Significant Alphas in Real Estate Funds: Do Fund Managers Add Value?

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INTRODUCTION

Since Jensen's 1967 study, the majority of academic research has supported the efficient market hypothesis, which implies that fund managers cannot routinely over time 'beat the market' with superior risk adjusted returns (Fama and French (2010)). However, sector fund managers may have access to informational asymmetries about their area that enable superior risk adjusted returns (Dellva, DeMaskey and Smith (2001)). In this environment, real estate fund management has presented a conundrum, with several studies reporting positive and significant alphas[1] for real estate funds (Cici, Corgel and Gibson (2011), Kallberg, Liu, Trzcinka (2000), Gallo, Lockwood and Rutherford (2000)). Other researchers report real estate fund managers do not outperform the market on a risk-adjusted basis, taking into account differing benchmarks, bootstrap methodology, control factors or test periods (Lin and Yung (2004), Rodriquez (2007), and Chiang, Kozhevnikov, Lee and Wisen (2008), Kaushik and Pennathur (2013)).

We evaluate two issues regarding Jensen's alpha as a measure of superior risk adjusted returns to real estate mutual funds. First, are the test statistics sensitive to alternative estimates of the standard errors that take heteroskedasticity into account?[2] Second, are the estimates of Jensen's alpha sensitive to specification error in the

[1] Jensen's alpha is used to evaluate a fund manager's ability to provide superior risk-adjusted returns. The Morningstar Box Score Report, http://global.morningstar.com/US/documents/Indexes/MorningstarBoxScoreReport2H09.pdf (accessed October 2013

[2] Real estate returns have been noted for heteroskedasticity (Young (2008), Cheng (2005), Yang and Chen (2009)).

model, i.e., is the apparent significance of alpha attributable to the pricing of factors other than “the” market return?

In in order to address the first issue, we evaluate the significance of Jensen’s alpha in real estate returns using several robust (i.e., heteroskedastic-consistent) estimators including the variants of the White (1980) estimator present by MacKinnon and White (1985), the Newey-West (1987) heteroskedastic-autocorrelation-consistent [HAC] estimator, and the ‘wild’ bootstrap suggested by Davidson and Flachaire (2008).[3]

In order to address the second issue, we estimate regression models that include innovations in the default spread, credit spread, market skewness and change industrial production growth as explanatory variables.

Additionally, three indexes as benchmarks are used in evaluating significant alphas in real estate returns. The Wilshire 5000 Index provides a total market benchmark, the Wilshire Real Estate Index provides a general real estate benchmark, and the NAREIT Index provides a REIT benchmark.

We compare the frequency of significant alphas obtained from OLS regression with OLS standard errors to the results obtained from three heteroskedastic-consistent [HC] estimators, the Newey-West (1987) standard errors, and a wild bootstrap applied to one of the HC estimators[4]. Examining several contiguous sub periods over the interval from 1990 to 2012, we find that the heteroskedastic-consistent standard errors reduced the number of significant alphas exhibited by real estate mutual funds.

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[3] We are not aware of any other study of Jensen’s alpha that relies on a wild bootstrap.

[4] More specifically, we apply the wild bootstrap to the HC standard error designated HC3 by MacKinnon and White (1985); HC3 is an exact equivalent to the ‘delete-one’ jackknife. Since HC3 is more conservative (less likely to reject the null) than other HC standard errors, our goal is to establish whether bootstrapping can further reduce the frequency of significant alpha.

Implementing a wild bootstrap consistently provides the most conservative result. Similar results were obtained with real estate investment trusts. [REITs]

Contrary to expectations, the HAC (Newey-West) standard error increased the per cent of REITs that exhibit significant alphas. For most of the time periods studied (regardless of benchmark) the number of REITs that exhibit significant alphas based on Newey-West standard errors was greater than any other estimator.

In general, adding explanatory variables to the regression failed to systematically attenuate the frequency of significant alphas in real estate fund returns. However, using the nominal 5% critical values, the inclusion of the short-run volatility proxy reduced the per cent of real estate mutual funds that exhibit significant alphas from 14% to 8% over the 2005 to 2012 period. A further reduction occurred with the wild bootstrap. This result suggests that financial constraints (proxied by market skewness) may be a factor in the pricing of real estate mutual funds returns during some time periods.

For the REIT sample, short- and long-run volatility proxies attenuated the frequency of significant alphas from 9% to 7% over the 1996 to 2012 period using OLS standard errors. When the REIT sample is examined with the wild bootstrap applied to heteroskedastic consistent standard errors, the proportion of REITs that exhibit significant alphas is reduced to less than 4% in most periods, i.e. less than the nominal significance level of the test.

The addition of innovations from the change in default and term spread had a minimal effect on explanatory power as measured by the adjusted R^2. Our proxies for short and long-run volatility increased the explanatory power only in the 1990 to 2012 sample for real estate mutual funds.

Our results suggest the potential for erroneous interpretation of significance of variables when the standard errors are not adjusted for heteroskedasticity. Tabulating critical values of the empirical distribution using a wild bootstrap provided the most conservative result regarding the frequency of real estate funds with significant alphas.

The remainder is organized as follows: Section 2 describes the methodology, Section 3 the data, Section 4 and 5 provides empirical results and concluding comments.

Section 2 Methodology

Jensen's alpha measures fund performance using the single index version of the capital-asset pricing model (CAPM). The model assumes that realized returns on the security or portfolio are expressed as a linear function of its systematic risk, the realized returns on the market portfolio, the risk-free rate and a random error. Following Gallo et. al. (2002), our first model specification is based on the CAPM:

$$R_{i,t} = \alpha_i + \beta_i R_{m,t} + \varepsilon_{i,t} \quad (1)$$

where R_{it} is the excess return on the real estate mutual fund or REIT *i* in month t net of the change in the monthly return on the Citi 3-month Treasury Bill[5] in month t. Alternative proxies for $R_{m,t}$ are the excess return on the Wilshire Real Estate Index, the NAREIT Index, or the Wilshire 5000 Index.

[5] Citi 3-month Treasury Bill was used by Gallo, Lockwood and Rutherford (2000) in their study.

Following Hahn and Lee (2006) our second model includes innovations from the default spread and the term spread:

$$R_{i,t} = \alpha_i + \beta_i R_{m,t} + \delta_i \Delta default_t + \lambda_i \Delta term_t + \varepsilon_{i,t} \qquad (2)$$

Where $R_{i,t}$, α_i, and $R_{m,t}$ are as previously defined and the $\Delta default_t$ is the white noise innovations from $def_t - def_{t-1}$ and $\Delta term_t$ is the white noise innovation from $term_t - term_{t-1}$, and def_t and $term_t$ are the default spread and term spread at time t.

Following Adrian and Rosenberg (2008), our third model is an attempt to improve explanatory power by including market skewness and the change in growth in industrial production as proxies for short-and long-term volatility:

$$R_{i,t} = \alpha_i + \beta_i R_{m,t} + \varphi_i \text{market skewness}_t + \omega_i \Delta \text{industrial production}_t + \varepsilon_{i,t} \qquad (3)$$

Where $R_{i,t}$, α_i, and $R_{m,t}$ are as previously defined and the market skewness variable and the change in industrial production are included in the model. Adrian and Rosenberg (2008) found that adding short- and long-run volatility components achieved lower pricing errors, as measured by the root-mean-squared pricing error, than the market model with size and book-to-value factors. The third specification of the model includes innovations from market skewness and the change in industrial production.[6]

To evaluate whether fund managers can add value (as measured by Jensen's alpha, i.e.,$\hat{\alpha}$ in equations 1-3) we evaluate the standard error of the estimate with several robust (i.e., heteroskedastic-consistent) estimators. Following MacKinnon and

[6] To explain the short-run volatility factor, Adrian and Rosenberg (2008) posit that 'negative shocks to the market return increase short-run volatility more than positive shocks to the market'. This relationship is reflected in the 'strong positive correlation in the market risk premia of short-run volatility and skewness'

White (1985) we tabulate test statistics and empirical rejection rates for the null $\alpha = 0$ from the following variants of the White (1980) standard error:

HC1 is the White (1980) estimator adjusted for degrees of freedom.

HC2 is the leverage-adjusted heteroskedastic-consistent estimator.

HC3 is an exact equivalence to the delete-one jackknife.

Following Davidson and Flachaire (2008), we also evaluate HC3 with a ‘wild’ bootstrap.[7] The empirical distribution of the test statistic is tabulated from 10,000 iterations of the bootstrap; the critical value for the test is determined by the 5 % tail of the empirical (bootstrapped) distribution after sorting the test statistic from high to low.

Section 3 DATA

Real Estate Mutual Funds

Domestic fund data available on Morningstar[8] from January 1990 to December 2012 comprises the real estate mutual fund sample in this study. Funds with a ‘successful return history tend to issue multiple-share classes’ (Chiang, Kozhevnikov, Lee and Wisen (2008)). By excluding all but a single fund from the multiple-share classes, ‘positive performance bias’ is avoided (Chiang, Kozhevnikov, Lee and Wisen (2008)). Therefore, funds with multiple share classes are limited to a one-time inclusion in the study. This criterion eliminated 269 funds. A minimum of 36 consecutive months of data is also required for inclusion in the sample which eliminated an additional 15 funds. The final sample contains 80 funds, with five of the 80 funds active from 1990 to 2012.

[7] In accordance with Davidson and Flachaire (2008) we select the Radermacher distribution for the auxiliary random variable.

[8] Morningstar Inc, provided monthly and daily return data for real estate mutual funds.

The number of real estate mutual funds from the Morningstar dataset increased from five funds in 1990 to a peak of seventy-two funds in 2007. The summary statistics for each period under study are listed in Table 1. The first period, 1991 to 1997, corresponds to the time period studied by Gallo, Lockwood and Rutherford (2000). Subsequent seven-year periods, an eight-year period and the full period of 1990 to 2012 follow the initial seven-year analysis.

There are 84 months in the first and second study periods, with an average number of 53 and 76 observations. The third period included eight years or 96 months with an average of 85 data points per fund. The entire study included 12,000 observations from 1990 to 2012.

Table 1 Real Estate Mutual Fund Summary Statistics

Real Estate Mutual Fund Summary Statistics[9]

	1991 to 1997	**1998 to 2004**	**2005 to 2012**	**1990 to 2012**
Number of firms	20	52	76	80
Average Number of Observations	53	76	85	150
Mean Excess Return (%)	1.277	0.867	0.585	0.779
Median Excess Return (%)	1.097	1.271	1.467	1.497
Standard Deviation	3.385	4.094	7.796	6.498
Skewness	0.605	-0.720	-0.644	-0.731
Kurtosis	3.697	4.593	7.035	8.574

The null hypothesis of no autoregressive conditional heteroskedasticity was rejected for 37 of the 80 real estate mutual funds on the basis of Engle's ARCH test at the 5% level. The Breusch-Godfrey test for serial correlation was rejected for 15 funds.[10] Forty-one funds exhibited serial correlation or heteroskedasticity.

[10] The significance level for the Breusch-Godfrey test was 5%.

A two-step process developed by Lee, Strazicich, and Meng (2012) is used to test for structural breaks in the mutual fund return series.[11] The majority of the breaks in the real estate funds were during the financial crisis period of 2008 and 2009.

REITs

Monthly returns of REITs was collected from the Center for Research in Security Prices (CRSP) database from 1990 to 2012 specifying security characteristic line (SCL) 18, which identified 363 REITs. The REIT sample includes equity, mortgage and hybrid REITS listed on the NYSE, AMEX and NASDAQ.[12] The initial sample is filtered further by requiring 36 months of continuous returns, which eliminated 60 firms. The final REIT sample consists of 33,120 firm-month observations and 313 REITs.

Table 2 REIT Summary Statistics

REITs	**1990 to 1994**	**1995 to 2000**	**2001 to 2006**	**2007 to 2012**	**1990 to 2012**
Number of REITs	117	176	151	139	313
Average Number of Observations	55	63	67	70	132
Mean Excess Return (%)	0.55	0.51	1.70	0.68	0.87
Median Excess Return (%)	-0.05	0.06	1.50	0.52	0.66
Standard Deviation	10.29	7.90	7.85	13.04	0.10
Skewness	0.43	0.34	0.01	0.33	0.39
Kurtosis	4.84	4.73	5.23	7.68	8.17

The null hypothesis of no heteroskedasticity was rejected for thirty per cent of the 313 REITs in the study sample with Engle's ARCH test at a 5% significance level. Additionally, twenty-eight per cent of the REITs rejected the null hypothesis of no serial

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[11] Each data series is jointly tested for a structural break by an endogenous trend-break LaGrange Multiplier unit root 2-step procedure. If multiple breaks are detected within a six month interval these are considered a single structural break. If no trend (slope) breaks are identified, the data is tested for breaks in level (intercept) with the Lee and Strazicich (2003) method. If no level breaks are identified, the series is analyzed for stationarity by the Schmidt and Phillips (1992) technique.
[12] Hybrid REITs invest in equity and mortgage trusts.

correlation at the 5% level for with the Breusch-Godfrey test for serial correlation. Forty-three per cent of the REITs displayed serial correlation or heteroskedasticity, while fifteen per cent of the firms exhibited both serial correlation and heteroskedasticity.

Benchmarks

Three benchmarks were employed in the single-index model of each period: the Wilshire Real Estate Index, the NAREIT Index and the Wilshire 5000 Index. The Wilshire Real Estate Index (RESI) is a 'market capitalization-weighted index of publicly traded securities including REITs and real estate operating companies' formed by Wilshire Associates to serve as 'proxies for direct real estate investment by institutions'[13].

REITs comprise the second benchmark selected. The NAREIT Index is a capitalization-weighted index including all tax-qualified publicly-traded REITs such as specialty, mortgage, healthcare, and finite life entities available through NAREIT[14]. For consistency the RESI benchmark is also used for the REIT dataset. The correlation between the RESI and NAREIT benchmark ranges from a low of 0.93 during 1991 to 1997, to a high of 0.99 to 1.00 for the other time periods under study.

To provide a benchmark for the total market, the Wilshire 5000 Index is chosen. The Wilshire 5000 Index represents the broadest index for the U.S. equity market[15]. Over 5,000 capitalization weighted security returns are included in the index. The descriptive statistics for the indices are listed in Table 3.

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[13] Information on Wilshire indices is from the Wilshire website, Wilshire Indexes Methodology, April 2009, http://web.wilshire.com/Indexes/RealEstate. Firms included in the RESI must be an 'operator of commercial and/or residential real estate, which excludes mortgage REITS, mortgage brokers, real estate brokerages, home builders, timber, hybrid REITS and real estate finance companies'. A 'minimum total market capitalization of $200 million is also a requirement of any firm included in the RESI'.

[14] Information on the NAREIT Index is from www.nareit.com

[15] Information on the Wilshire 5000 Index is from http://www.wilshire.com/Indexes/

Table 3 Descriptive Statistics for Indices

RESI	1991 to 1997	1998 to 2004	2005 to 2012	1990 to 2012
Number of Observations	84	84	96	276
Mean Return (%)	1.30	1.03	0.87	0.87
Median Return (%)	1.17	1.47	1.87	1.16
Standard Deviation	3.58	4.25	8.32	5.59
Skewness	0.44	-0.59	-0.60	-0.76
Kurtosis	2.99	4.02	7.67	11.32
NAREIT	**1991 to 1997**	**1998 to 2004**	**2005 to 2012**	**1990 to 2012**
Number of Observations	84	84	96	276
Mean Return (%)	1.59	0.98	0.88	1.03
Median Return (%)	1.67	1.32	1.82	1.29
Standard Deviation	3.31	4.08	7.97	5.91
Skewness	0.55	-0.66	-0.66	-0.66
Kurtosis	3.36	4.49	7.52	10.93
Wilshire 5000	**1991 to 1997**	**1998 to 2004**	**2005 to 2012**	**1990 to 2012**
Number of Observations	84	84	96	276
Mean Return (%)	1.56	0.54	0.49	0.79
Median Return (%)	1.71	1.39	1.11	1.39
Standard Deviation	3.14	5.02	4.69	4.43
Skewness	0.00	-0.61	-0.73	-0.68
Kurtosis	3.16	3.12	4.64	4.16

The real estate benchmarks and the Wilshire 5000 Index have the lowest correlation during the 1998 to 2004 period (Table 4). The RESI and NAREIT indices show an increase correlation from 0.93 in the early period, to 0.99 and 1.00 in the later periods. With the high correlation between NAREIT Index and RESI, the use of RESI for the REIT sample should not limit the explanatory value of the model.

Table 4 Correlation of Benchmarks

1991 to 1997	RESI	NAREIT	W5000
RESI	1		
NAREIT	0.93	1	
W5000	0.50	0.42	1
1998 to 2004	**RESI**	**NAREIT**	**W5000**
RESI	1		
NAREIT	0.99	1	
W5000	0.30	0.27	1

2005 to 2012	**RESI**	**NAREIT**	**W5000**
RESI	1		
NAREIT	1.00	1	
W5000	0.63	0.65	1
1990 to 2012	**RESI**	**NAREIT**	**W5000**
RESI	1		
NAREIT	0.99	1	
W5000	0.37	0.35	1

RESI is the Wilshire Real Estate Index, NAREIT is the NAREIT Index, and W5000 is the Wilshire 5000 Index.

The risk-free rate in this study is calculated as the first difference in the monthly Citi 3-month Treasury bill returns[16]. Structural trend breaks in the risk-free rate are found in February and November of 1994.

Variables

The macroeconomic variables selected represent changes in the financial environment regarding credit conditions, interest rates and financial constraints. The term spread, which captures the influence of the yield curve, is the difference of the yield on U.S. government 10 year and 1 year constant maturity bonds. To capture the changes in the credit conditions, the default spread is the difference in the yield to maturity in the 10 year constant maturity Treasury bond and the Baa corporate bond rates[17] (Hahn and Lee (2006)). These variables are used as proxies for credit market conditions and the monetary policy of the Federal Reserve Board (Hahn and Lee (2006), Keim and Stambaugh (1986), Fama and French (1989), and Kashyap, Lamont

[16] I am indebted to J. Gallo and L. Lockwood for helpful discussion and data sharing. The Citi 3-month Treasury bill is an index that 'measures monthly return equivalents of yield averages that are not mark to market. The Three-Month Treasury Bill Indexes consist of the last three-month Treasury bill issues' copyright Morningstar 2014.
http://awgmain.morningstar.com/webhelp/glossary_definitions/indexes/Citigroup_3-month_T-bill_Index.htm

[17] The Baa data series is from Moodys see www.federalreserve.gov/releases/h15/data.htm

and Stein (1994)). The Baa rates, the 1 year and 10 year constant maturity bond data are from the St. Louis Federal Reserve database.[18]

Following Adrian and Rosenberg (2008), the market skewness and the change in growth in industrial production serve as proxies for short- and long-run volatility. Adrian and Rosenberg (2008) suggest using daily data as the ‘higher frequency data increases the estimation precision’ (Merton (1980)). The skewness estimate is then normalized by dividing the monthly skewness by the cubed standard deviation. This standardization process ‘orthogonalizes the skewness variable to the volatility factor’ (Adrian and Rosenberg (2008))[19].

The daily Wilshire Real Estate Index is used to estimate the monthly real estate return skewness.[20] [21]The Wilshire Real Estate Daily Index is available since March 1996. Therefore, an equally weighted portfolio of publicly-held REITs and real estate firms was created as a proxy for the Wilshire RESI for the early time periods: 1991 to 1997 for real estate mutual funds and 1990 to 1994 for REITs. The market model regressions with the RESI-skewness as the explanatory variable begin in March 1996 and end in December 2012.

Growth in U.S. industrial production is obtained from the St. Louis Federal Reserve data series.[22] [23]

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[18] The term spread revealed trend breaks in July 2000 and March of 2004. The default spread reflects a structural break during February and June of 2009.
[19] White noise innovations in market skewness are residuals from a first order autoregressive model.
[20] The skewness factor for the real estate data indicated trend breaks in October 2000 and January 2001, possibly a reaction to REITs inclusion in the Standard & Poor’s indexes on October 1, 2001.
[21] A broad market skewness estimate using the daily returns from the Wilshire 5000 Index is available upon request.
[22] The change in industrial production growth displayed trend structural breaks in May and August of 1993.
[23] White noise innovations in the industrial production factor are computed as residuals of a third order autoregressive process.

The correlations in the indices and variables are provided for each time period analyzed as well as the overall time span of 1990 to 2012 in Table 5.

Table 5 Correlation Matrices for Variables

1991 to 1997	**Default**	**Term**	**Prod**
Default	1		
Term	0.06	1	
Prod	0.07	-0.15	1

1998 to 2004	**Default**	**Term**	**Prod**	**ReSkew**
Default	1			
Term	-0.37	1		
Prod	0.32	-0.37	1	
ReSkew	0.09	-0.04	0.18	1

2005 to 2012	**Default**	**Term**	**Prod**	**ReSkew**
Default	1			
Term	0.39	1		
Prod	0.06	0.02	1	
ReSkew	0.27	0.04	0.16	1

1990 to 2012	**Default**	**Term**	**Prod**	**ReSkew**
Default	1			
Term	-0.16	1		
Prod	0.24	-0.26	1	
ReSkew	0.08	-0.03	0.16	1

Default is the difference in the yield to maturity in the 10 year constant maturity Treasury bond and the Baa corporate bond rate. Term represents the difference of the yield from the monthly U.S. 10 year and 1 year constant maturity bond. Prod is the change in growth in U.S. industrial production. ReSkew is the standardized monthly skewness from the daily Wilshire RESI returns. All variables are residuals from a monthly autoregressive model from stationary series.

Section 4 Results

This research questions whether the heteroskedasticity and non-normal distribution frequently found in real estate returns contribute to erroneous interpretations as measured by significant alphas. The results reveal attenuation in the firms with significant alphas when the standard errors are adjusted for heteroskedasticity.

However, the response to standard error adjustments by Newey-West and the wild bootstrap HC3 differed among real estate mutual funds and REITs.

A second question includes macroeconomic variables in the model to improve explanatory power and address whether significant alphas are reflected the pricing of changes in these macroeconomic variables. The addition of macroeconomic variables did not systematically decrease the firms with significant alphas.

Heteroskedasticity Consistent Covariance Matrix Estimators

The findings from the single-index model with the Wilshire Real Estate Index (RESI) as the benchmark indicate that the use of the HC1, HC2 or HC3 in the real estate mutual fund sample reduced the per cent of funds with significant Jensen's alpha by 3% during 1998 to 2004 and 1% during 2005 to 2012 (Table 6). The other periods did not reflect any change from the OLS results.

In the REIT sample, the HCCME used in this study provided a more conservative estimate up to 3% of the sample with significant alphas (Table 6). Overall, the use of the HCCME did not inflate the per cent of significant alphas and in some periods provided a more conservative result than OLS standard errors in the single-index model.

Table 6 Per cent of Significant Alphas in the RESI model

	# of Firms	Avg # of Obs	Avg Adj R^2	OLS	HC1	HC2	HC3	NW	WB3T
A. Per cent of Real Estate Mutual Funds with Significant Jensen's Alpha									
1991 to 1997	20	53	0.86	35	35	35	35	25	35
1998 to 2004	52	76	0.91	13	10	10	10	8	10
2005 to 2012	76	85	0.96	14	13	13	13	14	12
1990 to 2012	81	150	0.95	19	19	19	19	19	18
B. Per cent of REITs with Significant Jensen's Alpha									
1990 to 1994	117	55	0.11	13	12	13	12	20	3

1995 to 2000	176	63	0.21	9	7	7	7	13	3
2001 to 2006	151	67	0.28	10	7	7	7	9	1
2007 to 2012	139	70	0.41	3	3	3	3	9	0
1990 to 2012	313	132	0.23	17	17	17	16	21	4

The table shows the per cent of alphas from Real Estate Mutual funds and REITs that are significant at 5%. The model is $R_{i,t} = \alpha_i + \beta_i R_{m,t} + \varepsilon_{i,t}$ where $R_{i,t}$ is the excess return of the Real Estate Mutual fund or REIT over the 1 month risk-free rate and $R_{m,t}$ is the excess return of the Wilshire Real Estate Index. HC1, HC2 and HC3 reflect the standard error with the HC1, HC2 and HC3 adjustment. NW adjusts the standard error with the Newey-West model. WB3T is the Wild Bootstrap with an HC3 adjustment to the standard error and 10,000 repetitions.[24]

Newey-West Standard Errors

The use of the Newey-West adjustment for autocorrelation and heteroskedasticity did not always lead to a more conservative interpretation of the number of significant alphas than the OLS standard errors. The per cent of real estate mutual funds with significant alphas declined with Newey-West standard errors in the 1991 to 1997 period and 1998 to 2004 data periods (Table 6). However with NAREIT or the Wilshire 5000 Index as the benchmarks, the per cent of funds with significant alphas increased with Newey-West standard errors from the OLS, HC1, HC2 and HC3 results (Table 7).

When the alphas from the REIT regressions are reported using Newey-West standard errors the results reflect a four to seven per cent increase in the number of REITs with significant alphas in the 1990 to 1994, 2007 to 2012 and 1990 to 2012 periods. When alternative benchmarks are used Newey-West standard errors continued to inflate the per cent of REITs with significant alphas. This is contradictory to a priori expectation of a more conservative interpretation when implementing the Newey-West standard error adjustment (Newey and West (1987)). A possible

[24] Results with bootstrap standard errors at 1000 repetitions, a robust outlier adjustment to the standard error and a GARCH model are available upon request.

explanation is provided by Wooldridge (2009) who cautioned that the Newey-West standard errors could perform ‘poorly’ in small samples, samples with severe autocorrelation or when the residuals are not positively serially correlated.

Wild Bootstrap with HC3 Standard Errors

The wild bootstrap with HC3 standard errors is included in the analysis to address potential heteroskedasticity in the sample (WB3T in Table 6). Brooks (2008) expresses a concern that using a fixed 5% size of test may lead to an increase in the rejection of the null hypothesis in large samples. In the context of this study 10,000 repetitions are used for the wild bootstrap, potentially creating an increase in the rejection of the null hypothesis. Therefore, to find the 5% chance that a random variable will take the extreme values, the pseudo t-ratios for each 10,000 bootstrap repetitions are sorted and the 2.5 lower percentile and 97.5 upper percentile are used as the critical values for the individual firm[25].

For the post 1997 data the use of the wild bootstrap HC3 reduced the per cent of real estate mutual funds with significant alphas relative to the OLS results. The significant alphas were reduced 3% in the 1998 to 2004 period, 2% in 2005 to 2012 and 1% in the 1990 to 2012 period.

In the REIT samples, the wild bootstrap HC3 results reduce the number of firms with significant estimates of alphas 13% over the 1990 to 2012 period when compared to the OLS results. All other REIT sub-periods report less than 4% per cent with significant alphas. The wild bootstrap with HC3 standard errors provides the more conservative result in the REIT sample.

[25] Initially the data was analyzed at the 5% significance level using the p-value. The percentile method of the t-ratios was suggested by Imre Karafiath.

Model Benchmarks

Three benchmarks are compared for each time period in the study for real estate mutual funds (Table 7) and REITs (Table 8). The average adjusted R^2s for real estate mutual funds using the real estate indices only varied minimally (Table 7) and were equivalent during all periods for the REIT sample (Table 8). However the number of funds with significant alphas did vary with the choice of benchmark.

The benchmark representing the total market is the Wilshire 5000 Index. The explanatory power was the lowest in the early periods and increased in subsequent periods for real estate mutual funds and REITs. However, the explanatory power of the Wilshire 5000 Index was much less than the real estate indices (Table 7 & 8).

Real Estate Funds and Differing Benchmarks

The per cent of real estate funds with significant alphas at the 5% level varied by benchmark and time period. With the Wilshire 5000 Index as the benchmark, the unadjusted OLS model for the 1998 to 2004 time period indicated 23% of the real estate mutual funds had significant alphas (Table 7). The next period of 2005 to 2012 indicated no firms earned a significant alpha relative to a passive investment in the total market. Over the 23 year study of 1990 to 2012, only 2% of the real estate mutual funds provided a significant alpha as reflected in the OLS results. The wild bootstrap with HC3 standard errors reduced the per cent of the sample with significant alphas by 5% from 1991 to 1997 and 17% in the 1998 to 2004 time period. The wild bootstrap with the HC3 standard errors in the last sub-period (2005 to 2012) and the full period (1990 to 2012) indicated zero funds with significant alphas with the Wilshire 5000 benchmark.

Table 7 Per cent of Real Estate Mutual Funds with significant alphas using different benchmarks

	Avg Adj R^2	OLS	HC1	HC2	HC3	NW	WB3T
1991 to 1997							
Wilshire 5000	0.13	10	10	10	10	20	5
NAREIT	0.87	0	0	0	0	5	0
RESI	0.86	35	35	35	35	25	35
1998 to 2004							
Wilshire 5000	0.1	23	23	23	23	23	6
NAREIT	0.91	15	10	10	10	12	10
RESI	0.91	13	10	10	10	8	10
2005 to 2012							
Wilshire 5000	0.66	0	0	0	0	0	0
NAREIT	0.96	17	17	16	16	17	15
RESI	0.96	14	13	13	13	14	12
1990 to 2012							
Wilshire 5000	0.45	2	2	2	2	2	0
NAREIT	0.94	15	15	15	14	14	11
RESI	0.95	19	19	19	19	19	18

The table shows the per cent of Real Estate Mutual funds that are significant at 5%. The model is $R_{i,t} = \alpha_i + \beta_i R_{m,t} + \varepsilon_{i,t}$ where $R_{i,t}$ is the excess return of the Real Estate Mutual fund or REIT over the 1 month risk-free rate and $R_{m,t}$ is the excess return of the Wilshire 5000 Index (Wilshire 5000), NAREIT Index, or Wilshire Real Estate Index (RESI). HC1, HC2 and HC3 reflect the standard error with the HC1, HC2 and HC3 adjustment. NW adjusts the standard error with the Newey-West model. WB3T is the Wild Bootstrap with an HC3 adjustment to the standard error and 10,000 repetitions.

Using the NAREIT Index or RESI as the benchmark and comparing the OLS results, the per cent of significant alphas varied by time period for real estate funds and REITs (Table). This is unexpected since the indexes are highly correlated (Table 4) and the explanatory power (as measured by the adjusted R^2) of the models is comparable.

Table 8 Per cent of REITs with significant alphas using different benchmarks

	Avg Adj R^2	OLS	HC1	HC2	HC3	NW	WB3T
1990 to 1994							

Wilshire 5000	0.05	8	9	9	9	17	3
NAREIT	0.11	7	9	9	9	13	3
RESI	0.11	13	12	13	12	20	3
1995 to 2000							
Wilshire 5000	0.03	9	7	7	7	9	3
NAREIT	0.21	9	8	8	7	13	3
RESI	0.21	9	7	7	7	13	3
2001 to 2006							
Wilshire 5000	0.07	48	49	49	47	53	3
NAREIT	0.28	8	7	7	7	9	1
RESI	0.28	10	7	7	7	9	1
2007 to 2012							
Wilshire 5000	0.31	2	1	1	1	6	0
NAREIT	0.41	3	4	4	3	8	0
RESI	0.41	3	3	3	3	9	0
1990 to 2012							
Wilshire 5000	0.1	7	7	7	6	12	4
NAREIT	0.23	14	14	14	14	19	3
RESI	0.23	17	17	17	16	21	4

With real estate indices as the benchmarks for real estate mutual funds and REITs, the use of HC1, HC2 and HC3 reduced the per cent of significant alphas up to 5% or indicated no change from the OLS (Table 7 and 8). In REITs the HC3 standard errors proved a more conservative or equivalent number of significant alphas than OLS, HC1 or HC2.

When Newey-West standard errors are implemented with the real estate benchmarks, during 1991 to 1997, the NAREIT benchmark regressions showed an increase in the per cent of funds with significant alphas for real estate mutual funds (Table 7). In the REIT sample the use of Newey-West standard errors increased the per cent of significant alphas in all periods (Table 8).

The wild bootstrap with the HC3 standard errors for the real estate mutual fund and REIT data provided a more conservative or an equivalent result to the OLS regressions in all time periods (Table 7 and 8).

Model Specification

The addition of the change in the default and term spread did not systematically change the number of real estate mutual funds with significant alphas. For the REIT sample the change in default spread and term spread increased the explanatory power in the all the sub-periods, howbeit minimally.

Table 9 Average Adjusted R^2

Real Estate Mutual Funds	OLS	Default and Term	RESI short- and long-run volatility
1991 to 1997	0.86	0.87	0.86
1998 to 2004	0.91	0.92	0.91
2005 to 2012	0.96	0.96	0.96
1990 to 2012	0.95	0.95	0.98
REITs			
1990 to 1994	0.11	0.13	0.11
1995 to 2000	0.21	0.22	0.22
2001 to 2006	0.28	0.30	0.28
2007 to 2012	0.41	0.42	0.40
1990 to 2012	0.23	0.24	0.27

The addition of proxies for short and long-run volatility increased the average adjusted R^2 for real estate mutual funds and REITs only for the 1990 to 2012 period (Table 9). In real estate mutual funds the adjusted R^2 increased from 0.95 in the single-index model to 0.98 when RESI short-run and long-run volatility factors are included in the model[26]. These variables improved explanatory power for REITs from 0.23 to 0.27.

The inclusion of the macroeconomic variables to potentially improve the explanatory power of the model provides a secondary insight. Any attenuation in the significant alpha would suggest that seemingly superior performance of the fund is attributable to these variables. The use of HC1, HC2, HC3, Newey-West and the wild bootstrap HC3 standard errors compares the number of significant alphas when returns are adjusted for heteroskedasticity.

The addition of default and term spread over most time periods did not appreciably change the number of firms with significant alphas (Table10)[27]. The exception being the 1991 to 1997 period. The number of significant alphas in the real estate mutual funds declined from 35 per cent to 30 per cent (Table 10).[28] The influence is limited to this period, suggesting a unique environment, anomaly or smaller variations in the spread during later periods. The HCCME variants when the default and term spread are included in the model indicate the HC3 adjustment provides a more conservative interpretation of the per cent of funds with significant alphas (Table 10). The wild bootstrap with HC3 standard errors provides the most conservative result for REITs and real estate mutual funds in most cases.

[26] Similar increase in real estate mutual fund in adjusted R^2 was found when the innovation from the Wilshire 5000 short-run volatility factor is included in the model.
[27] Results from the default and term spread included individually to the model is available upon request.
[28] The number of significant alphas in the real estate mutual funds declined from 35 per cent to 30 per cent when the change in default spread is included in the model. The average adjusted R^2 increased from 0.86 to 0.95 in this period.

Table 10 Term and Default Spread

	Avg Adj R^2	RESI OLS	OLS	HC1	HC2	HC3	NW	WB3T
Real Estate Mutual Funds								
1991 to 1997	0.87	**35**	30	40	40	25	25	35
1998 to 2004	0.92	**13**	17	15	15	12	12	12
2005 to 2012	0.96	**14**	16	13	13	12	16	11
1990 to 2012	0.95	**19**	20	20	17	17	20	16
REITs								
1990 to 1994	0.13	**13**	15	15	15	12	22	3
1995 to 2000	0.22	**9**	10	10	10	9	10	3
2001 to 2006	0.30	**10**	8	8	7	5	12	1
2007 to 2012	0.42	**3**	5	5	5	4	12	0
1990 to 2012	0.24	**17**	14	15	14	13	18	4

The table shows the per cent of alphas from Real Estate Mutual funds and REITs that are significant at 5%. The model is $R_{i,t} = \alpha_i + \beta_i R_{m,t}$ + term spread$_t$ + default spread$_t$ + $\varepsilon_{i,t}$ where $R_{i,t}$ is the excess return of the Real Estate Mutual fund or REIT over the 1 month risk-free rate and $R_{m,t}$ is the excess return of the Wilshire Real Estate Index. The RESI OLS column provides the comparable to the single-index model OLS results. HC1, HC2 and HC3 reflect the standard error with the HC1, HC2 and HC3 adjustment. NW adjusts the standard error with the Newey-West model. WB3T is the Wild Bootstrap with an HC3 adjustment to the standard error and 10,000 repetitions.

Table 11 shows the per cent of funds with significant alphas at the 5% level when the real estate index market skewness factor and the change in industrial production are included in the model (proxies for short- and long-run volatility). From 1996 to 2012 the addition of the volatility factors reduced the per cent of significant alphas in real estate mutual funds and REITs. Of the HCCME compared, the HC3 is the most conservative while the Wild Bootstrap with HC3 standard errors provides the most conservative interpretation.

Table 11 Wilshire Real Estate Index short-run and long-run volatility

	Avg Adj R^2	RESI OLS	OLS	HC1	HC2	HC3	NW	WB3T
Real Estate Mutual Funds								
1991 to 1997*	0.86	**35**	30	30	30	30	30	25
1998 to 2004	0.91	**13**	17	10	10	8	8	10
2005 to 2012	0.96	**14**	7	11	9	7	11	7
1996 to 2012	0.98	**15**	12	14	12	10	12	10
REITs								
1990 to 1994*	0.11	**13**	3	8	7	5	10	27
1995 to 2000	0.22	**9**	10	10	9	7	11	4
2001 to 2006	0.28	**10**	12	9	9	7	11	1
2007 to 2012	0.40	**3**	7	9	9	6	15	9
1996 to 2012	0.27	**9**	7	8	7	7	11	4

The table shows the per cent of alphas from Real Estate Mutual funds and REITs that are significant at 5%. The model is $R_{i,t} = \alpha_i + \beta_i R_{m,t}$ + RESI short-run volatility$_t$ + long-run volatility$_t$ + $\varepsilon_{i,t}$ where $R_{i,t}$ is the excess return of the Real Estate Mutual fund or REIT over the 1 month risk-free rate and $R_{m,t}$ is the excess return of the Wilshire Real Estate Index. The innovation from the skewness from the Wilshire Real Estate Index proxy for the short-run volatility. The innovation of the change in growth industrial production proxy for the long-run volatility. The RESI OLS column provides the comparable to the single-index model OLS results with the RESI benchmark. HC1, HC2 and HC3 reflect the standard error with the HC1, HC2 and HC3 adjustment. NW adjusts the standard error with the Newey-West model. WB3T is the Wild Bootstrap with an HC3 adjustment to the standard error and 10,000 repetitions. *An equally weighted portfolio of REITs and publicly traded real estate firms was created and the standardized skewness factor calculated from daily returns for 1990 to February 1996.

Section 5 Conclusion

This study addresses the question if superior significant alphas in real estate funds as reported are in fact attributable to bias from data with non-constant error variances. HC1, HC2 and HC3 standard errors were calculated in the single-index model to address the heteroskedasticity reported in real estate returns (Young (2008)). Results comparing the number of funds with significant alphas in the OLS model with no standard error adjustments ranged from no change to a 3% decline when the HCCME used in this study are implemented.

From the literature (MacKinnon (2013)) there was an expectation of a reduction in size distortion with the application of HC1, HC2 or HC3 standard errors. The empirical data in comparing the HC1 to the HC2 to the HC3 produced no systematic changes in the significance of alpha of real estate mutual funds. In REITs the HC3 reduced the number of firms with significant alphas by 1% in two periods. When additional economic variables are included in the model the HC3 standard error provided a more conservative result with up to 3% fewer significant alphas than the unadjusted OLS result.

When real estate mutual fund alphas are evaluated with the Newey-West standard errors highly variable changes are obtained in comparison with the OLS results. When the Newey-West standard error results are compared for the REITs, their results often show a larger number of significant alphas than the OLS results. Since several studies reviewed for this research reported Newey-West results, for example Wiley and Zumpano (2009), Kallberg, et.al (2000), Cici, Corgel, and Gibson (2011) and Kaushik et.al. (2013), the use of Newey-West standard errors in studies of real estate returns would warrant further examination.

In the first use of the wild bootstrap with an HC3 adjustment for heteroskedasticity the real estate mutual fund sample reduced the per cent of funds with significant alphas by up to 3% in the single index model and up to 7% in the multifactor models. When applied to REIT data, the wild bootstrap HC3 standard errors reduced the number of firms with significant Jensen's alpha as much as 13%. The REIT and the real estate mutual fund results would suggest that the use of the wild bootstrap HC3 standard errors should be considered in real estate research since this test provided the most conservative result. This is consistent with Davidson and Flachaire (2008) Monte Carlo simulation and MacKinnon's (2013) recommendation.

This research is possibly the first use of market skewness as a proxy for short-run volatility (as described by Adrian and Rosenberg (2008)) as an explanatory with real estate fund returns as the dependent variable. The results would indicate over the longer period the RESI skewness variable attenuates some of the significant alphas for real estate firms. This would suggest that part of the alphas in real estate mutual funds may be attributable to real estate market skewness as a proxy for financial constraints in the real estate market.

The conclusion that the benchmark makes a difference is not a new finding (Roll (1978) and Hartzell, Muhlhofer and Titman (2010)). For real estate mutual funds, while the average adjusted R^2 of the two real estate benchmarks (Wilshire Real Estate Index (RESI) and NAREIT) were within 1/100 of each other, the number of firms that had significant alphas varied. In the 1990s with RESI as the benchmark in the model a higher number of firms (35%) exhibited significant alphas than the same firms with the NAREIT benchmark (0). The order is reversed during the 2005 to 2012 period where there is a higher per cent of funds with significant alphas when the NAREIT index is the benchmark.

This research provides a comparison of several adjustments to the estimated standard errors of regression parameters in a dataset with non-constant variance: real estate returns. The results indicate minimal reduction in the number of significant Jensen's alpha with the use of HC1, HC2 or HC3 standard errors. Among the HCCME used in this study, the HC3 standard error provided the most conservative results. The

Newey-West standard error provides variable and often a larger number of significant alpha results than OLS. The wild bootstrap with HC3 standard errors provides a more conservative standard error when compared to OLS and other measures of standard errors. Based on this research, the wild bootstrap with HC3 standard errors is recommended as the most conservative approach to testing for superior performance by REIT or real estate mutual fund managers.

REFERENCES

Adrian, T., & Rosenberg, J. (2008). Stock Returns and Volatility: Pricing the Short-Run and Long-Run Components of Market Risk. The Journal of Finance, 63(6), 2997-3030.

Brooks, Chris (2008). Introductory Econometrics for Finance. Second Edition, Cambridge University Press,IBSN 978-0-521-69468-1.

Cheng, P. (2005). Asymmetric risk measures and real estate returns. *The Journal of Real Estate Finance and Economics*, *30*(1), 89-102.

Chiang, K., K. Kozhevnikov, M. Lee and C. Wisen. 2008. Further Evidence on the Performance of Fund of Funds: The Case of Real Estate Mutual Funds. *Real Estate Economics* 36: 47–61.

Cici, G., Corgel, J., & Gibson, S. (2011). Can fund managers select outperforming REITs? Examining fund holdings and trades. *Real Estate Economics*, *39*(3), 455-486.

Davidson, R., & Flachaire, E. (2008). The wild bootstrap, tamed at last. Journal of Econometrics, 146(1), 162-169.

Dellya, W. L., DeMaskey, A. L., & Smith, C. A. (2001). Selectivity and market timing performance of fidelity sector mutual funds. *Financial Review*, *36*(1), 39-54.

Engle, R., 1982. Autoregressive Conditional Heteroskedasticity with Estimates of the Variance of United Kingdom Inflation. *Econometrica* 50: 987-1007.

Fama, E. F., & French, K. R. (1989). Business conditions and expected returns on stocks and bonds. Journal of financial economics, 25(1), 23-49.

Fama, E. F., & French, K. R. (2010). Luck versus Skill in the Cross-Section of Mutual Fund Returns. *The Journal of Finance*, *65*(5), 1915-1947.

Gallo, J. G., Lockwood, L. J., & Rutherford, R. C. (2000). Asset allocation and the performance of real estate mutual funds. *Real Estate Economics*, *28*(1), 165-185.

Hahn, J., & Lee, H. (2006). Yield spreads as alternative risk factors for size and book-to-market. Journal of Financial and Quantitative Analysis, 41(02), 245-269.

Hartzell, J. C., Mühlhofer, T., & Titman, S. D. (2010). Alternative benchmarks for evaluating mutual fund performance. *Real Estate Economics*, *38*(1), 121-154.

Jensen, Michael C., 1968, The performance of mutual funds in the period 1945–1964, *Journal of Finance* 23, 389–416.

Kallberg, J., C. Liu and C. Trzcinka. 2000. The Value Added from InvestmentManagers: An Examination of Funds of REITs. *Journal of Financial and Quantitative Analysis* 35:387–408.

Kaushik, A., & Pennathur, A. K. (2013). Performance And New Money Cash Flows In Real Estate Mutual Funds. *Journal of Financial Research*, *36*(4), 453-470.

Lee, J., & Strazicich, M. C. (2003). Minimum Lagrange multiplier unit root test with two structural breaks. *Review of Economics and Statistics*, *85*(4), 1082-1089.

Lee, J., Strazicich, M. C., & Meng, M. (2012). Two-Step LM Unit Root Tests with Trend-Breaks. *Journal of Statistical and Econometric Methods*, *1*(2), 81-107.

Long, J. S., & Ervin, L. H. (2000). Using heteroscedasticity consistent standard errors in the linear regression model. *The American Statistician*, *54*(3), 217-224.

MacKinnon, J. G., & White, H. (1985). Some heteroskedasticity-consistent covariance matrix estimators with improved finite sample properties. Journal of Econometrics, 29(3), 305-325.

MacKinnon, J. G. (2013). Thirty years of heteroskedasticity-robust inference. In *Recent Advances and Future Directions in Causality, Prediction, and Specification Analysis* (pp. 437-461). Springer New York.

Merton, Robert C., 1980, On estimating the expected return on the market: An exploratory investigation, Journal of Financial Economics 8, 323-361.

Newey, Whitney K. and Kenneth D. West, 1987,A simple, positive semi-definite, heteroskedasticityand autocorrelation consistent covariance matrix, Econometrica 55, 703–708.

Rodriguez, J. 2007. A Critical Look at the Forecasting Ability of Real Estate Mutual Fund Managers. Journal of Real Estate Portfolio Management 13: 99–106.

Schmidt, P., & Phillips, P. C. (1992). LM TESTS FOR A UNIT ROOT IN THE PRESENCE OF DETERMINISTIC TRENDS*. Oxford Bulletin of Economics and Statistics, 54(3), 257-287.

White, H. (1980). A heteroskedasticity-consistent covariance matrix estimator and a direct test for heteroskedasticity. *Econometrica: Journal of the Econometric Society*, 817-838.

Wiley, Jonathan, and Leonard Zumpano (2009). Institutional Investment and the Turn-of-the-Month Effect: Evidence from REITs Journal of Real Estate Finance and Economics 39:180-201.

Woodridge, Jeffrey M. (2009). Introductory Econometrics, Fourth Edition. Cengage learning. IBSN: 978-0-324-66054-8

Young, M. S. (2008). Revisiting non-normal real estate return distributions by property type in the US. The Journal of Real Estate Finance and Economics, 36(2), 233-248.