Unsmoothing Commercial Property Returns: A Revision to Fisher-Geltner-Webb's Unsmoothing Methodology

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Abstract

Fisher, Geltner, and Webb (1993), in a highly influential paper, develop a procedure to recover the underlying market values from a smoothed valuation-based commercial property return index, without assuming that the underlying property market is informationally efficient. Many papers since then have used the Fisher-Geltner-Webb unsmoothing technique to desmooth commercial property returns. We show, however, that there is an inherent bias in Fisher-Geltner-Webb unsmoothing technique and propose a simple extension of their model to correct for this bias. We then compare the performance of our improved specification to that of the Fisher-Gelter-Webb model.

I. Introduction

One of the most challenging puzzles in real estate finance today is how to unsmooth commercial property returns. This unsmoothing problem arises because most commercial property return series are universally based on appraised market values rather than actual sales transactions, and because appraisers tend to be more "backward-looking," dependent on transaction price observations, than property market participants.

A number of approaches have been put forth to unsmooth commercial property returns. The most widely used approach (as far as we can tell) is the Fisher-Geltner-Webb (1993), hereafter FGW, unsmoothing technique.[1] The approach begins with the assumption that valuations will normally smooth the underlying property market values across time. We can think of the appraiser, in estimating the property value at each in point in time, as lacking perfect confidence in the most recent empirical evidence regarding property value. Under these circumstances, Quan and Quigley (1989) have shown that it is optimal in a formal sense for the appraiser to use a simple Bayesian updating rule, which amounts to an adaptive expectations or exponential smoothing approach to estimating property value. In this way, if the appraiser attaches too much weight to the most recent empirical evidence regarding property value, when this evidence is highly spurious, a large potential error in the valuation of the property will result. Likewise, if the appraiser does not give enough weight to the most recent empirical evidence regarding property value, then the average valuation is likely to lag the current true market value as of the time of the valuation. On average, however, when numerous individual property valuations made by numerous independent appraisers are averaged together (across properties, as of a given point in time), the random valuation errors should diversify away. Furthermore, over a long sample period of time, the average valuation is likely to be unconditionally biased.

To expand this valuation into a returns series, we can represent the "true" or underlying market value of the commercial property as the accumulation of the true returns over time. Then by the definition of the observable valuation-based return, we can express the observed return as a function of present and past true returns. This is a consequence of the exponential smoothing procedure assumed for the property-level valuation.

We can also express the observable valuation-based return by an autoregressive model involving an infinite moving average of the present and past smoothed returns, and presumably an error term. The advantage of this representation is that the equation can be inverted to obtain an expression for the unobservable underlying return as a function of the present and past values of the observable return.

[1] Other unsmoothing methodologies are described in Barkham and Geltner (1994), Blundell and Ward (1987), Brown and Matysiak (1998), Geltner (1993), and Ross and Zisler (1991).

The model is operationalized by applying standard univariate time-series estimation procedures to the appraisal-based returns. The parameter values from the autoregressive model are then used to unsmooth the return series. The perceived advantage of this approach is that it does not rely on the assumption of efficient markets.

However, the major disadvantage of this approach is that the error term in the regression model does not necessarily have an expectation zero. Using inflation-adjusted appreciation returns to estimate the model helps, but does not completely solve the problem. Even with inflation-adjusted appreciation returns, a non-zero error term is possible. And a large non-zero error term that varies over time can be quite troubling (as we shall see), since it would mean that a systematic bias is present.

This paper proposes a simple solution to this problem. The solution involves the use of generalized differences to alter the model into one in which the errors are independent. A constant term is then added to the model to control for omitted effects moving uniformly over time as well as potential spurious correlation of the expected return with time. In this context, the constant term removes any deterministic nonstationary component from the data, leaving only stochastic nonstationarity to be considered.

II. The FGW Model

Briefly, the FGW model takes the form

$$r_t^* = w_0 r_t + w(B) r_{t-1}, \tag{1}$$

where r_t^* is the smoothed return index during period t, r_t is the corresponding underlying true or unsmoothed return during period t, w_0 is a weight between 0 and 1, $w(B) = w_1 + w_2 B + w_3 B^2 + \cdots$, and B is a lag operator.[2]

Substituting and expressing r_t^* in terms of present and past values of r_{t-1}^*, we get

$$r_t^* = \phi(B) r_{t-1}^* + e_t \tag{2}$$

where $\phi(B) = \phi_1 + \phi_2 B + \phi_3 B^2 + \cdots$ is a lag operator polynomial and $e_t = w_0 r_t$.

It is convenient to write the autoregressive representation in (2) as:

$$r_t^* = (\phi_1 + \phi_4 B^3) r_{t-1}^* + e_t. \tag{3}$$

[2] The notation here follows as closely as possible the notation in Fisher, Geltner, and Webb (1993).

The first-order lagged value r^*_{t-1} is included in (3) to capture the tendency of appraisers to use a Bayesian updating rule. The fourth-order lagged value $B^3 r^*_{t-1}$ is included in (3) to deal with the fact that many properties are effectively reappraised only annually occurring in the fourth calender quarter.

The foremost requirement in (3) is that e_t be regarded as a random variable that has a normal distribution with mean 0 and a constant variance σ^2. It follows from this assumption that the conditional distribution of r^*_t given r^*_{t-1} will be a normal distribution with mean $(\phi_1 + \phi_4 B^3) r^*_{t-1}$ and variance σ^2.

Solving (3) for r_t we are able to obtain the unsmoothed value of r^*_t

$$r_t = \left(r^*_t - (\phi_1 + \phi_4 B^3) r^*_{t-1}\right) / w_0. \tag{4}$$

In turn, we can compute w_0 from the assumption that $\sigma(r_t) \equiv \sigma[r^*_t - (\phi_1 + \phi_4 B^3 r^*_{t-1})/w_0] \equiv \sigma^{SP}/2$, where $\sigma[r^*_t - (\phi_1 + \phi_4 B^3 r^*_{t-1})/w_0]$ is some number based on the observable historical r^*_t series and the empirical estimation of ϕ_1 and ϕ_4, and σ^{SP} represents the volatility of the S&P500 index of stock market values (in real terms). This condition amounts to assuming that the true volatility of commercial property values is approximately half the volatility of the S&P500 stock market index.

With this constraint on $\sigma[r^*_t - (\phi_1 + \phi_4 B^3 r^*_{t-1})/w_0]$, we obtain

$$w_0 = 2\sigma[r^*_t - (\phi_1 + \phi_4 B^3) r^*_{t-1}]/\sigma^{SP}. \tag{5}$$

FGW use this condition to obtain values of r_t. FGW also use inflation-adjusted appreciation returns to estimate (3), the thought being that $E(e_t) = 0$ is more theoretically sound for real returns. Additionally, FGW add a constant term ϕ_0 to (3) to make the mean of e_t zero. FGW's estimates are not fully efficient, however, because they ignore potential variations in r_t over time. And as we shall see, FGW's unsmoothing methodology will prove to be very sensitive to variations in r_t over time.

III. A Revision to the FGW Model

To correct for this problem, let us write an expression for r^*_{t-2} using (2). The result is as follows:

$$r^*_{t-2} = \phi(B) r^*_{t-3} + e_{t-2}. \tag{6}$$

Next, subtracting (6) from (2), we have

$$\Delta r_t^* = \phi(B)\Delta r_{t-1}^* + \Delta e_t \tag{7}$$

where $\Delta r_t^* \equiv r_t^* - r_{t-2}^*$ and $\Delta e_t \equiv e_t - e_{t-2} = w_0(r_t - r_{t-2})$. Assuming $r_t \approx \mu + r_{t-2} + \epsilon_t$ or $r_t - r_{t-2} = \mu + \epsilon_t$ where ϵ_t has a zero mean and variance σ_ϵ^2 and where ϵ_t is serially uncorrelated, then (7) becomes

$$\Delta r_t^* = \mu^{'} + \phi(B)\Delta r_{t-1}^* + \epsilon_t^{'} \tag{8}$$

where $\mu^{'} = w_0\mu$, $\epsilon_t^{'} = w_0\epsilon_t$, and, hence, $E(\epsilon_t^{'})$ will equal zero, as it should.

The advantages of estimating (8), as opposed to (3), are three-fold. First, the error term, $\epsilon_t^{'}$, now should have a mean zero and constant variance. Second, by using a generalized difference equation, instead of a first difference equation, we avoid the problem of having r_{t-1} on both the left and right-hand side of (8). Consequently, with truly exogenous variables on the right-hand side of (8), we can estimate $\phi(B)$ consistently and with lack of bias via ordinary least squares procedures. Third, by adding a constant term to (8), we are able to control for any omitted effects moving uniformly over time in the initial model (e.g., a reduction in long-term risks), something that FGW are unable to do. This fact permits a more efficient set of parameter estimates. Of course, with no omitted effects, the two models should be identical, and indeed should lead to similar estimates of the unsmoothed return.

The remainder of this paper reports results which verify the general proposition that, where there are possible nonstationarities or evidence of strong positive autocorrelation in the observed return series, the effects of such trends are to introduce a bias in the FGW model. To correct for this bias, (3) can be estimated using generalized differences. We illustrate this result by comparing and contrasting the results of (3) and (8) for three countries: the United States, the United Kingdom, and Japan. These are only three of a number of possible examples, but they deal with several important questions.

IV. Three Unsmoothing Cases Examined

A. The United States

We base our analysis on quarterly NCREIF (National Council of Real Estate Investment Fiduciaries) inflation-adjusted appreciation returns (adjusted by using the CPI index) in the period from 1978 to 1999. NCREIF returns represent an aggregate of individual property returns on income-producing properties owned by commingled funds on behalf of qualified pension and profit-sharing trusts or owned directly by these trusts and managed on a separate-account basis by life insurance companies or pension fund investment managers. The capital appreciation or depreciation component is estimated for every property, every quarter based on appraised market values (as of one or more quarters ago) rather than actual sales transactions.

NCREIF returns have several facets which appear inconsistent with the FGW model. First, in nominal terms, we observe that the mean (quarterly) capital appreciation return before 1992 is 0.44 percent in contrast to 0.12 percent after 1992 (see table 1). We also observe that the standard deviation of the capital appreciation return before 1992 is slightly larger than after 1992 (1.96 versus 1.22). Second, in real terms, the pre-1992 sample has a mean (quarterly) inflation-adjusted capital appreciation return of -1.00 percent in contrast to -0.49 percent for the post-1992 sample. Over the entire 1978-99 period, the mean (quarterly) inflation-adjusted capital appreciation return is -0.84 percent and the standard deviation is 1.65. Third, the inflation-adjusted capital appreciation returns are strongly positively correlated. The correlation is generally close to 0.55. One concern is that this strong positive correlation could lead to spurious results when (refeq:model) is estimated in levels.

The empirical analysis below will regress NCREIF inflation-adjusted appreciation returns on their lagged one and four quarter values. This specification is similar to that of FGW. Absent an external appraisal effect, we can test the proposition that $r_t^* = \phi(B)r_{t-1}^* + e_t$ by regressing the current NCREIF inflation-adjusted appreciation return on its lagged one value. To the extent that external appraisals have any effect on the calculation of the NCREIF index (and assuming, as FGW do, that most properties are effectively reappraised only annually occurring in the fourth calendar quarter), it should generally be a lagged fourth quarter one.

We obtain, with standard errors in parentheses, the coefficient estimates given in row 1 of table 2. The estimates are consistent with those reported by FGW. The NCREIF inflation-adjusted appreciation return is positively related to its lagged one and four quarter values, with their coefficients being 0.25 and 0.59. The regression coefficients are statistically significant at the 0.05 level, and the mean squared error of the equation is 0.0059. Solving (5) yields a value of $w_0 = 0.31$.

The next row contains coefficient estimates for the 1978-92 period. As can be seen, while the results remain qualitatively the same, the coefficients are changed somewhat. The r_{t-1}^* coefficient changes from positive and statistically significant to positive and statistically insignificant (a p-value of 0.08). The r_{t-4}^* coefficient increases from 0.59 to 0.72 (a 22 percent increase), but remains statistically significant. This yields an implied value of $w_0 = 0.25$.

We also ran a similar regression for the 1993-99 period. These estimates are presented in row 3 of table 2. The coefficient estimates have the predicted signs and, unlike the coefficient estimates in row 2, are 2 to $3\frac{1}{2}$ times their respective standard errors. Focusing first on the coefficient of r_{t-1}^*, the results show a lagged one effect of 0.32, up from 0.25

in row 1 and 0.20 in row 2. In contrast, the coefficient of r_{t-4}^* is 0.42, down from 0.59 in row 1 and 0.72 in row 2. Also note that b_0 changes from negative and statistically insignificant in rows 1 and 2 to positive and statistically insignificant. These effects are generally consistent with a negative mean e_t term, which biases upward the coefficients b_0 and b_1.

The implications of these findings are shown in figure 1, which graphs the smoothed and unsmoothed NCREIF capital appreciation indices. The unsmoothed NCREIF capital appreciation return is calculated according to $r_t = (r_t^* - b_1 r_{t-1}^* - b_4 r_{t-4}^*)/w_0$. Both series are expressed in nominal terms by adding back in the inflation rate. The solid line plots the smoothed NCREIF capital appreciation return index. The dotted line plots the unsmoothed index.

It is apparent from figure 1 that the two indices generally move together over the 1978-92 period. However, this strong association stops in 1992. In the post-1992 period, the smoothed index is generally decreasing, while the unsmoothed index wanders off. This result is consistent with a nonstationary e_t term.

As a means of salvaging the FGW unsmoothing methodology, we re-estimated their model using generalized differences. The estimated coefficients and standard errors are shown in table 3. The left-hand side variable is $r_t^* - r_{t-2}^*$. The right-hand side variables are $r_{t-1}^* - r_{t-3}^*$ and $r_{t-4}^* - r_{t-6}^*$. Our results suggest that by estimating (3) using generalized differences we obtain vastly different results. For 1978-99, we find a negative and statistically insignificant b_1 coefficient. Our b_4 coefficient estimate is 0.52 and statistically significant. Despite these vastly different results, the resulting w_0 is 0.33.

We also estimated (8) for the 1978-92 and 1993-99 time periods. These estimates are summarized in rows 2 and 3 of table 3. The results generally lead to the same qualitative conclusions as the regression in row 1. The b_4 coefficient, for example, is the only coefficient which is statistically significant in both cases. Further, there is no significant effect from $r_{t-1}^* - r_{t-3}^*$. Our b_1 coefficient remains statistically insignificant in both cases.

In figure 2, we explore the implications of these results for r_t. Here we graph the smoothed NCREIF capital appreciation index and our corresponding unsmoothed index. The solid line plots the unsmoothed NCREIF capital appreciation index. The dotted line plots our unsmoothed index. As above, the unsmoothed capital appreciation return is calculated according to $r_t = (r_t^* - b_1 r_{t-1}^* - b_4 r_{t-4}^*)/w_0$. The parameter estimates, however, are those from table 3 rather than table 2.

As figure 2 shows, our unsmoothed index now lies much closer to the smoothed index than in figure 1. Our unsmoothed index starts off by increasing from years 1978 through 1990, and then declines from years 1991 through 1997, ending the entire sample period 10 percent lower than in 1978. As we noted in the introduction, the key to understanding why our unsmoothed index does not wander off after 1992 as does the FGW index is in the generalized-differences transformation. Our generalized-difference transformation places much less weight on a lag one effect, reflecting the fact that $r^*_{t-1} - r^*_{t-3}$ varies much more rapidly over time than r_{t-1}.

B. The United Kingdom

The IPD (Investment Property Databank) index, which is described in detail in Brown and Matysiak (1998) and elsewhere, is produced directly from survey data collected from large professional investors (including institutions, property companies, and open-ended investment funds). The data consist of a time series of income and capital growth (based on appraised values) over the period from December 1986 to December 2000.

The primary feature of the IPD index is that it is constructed monthly. In addition, the IPD collects considerably greater depth of property appraisal and operational-level information (including appraisal assumptions and inputs, lease, rental and occupancy information, and detailed breakdowns of operating expenses) than, for example, NCREIF.

The IPD index is, according to our understanding, the most widely used commercial property index in the UK. Other available commercial property indices include the JLL (Jones Lang LaSalle) index and the REI (Richard Ellis) index.

Table 4 presents descriptive statistics on the IPD index. The IPD index seems stable through time. For example, in the months from 1987 through 1993, the mean (nominal) capital appreciation return is 0.22 percent. Similarly, for the months from 1993 through 2000, the mean (nominal) capital appreciation return is 0.26 percent. However, in real terms (deflated by the retail price index), the mean monthly capital appreciation returns are -0.30 and 0.03 percent, respectively. Further, the positive autocorrelation in the monthly capital appreciation inflation-adjusted returns is 0.77.

We have estimated a regression of IPD inflation-adjusted appreciation returns on their lagged one and twelve values. Our results appear in table 5. We estimated the regressions in table 5 over three periods: 1987-2000, 1987-93, and 1994-2000. The fits obtained here are similar to those obtained with the quarterly NCREIF data. However, the optimal weight that one should attach to the current true market value as of the time of the valuation (i.e., the optimal value of w_0) is considerably different. Those range from -0.06 (which is outside the feasible range) to 0.18.

For the whole period (see row 1 of table 5), we find a strong lag one effect. The coefficient on r_{t-1} is 0.74 with a t-statistic of 14.3 (0.74 divided by 0.05). There also is a positive and significant lag twelve effect. The coefficient on r_{t-12} is 0.11 with a t-statistic of 2.09 (0.11 divided by 0.05). Both of these results seem plausible.

In the regression from 1987 through 1993 (see row 2 of table 5), the coefficient of r_{t-1} is 0.81 and remains statistically significant. The coefficient of r_{t-12}, on the other hand, is cut in half and the effect is statistically insignificant (a p-value of 0.33). The problem seems to be one of power.

The final set of results in table 5, reported in row 3, pertain to the subperiod 1994-2000. First we note that the coefficient of r_{t-1} in this case is 0.56 and statistically significant. Next we note that the coefficient of r_{t-12} is positive, but insignificant (p-value of 0.29). These results make it impossible to interpret the value of w_0. With $b_1 = 0.56$ and $b_{12} = 0.08$, the implied value of $w_0 = -0.06$. Note that a value of $b_{12} = 0$ (i.e., no lag twelve effect) yields a more reasonable measure of $w_0 = 0.12$.

Next, figure 3 presents the smoothed and unsmoothed IPD capital appreciation index from 1987 through 2000. The unsmoothed index in this case depends sensitively on the value of w_0. With a low value of w_0, differences between r_t and r_{t-1} are greatly magnified. Also, as we shall see below, the unsmoothed index depends sensitively on the lag effects. Given a low value for b_{12}, for example, the unsmoothed index in figure 3 turns down quickly in August 1989. It then turns up quickly in May 1993, only to turn down quickly again in March 1994. The problem lies in these turn points. Appraisal smoothing should dampen the cyclical influences on the capital appreciation return. Unsmoothing should amplify them, as opposed to the other way around.

To explore this issue further, we re-estimated (3) using generalized differences. The results of this generalized-difference transformation are presented in table 6. The model removes the independent trend from the data before estimating the value of b_1 and b_{12}. Three equations are estimated: one for the whole period 1987-2000, one for the subperiod 1987-93, and one for 1994-2000. The left-hand side variable is $r^*_t - r^*_{t-2}$. The right-hand side variables are $r^*_{t-1} - r^*_{t-3}$ and $r^*_{t-12} - r^*_{t-14}$.

The first set of results in table 6, reported in row 1, pertain to the whole period 1987-2000. Here the variable $r^*_{t-1} - r^*_{t-3}$ has a strong positive and highly significant effect on $r^*_t - r^*_{t-2}$. In contrast, the variable $r^*_{t-12} - r^*_{t-14}$ has a rather puzzling, but very significant negative effect. This negative effect suggests that commercial property prices (at least in the United Kingdom) come back slowly and partially following a shock, so that a high r^*_{t-12} implies a low r^*_t today.

Row 2 of table 6 reports results pertaining to the subperiod 1987-93. These results bear on the issue of stability through time. Here there is evidence that a high r^*_{t-12} implies a high r^*_t today. These results are quite different from those reported in row 1. The coefficient of $r^*_{t-12} - r^*_{t-14}$ is now 0.45 with a t-statistic of 3.9 (0.45 divided by 0.12). We also find a positive and significant coefficient on $r^*_{t-1} - r^*_{t-3}$.

In row 3 of table 6, we report the results of estimating (8) for 1994-2000. An interesting finding is that the coefficient of $r^*_{t-1} - r^*_{t-3}$ is now positive, but insignificant. However, the coefficient $r^*_{t-12} - r^*_{t-14}$ continues to be positive and significant.

Figure 4 shows that the profile of our unsmoothed IPD capital appreciation index tracks the smoothed index. Furthermore, there is a wider variation in our unsmoothed index than in the FGW index, particularly in the months from May 1989 through August 1993. Thereafter, the two series – ours and FGW's – are quite similar. We interpret these results as support for our generalized-difference estimation.

C. Japan

The time series on inflation-adjusted appreciation returns we use is the Mitsubishi Trust and Banking Corporation-IKOMA annual real estate investment index. We choose this series because it is available for a long sample. For the period for which it overlaps with existing other series, such as those based on official commercial land price data, the different series move closely together. In particular, all series show booming real estate and commercial land prices during the 1980s. Then between 1990 and 1996, all series show falling real estate and commercial land prices.

The MTB-IKOMA index calculates total returns (income and appreciation) for 73 zones in 13 cities around Japan, including such areas as Marunouchi in Tokyo, Umeda in Osaka, and Sakae in Nagoya, using yields and appraised values. Specifically, total return is the sum of the estimated yield and the change in appraised market values. Effective rents are constructed from a multiple linear regression equation, given information on contract rents for approximately 12,000 properties. The multiple linear regression controls for four categories of buildings based on current conditions in the Japanese office market: total floor space of up to less than 1,815 square meters, 1,815 up to 3,300 square meters, 3,300 to 9,900 square meters, and 9,900 or more square meters. MTB-IKOMA also provides an aggregate index, and it is this index (deflated by the CPI index) that we use in our analysis below. End of year appraisal values are the result of adding the market value of the land to the replacement cost (net of depreciation) of a hypothetical office building. In this measure, land is valued at a price designated by the government – thus, land values correspond to the so-called "Official Land Price," or Chika-koji.

Table 7 shows the time-series statistics of the nominal annual capital component of the returns to the unsmoothed MTB-IKOMA index. For comparison, annual capital appreciation returns are reported for two subperiods: 1970-85 and 1986-1996. In nominal terms, the average annual capital appreciation return is 10.6 percent in the years from 1970 through 1985 and 1.13 percent thereafter. In real terms (deflated using the CPI index), the average annual capital appreciation returns are 2.37 and -0.72 percent, respectively. The first-order autocorrelation of the capital appreciation inflation-adjusted returns is 0.69, indicating some sluggishness or inertia in the return series.

Here we run regressions analogous to those above: we regress MTB-IKOMA inflation-adjusted annual appreciation returns on their lagged one values. The results in table 8 show a strong and highly significant relation between r_t^* and r_{t-1}^* for the whole sample and in the subperiod 1986-96. In the subperiod 1970-85 in row 2, the coefficient of r_{t-1}^* is positive, but insignificant.

Figure 5 shows the nominal property value levels implied by the smoothed MTB-IKOMA index and the FGW unsmoothed index. The FGW unsmoothed index is obtained by solving $r_t = (r_t^* - b_1 r_{t-1}^*)/w_0$, using a value of $w_0 = 0.12$.[3] As figure 5 shows, it is hard to take the FGW unsmoothed index seriously. In the years from 1970 through 1980, the FGW unsmoothed index increases by over 15,000 percent. Thereafter, the FGW unsmoothed index decreases by 96 percent, to just about 440 (1970 = 100) by year-end 1996. In comparison, the smoothed MTB-IKOMA index rises to about 1,000 in 1990, before decreasing to about 440 in 1996.

We also investigated regressions of $r_t^* - r_{t-2}^*$ on $r_{t-1}^* - r_{t-3}^*$. The results of these regressions are shown in table 9. For comparison, we present results for the whole sample in row 1, for the subperiod 1970-85 in row 2, and for 1986-96 in row 3. In row 1, our estimate of b_1 is 0.31, which is relatively low compared to FGW's estimate, but the t-statistic is 1.52. This regression has only 23 observations, however. The implied value of w_0 is 0.61, compared with FGW's estimate of $w_0 = 0.12$.

In rows 2 and 3, our estimate of b_1 varies from 0.13 to 0.50. The coefficient is insignificant in both regressions, however. A better fit is obtained in the regression for 1986-96, but the implied value of w_0 is 1.38, with a 95 percent confidence interval between 0.94 and 1.81.

Turning now to the results in figure 6, we see that our generalized-difference index follows the smoothed MTB-IKOMA index much better than the FGW index. In the years from 1970 through 1990, our index increases by about 1100 percent (equivalent to an annual growth rate of 13 percent). Thereafter, our index declines by 65 percent, or 16 percent per year.

V. Conclusion

Our two related goals in this paper were the following: Firstly and mainly, we wanted to point out a potential source of bias in the FGW unsmoothing model. Secondly, we wanted to compare and contrast the unsmoothing approach originated by FGW with the extension suggested herein. To do this, we conducted an analysis much like FGW's, using valuation-based commercial property returns for the United States, the United Kingdom, and Japan. We then presented and discussed the results obtained for each of these countries.

[3] Values of w_0 range from 0.12 for the whole sample to 1.63 (which is outside the feasible range) for the 1986-96 period.

Our results can be summarized as follows:

1. We do find that valuation-based commercial property returns can be better modelled using a generalized-difference specification. This improved specification leads to real and economically meaningful changes in all three countries.
2. In the United States, the improved specification behaves more reasonably by not wandering off (as the FGW model does) in the years after 1992.
3. In the United Kingdom, the improved specification behaves more reasonably by amplifying, as opposed to dampening, the cyclical influences on the capital appreciation return.
4. In Japan, the improved specification tracks the MTB-IKOMA smoothed return index more closely. In contrast, the FGW model produces spurious results.

These findings bear on two concerns that were raised in connection with FGW's unsmoothing model. One is that the error term in their model does not necessarily have an expectation zero. This can easily lead to biased coefficient estimates depending on the choice of sample period, which helps to explain why, for instance, the FGW unsmoothed index in the United States wanders off in the years after 1992, while the generalized-difference specification does not.

A second concern is that, because valuation-based commercial property returns vary smoothly over time, a generalized-difference specification is needed to transform the resulting error terms in FGW's unsmoothing model into a stationary series. In so doing, however, there is a potential danger that negative autocorrelation will be introduced. To account for this problem, we added a constant term to our generalized-difference transformation. Our modelling approach was seen to provide much more plausible results than FGW's model. This is particularly evident in the case of Japan.

One drawback of our generalized-difference approach is that it does not capture the possibility of nonstationary variances, nor does it capture the potential instability of coefficients in a given period.

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Table 1.
Summary Statistics: NCREIF Index*
(quarterly)

	Income Yield	Capital Appreciation	Total Return
1978-1999	1.93%	0.34%	2.27%
	(0.20)	(1.77)	(1.84)
1978-1992	1.85%	0.44%	2.29%
	(0.17)	(1.96)	(2.05)
1993-1999	2.13%	0.12%	2.24%
	(0.08)	(1.22)	(1.26)

* Standard deviations are in parentheses (in percent).

Table 2.
Regression Results: FGW Methodology*
Dependent variable: NCREIF inflation-adjusted appreciation return

$$\mathbf{r_t^* = b_0 + b_1 r_{t-1}^* + b_4 r_{t-4}^* + \varepsilon_t}$$

	b_0	b_1	b_4	R^2	MSE	F-value	w_0
1978 - 1999	-0.0010	0.2532	0.5887	0.54	0.0058	42.7	0.3133
	(0.0015)	(0.09)	(0.09)				
1978 - 1992	-0.0029	0.2007	0.7208	0.55	0.0047	30.5	0.2456
	(0.0020)	(0.11)	(0.12)				
1993 - 1999	0.0026	0.3210	0.4167	0.71	0.0015	27.2	0.2360
	(0.0018)	(0.12)	(0.12)				

*Standard errors are in parentheses.

Table 3.
Regression Results: Generalized-Difference Specification*
Dependent variable: NCREIF inflation-adjusted appreciation return

$$r_t^* - r_{t-2}^* = b_0 + b_1(r_{t-1}^* - r_{t-3}^*) + b_4(r_{t-4}^* - r_{t-6}^*) + \varepsilon_t$$

	b_0	b_1	b_4	R^2	MSE	F-value	w_0
1978 - 1999	0.0002	-0.0079	0.5224	0.31	0.0026	15.5	0.3315
	(0.001)	(0.10)	(0.09)				
1978 - 1992	-0.0001	-0.0565	0.5836	0.33	0.0025	11.4	0.3616
	(0.002)	(0.12)	(0.12)				
1993 - 1999	0.0015	0.2149	0.2930	0.20	0.0002	2.7	0.2616
	(0.002)	(0.18)	(0.14)				

*Standard errors are in parentheses.

Table 4.
Summary Statistics: IPD Index*
(monthly)

	Income Yield	Capital Appreciation	Total Return
1987-2000	0.32% (0.74)	0.24% (0.88)	0.85% (0.85)
1987-1993	0.39% (1.01)	0.22% (1.14)	0.80% (1.10)
1994-2000	0.25% (0.27)	0.26% (0.50)	0.90% (0.50)

* Standard deviations are in parentheses (in percent).

Table 5.
Regression Results: FGW Methodology
Dependent variable: IPD inflation-adjusted capital appreciation return

$$r_t^* = b_0 + b_1 r_{t-1}^* + b_{12} r_{t-12}^* + \varepsilon_t$$

	b_0	b_1	b_{12}	R^2	MSE	F-value	w_0
1987-2000	-0.0003	0.7448	0.1059	0.61	0.0046	118.3	0.1870
	(0.0005)	(0.05)	(0.05)				
1987-1993	-0.0002	0.8129	0.0762	0.67	0.0038	67.7	0.0587
	(0.0009)	(0.08)	(0.08)				
1994-2000	0.0001	0.5602	0.0780	0.36	0.0005	23.5	-0.0571
	(0.0005)	(0.08)	(0.07)				

*Standard errors are in parentheses.

Table 6.
Regression Results: Generalized-Difference Specification
Dependent variable: IPD inflation-adjusted capital appreciation return

$$r_t^* - r_{t-2}^* = b_0 + b_1(r_{t-1}^* - r_{t-3}^*) + b_{12}(r_{t-12}^* - r_{t-14}^*) + \varepsilon_t$$

	b_0	b_1	b_{12}	R^2	MSE	F-value	w_0
1978 - 1999	-0.0003	0.6426	-0.8358	0.42	0.0021	54.7	0.8656
	(0.0005)	(0.08)	(0.08)				
1987-1993	0.0004	0.2155	0.4462	0.26	0.0009	11.5	0.4053
	(0.001)	(0.11)	(0.12)				
1994-2000	-0.0005	0.1325	0.3079	0.15	0.0003	7.1	0.1005
	(0.0006)	(0.10)	(0.09)				

*Standard errors are in parentheses.

Table 7.
Summary Statistics: MTB-IKOMA Index*
(annual)

	Income Yield	Capital Appreciation	Total Return
1970 - 1996	3.36%	3.49%	6.83%
	(0.72)	(1.49)	(14.89)
1970 - 1985	3.82%	6.76%	10.58%
	(0.34)	(9.90)	(9.76)
1986 - 1996	2.65%	-1.28%	1.37%
	(0.63)	(19.64)	(19.45)

* Standard deviations are in parentheses (in percent).

Table 8.
Regression Results: FGW Methodology
Dependent variable: MTB-IKOMA inflation-adjusted capital appreciation return

$$r_t^* = b_0 + b_1 r_{t-1}^* + \varepsilon_t$$

	b_0	b_1	R^2	MSE	F-value	w_0
1971 - 1996	0.0018	0.7019	0.48	0.2457	21.5	0.1223
	(0.02)	(0.15)				
1971 - 1985	0.0294	0.3153	0.08	0.0129	1.1	0.7971
	(0.30)	(0.30)				
1986 - 1996	-0.0231	0.8299	0.75	0.2530	26.9	1.6277
	(0.29)	(0.16)				

*Standard errors are in parentheses.

Table 9.
Regression Results: Generalized-Difference Specification
Dependent variable: MTB-IKOMA inflation-adjusted capital appreciation return

$$r_t^* - r_{t-2}^* = b_0 + b_1(r_{t-1}^* - r_{t-3}^*) + \varepsilon_t$$

	b_0	b_1	R^2	MSE	F-value	w_0
1971 - 1996	-0.0177	0.3127	0.10	0.0546	2.3	0.6090
	(0.32)	(0.20)				
1971 - 1985	-0.0032	0.1289	0.02	0.0049	0.2	0.9170
	(0.05)	(0.31)				
1986 - 1996	-0.0253	0.4958	0.25	0.0619	3.0	1.3750
	(0.05)	(0.29)				

*Standard errors are in parentheses.

Figure 1. Original and Unsmoothed FGW Index:
NCREIF Capital Values (19793 = 100)

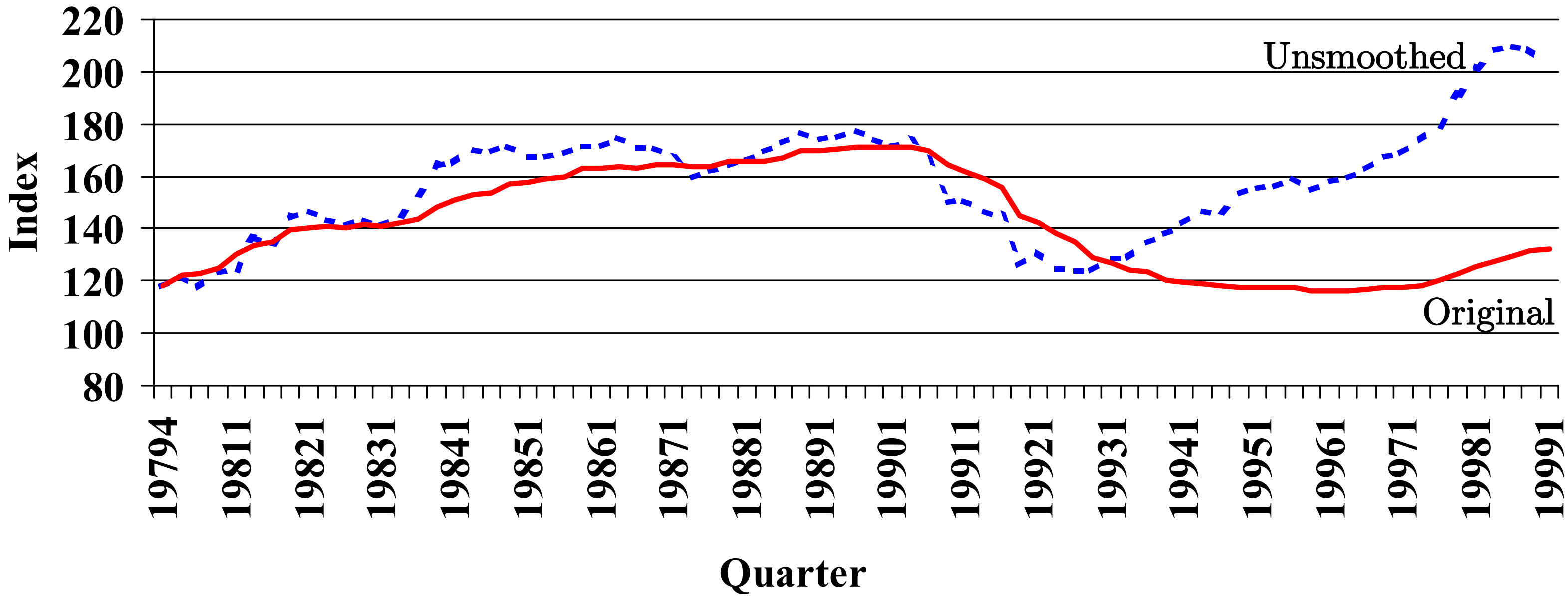

Figure 2. Original and Unsmoothed Generalized-
Difference Index:
NCREIF Capital Values (19793 = 100)

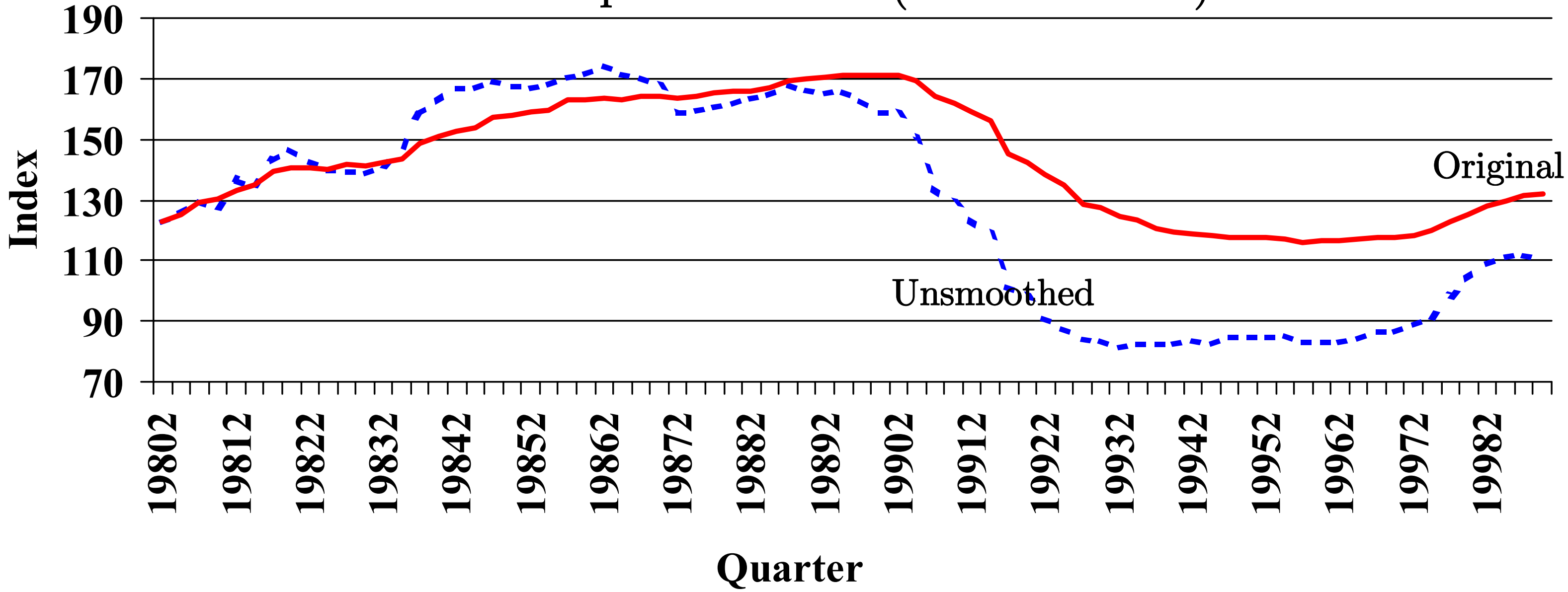

Figure 3. Original and Unsmoothed FGW Index: IPD Capital Values (December 1986 = 100)

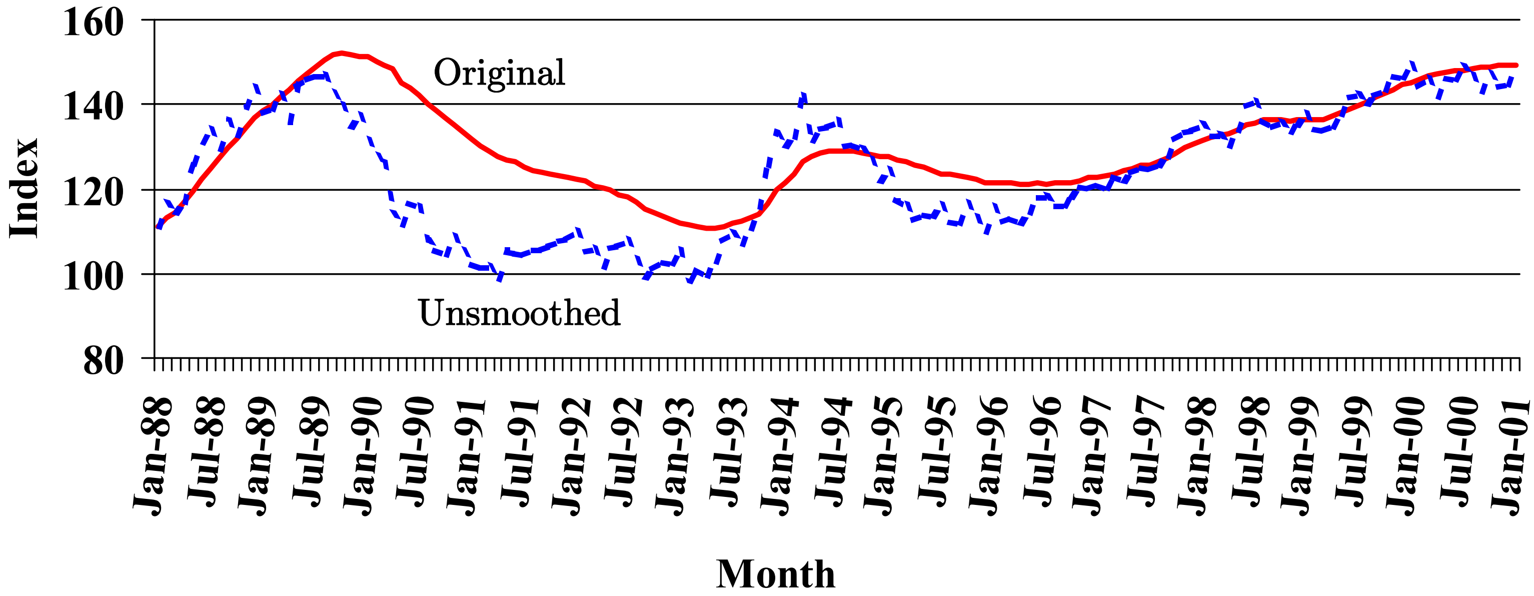

Figure 4. Original and Unsmoothed Generalized-Difference Index:
IPD Capital Values (December 1986 = 100)

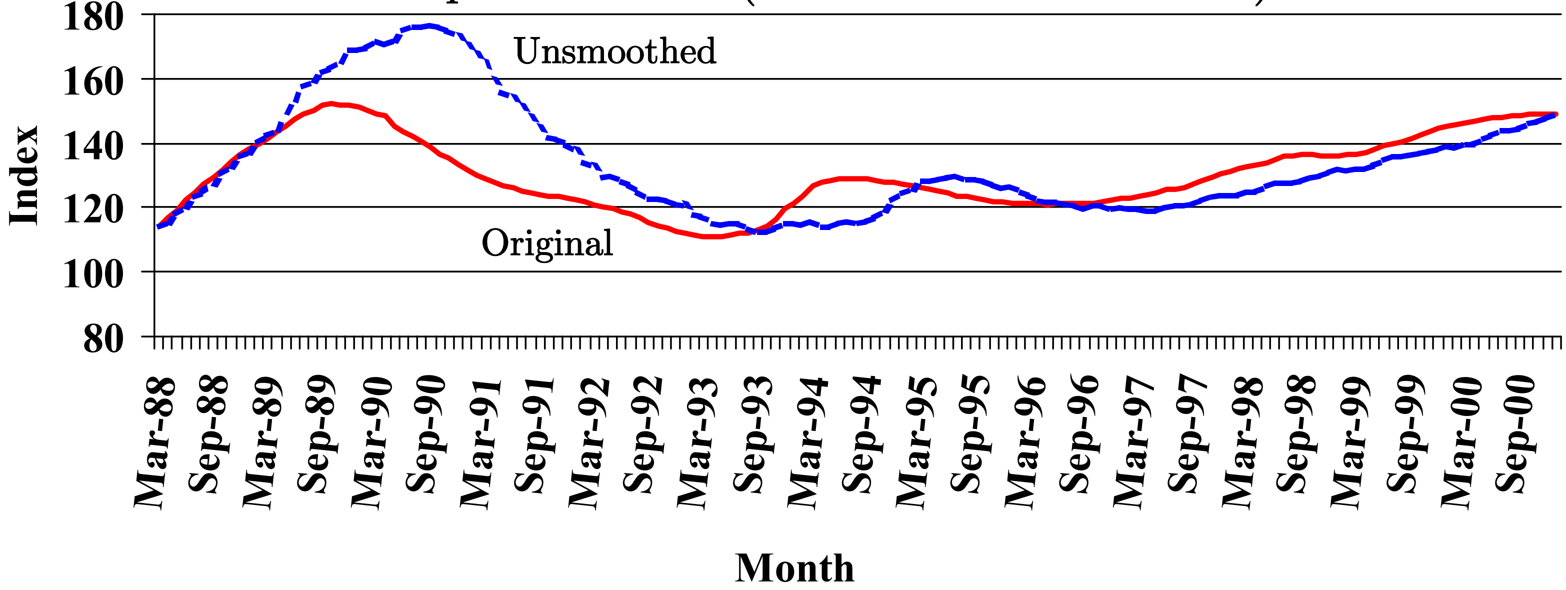

Figure 5. Original and Unsmoothed FGW Index:
MTB-IKOMA Capital Values (1970 = 100)

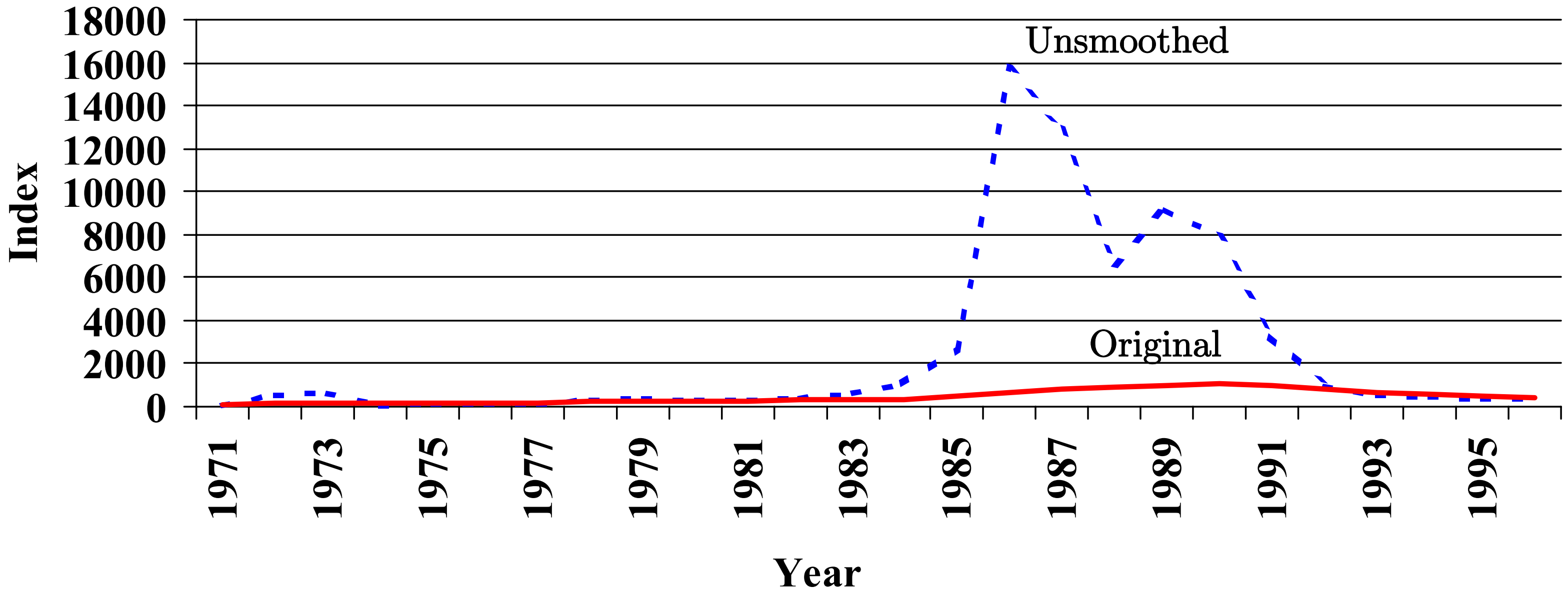

Figure 6. Original and Unsmoothed Generalized-Difference Index: MTB-IKOMA Capital Values (1970 = 100)