Noise vs. News In Equity Returns

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Abstract

What role does noise play in equity markets? Answering this question usually leads immediately to specifying a model of fundamentals and hence the pervasive joint hypothesis quagmire. We avoid this dilemma by measuring noise volatility directly by focusing on the behavior of country closed-end funds (CCEFs) during foreign (i.e., non-U.S.) holidays – for example, the last days of Ramadan in Islamic countries. These holiday periods are times when the flow of fundamental information relevant to foreign equity markets is substantially reduced and hence trading of CCEFs in U.S. markets can be responding only weakly, if at all, to fundamental information. We find that, controlling for the effects of industry and global shocks and of the overall U.S. market, there remains a substantial amount of noise in the equity returns of U.S. CCEFs. In the absence of noise, the noise ratio statistic would be near zero. However, our results indicate statistically significant departures from zero, with values averaged over all U.S. CCEFs ranging from 68% to 77% depending on assumptions about the leakage of information during holiday periods and kurtosis. Noise is negatively related to institutional ownership of U.S. CCEFs and is much less important for U.K. CCEFs. The lower levels of noise for matched U.K. and U.S. CCEFs provide some initial evidence that the U.K. securities transaction tax is effective in reducing stock market noise.

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Table Of Contents

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> The quick and accurate reaction of security prices to information, as well as the non-reaction to non-information, are the two broad predictions of the efficient markets hypothesis.
>
> Shleifer (2000, p. 5)

> Perhaps no single empirical issue is of more fundamental importance to both the fields of financial economics and macroeconomics than the question of whether or not stock prices are a well-informed and rational assessment of the value of future earnings available to stockholders.
>
> Fischer and Merton (1984, p. 94)

> ... it makes a great deal of difference to an investment market whether or not they [skilled long-term investors] predominate in their influence over the game-players.
>
> Keynes (1936, p. 156)

I. Introduction

What role does noise play in equity markets? Textbooks provide elegant treatments that describe how rational investors efficiently use information to value equities and extract the subsequent implications for equity prices. In the 1980's, research began to suggest that there is substantial distance between this textbook view and the actual performance of equity markets. For example, Shiller (1981) LeRoy and Porter (1981), and West (1988) show that equity prices are too volatile to be consistent with a net present value model based on fundamentals. Cutler, Poterba, and Summers (1989) and Roll (1984, 1988) document the surprising inability of fundamental economic variables and important macroeconomic news known ex-post to explain equity returns. Many other studies – surveyed by LeRoy (1989), Shiller (1990), and Campbell, Lo, and MacKinlay (1997) and collected in Thaler (1993, 2005) – have highlighted several important problems with the view that equity prices conform to the two broad predictions of the efficient markets hypothesis mentioned by Shleifer in the above quotation.[1]

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[1] See Malkiel (2003) and Shiller (2003) for an informative exchange about the status of the efficient markets hypothesis.

A quantitatively important role for noise -- market outcomes based on non-information (Black, 1986) or on "heuristics rather than Bayesian rationality" (Shleifer, 2000, p. 12) -- raises serious doubts as to whether equity markets are able to discharge their central function of allocating resources.[2] Noise disrupts equity markets in several ways. Discount rates are raised by noise, thus lowering the level of the equilibrium capital stock and pressuring managers to adopt a short-term focus.[3] If noise is concentrated in certain industries (e.g., those where intangible capital and information technologies are important), resources directed through equity markets may be seriously misallocated. Noise raises the cost of risk arbitrage and thus curbs the capacity of arbitrageurs to align market values with fundamental values. If Keynes' "game players" loom large, equity markets will be unable to allocate capital efficiently.[4] Moreover, noise creates deadweight losses, as real resources need to be expended to disentangle noise from fundamentals. Noise-induced movements in equity prices may lead to the onset of bubbles, and bursting bubbles have substantial impacts on real spending.[5] This channel is amplified by the increasing proportion of assets held in equities and may well disrupt efforts by policymakers to stabilize the macroeconomy (Issing, 2004).

How much of the volatility in equity returns can be attributed to noise relative to news? Addressing this question usually leads immediately to specifying a model of fundamentals and the pervasive joint hypothesis quagmire (Fama, 1970, 1991), which implies that, in any asset

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[2] This list of problems stemming from noise is drawn from Shleifer and Summers (1990), Shiller (1990), and Shleifer (2000).

[3] De Long, Shleifer, Summers and Waldman (1990) and Campbell and Kyle (1993) present models of noise trading risk, and Lee, Shleifer, and Thaler (1991) provide theoretical arguments and empirical evidence that noise qua investor sentiment is a systematic risk factor raising the required equity return. This result is controversial; see the spirited exchange between Chen, Kan and Miller (1993) and Chopra, Lee, Shleifer, and Thaler (1993). Even if noise is idiosyncratic, it can raise the required equity return through unbalanced portfolios (Levy, 1978; Malkiel and Xu, 2002), incomplete information (Merton 1987a), or information risk (Easley, Hvidkjaer, and O'Hara, 2002). Goyal and Santa-Clara (2003) document that idiosyncratic risk predicts aggregate returns. Their key result is confirmed by Ball, Cakici, Yan, and Zhang (2005) and Brown and Ferreira (2005), who argue that it is limited to small firms.

[4] Inefficiencies created by noise are the driving force in the "fundamental indexation" approach to investing advanced recently by Arnott, Hsu, and Moore (2005) and Siegel (2006).

[5] Lei, Noussair, and Plott (2001) generate experimental evidence showing how noise impacts bubbles, and Baker, Stein, and Wurgler (2003), Chirinko and Schaller (2001, 2008), Gilchrist, Himmelberg, and Huberman (2005), and Polk and Sapienza (forthcoming) document the effects of equity misvaluations on investment in property, plant, and equipment.

pricing model, the volatility attributable to noise can equally well be interpreted as the volatility attributable to specification error. This criticism necessarily confronts prior studies decomposing returns into fundamental and noise components. These studies calculate noise by estimating the volatility attributable to fundamentals and then subtracting this estimate from overall volatility. Since noise volatility is a residual in this procedure, specification errors in the fundamental model are impounded in the subsequent estimate of noise. Noise in the eyes of one researcher is specification error in the eyes of another.

We avoid this quagmire by measuring noise volatility directly. We are able to employ a direct estimation strategy by focusing on the behavior of country closed-end funds (CCEFs) during holidays in foreign markets, where the latter is defined for purposes of this paper as the local equity market for countries other than the United States (or the United Kingdom in Section V). Holiday periods – for example, the last days of Ramadan in Islamic countries – are times when the flow of fundamental information relevant to foreign equity markets is substantially reduced. Consequently, trading of CCEFs in U.S. markets can be responding only weakly, if at all, to fundamental information flowing from the foreign markets. CCEF return volatility during foreign holidays thus provides an accurate estimate of noise in the highly developed equity markets of the United States and the United Kingdom. Our approach is most closely related to the pioneering study of French and Roll (1986), who use the closing of the New York Stock Exchange on Wednesday afternoons during the second half of 1968 to separate news from noise by “turning-off” the noise channel. In contrast, the “natural experiment” afforded by foreign holidays turns-off (or tones-down) the news channel, thus allowing for a more direct estimate of equity market noise.

This paper evaluates the importance of noise by estimating return variances from three nested models of increasing generality. Section II models the return distribution for individual CCEFs as a mixture of the underlying distributions for holidays and standard days, where the former depends on the noise variance while the latter depends on both noise and information variances. The mixing variable is the exogenous occurrence of a foreign holiday. The estimation problem is simplified by the non-stochastic nature of the mixing variable, and maximum likelihood and method-of-moments estimates are equivalent. Based on 54 CCEFs traded in the United States and controlling for the effects of industry shocks and the overall U.S. market, we compute a noise ratio statistic – NR, the noise variance divided by the sum of the noise and information variances – and find that, for CCEF returns adjusted for exchange rates and the overall U.S. market, the mean of NR is 0.767. A likelihood ratio test against the null hypothesis of a noise ratio of 0.10 is rejected at the 1% level for 52 of the 54 funds in the sample.

An attractive feature of these baseline results is that they rely on an exogenous instrument (holidays) and few modeling assumptions, thus largely avoiding the joint hypothesis quagmire. The only two assumptions are the normality of returns and a substantial reduction in the flow of information on holidays. Each of these assumptions is relaxed in Section III. In order to account for the kurtosis that characterizes equity returns, we replace the normal distribution underlying the baseline results with a Scaled-t distribution. The mean NR rises to 0.752 and is statistically different from a noise ratio of 0.10 at the 1% level for all 54 funds. We then allow for the possibility that some information germane to CCEF returns is nonetheless transmitted during a holiday and develop a double mixture model, where the second mixture represents the possibility of information "arrival" or "leakage" during holidays. Under this more general model, the mean NR falls to 0.673. Formal hypothesis tests indicate that the NR is significantly far from 0.1; this null hypothesis is rejected at the 1% level for 49 of the 54 CCEFs and at the 5% level for all but one of the funds. Thus, our general conclusion about the importance of noise is largely robust to the leakage of some information during holidays.

Section IV briefly examines the role of institutional owners. These financially sophisticated and patient investors are presumably less susceptible to the "sentiments" and "irrationalities" that drive noise trading. We find some support for this hypothesis, as greater institutional ownership of CCEFs is associated with less noise. This result is economically and statistically significant, and the elasticity of noise with respect to institutional ownership is between -0.163 and -0.193.

Section V repeats the analysis for 32 CCEFs traded in the United Kingdom. These results provide insights into the potency of a securities transaction tax that affects trades in the United Kingdom but not in the United States. We find that the NRs for U.K. CCEFs are generally much lower than those for U.S. CCEFs. This result continues to hold when we compare U.K. to U.S. CCEFs that invest in the same country. These results for matched funds provide some initial evidence that the securities transactions tax lowers noise volatility.

Section VI discusses several key papers that have assessed noise and its role in equity markets, and Section VII concludes.

II. Baseline Evidence

Our estimation strategy permits us to estimate noise directly. We exploit the natural experiment produced by holidays that occur in a foreign country but that are not held at the same time in the United States. During these holiday periods, the flow of fundamental information relevant to securities in the foreign market is substantially reduced. Consequently, trading of CCEFs in the U.S. markets can be responding only weakly, if at all, to fundamental information. CCEF return volatility (with suitable adjustments discussed later in this section) during foreign holidays thus provides an accurate estimate of noise in the U.S. equity market.

We use a mixture model of returns and our identifying assumption to generate maximum likelihood estimates of the proportion of return variance due to noise. We assume that, in general, the CCEF return on a standard day or holiday ($R_{d,t}, d = H,S$) can be decomposed into two orthogonal components due to noise (R_t^N) and information (R_t^I),

$$R_{d,t} = R_t^N + R_t^I \qquad d = H,S, \qquad \forall t. \tag{1}$$

On standard days (non-holidays), CCEF returns are affected by both noise and information. By contrast, the identifying assumption that the information flow is largely shut-off during holidays implies that R_t^I has at most a modest impact on holiday returns. Thus, R_t is a mixture of distributions that apply during holidays and standard days with mixing weights determined by the occurrence of holidays,

$$\begin{aligned} R_t &= \rho_t R_{H,t} + (1-\rho_t) R_{S,t} \qquad \forall t, \\ &= \rho_t R_t^N + (1-\rho_t)(R_t^N + R_t^I) \end{aligned} \tag{2}$$

where the mixing weight, ρ_t, is one for holidays and zero for standard days. That the mixing weights are known and exogenous substantially simplifies the estimation problem. In this case, the likelihood function is additively separable over holidays and standard days, and the conventional maximum likelihood estimator is identical to a method of moments estimator (see the Appendix, Section 3). Volatility is measured in terms of variances, and the variance of the return ($V[R_t]$) during holidays (V_H) and standard days (V_S) is related to the noise variance (V^N) and the information variance (V^I) as follows,

$$V[R_t] \equiv V_H = V^N \qquad \forall t \in \mathfrak{I}^H \text{ where } \mathfrak{I}^H = \{t : t \text{ is a holiday}\}, \tag{3a}$$

$$V[R_t] \equiv V_S = V^N + V^I \qquad \forall t \in \mathfrak{I}^S \text{ where } \mathfrak{I}^S = \{t : t \text{ is a standard day}\}. \tag{3b}$$

Note that the subscripts on V_H and V_S denote sample moments, while the superscripts on V^N and V^I denote unknown parameters entering the likelihood function. We assume that returns are distributed normal (an assumption that will be relaxed in Section III) and are mean zero (to be discussed below). The impact of noise is assessed by a noise ratio statistic, NR, defined for each CCEF as follows,

$$NR \equiv V^N / (V^N + V^I). \tag{4a}$$

To estimate the V^N and V^I parameters in equation (4a), we need data on foreign holidays, returns, and exchange rates. Holidays are identified by zero daily positive trading volume of a broad foreign market index (e.g. the Nikkei 225) and trading in U.S. markets. In several cases, we confirmed the accuracy of this procedure by cross-checking the holidays identified by the two criteria against the list of fixed and algorithmic holidays in Weaver (1995). Thus, holidays are when business activity and hence the flow of relevant information are substantially reduced in the foreign country.

Returns are drawn from the CRSP database on Wharton Research Data Services (WRDS) for the period beginning January 1, 1980 (or the date that the CCEF was first listed) and ending December 31, 2004 (or the date that the CCEF was delisted). CCEFs that cover more than one country (with one exception noted in Table 1) are excluded. Table 1 lists the 54 CCEF funds for the U.S. used in this study, and the table notes provide additional details about data collection.

Daily exchange rate data are obtained from Datastream and augmented where missing with exchange rate data from the Federal Reserve Bank of St. Louis. The return series used to compute the sample moments (V_H and V_S) are adjusted for movements in exchange rates by subtracting the log difference daily exchange rate from the CCEF return.

A second adjustment is made to the return series with respect to market-wide factors. The sensitivity of CCEF returns to the U.S. market has been documented by Hardouvelis, La Porta, and Wizman (1994, Table 8.8A) and Bodurtha, Kim, and Lee (1995, Table 8). In estimating V^N, we remove this effect by estimating for each country a market model of the exchange-rate-adjusted CCEF return against a constant and the return on the S&P 500.[6] The residuals from this regression define the mean zero return series, R_t, used to compute V^N in equation (3). Note that this estimate is a lower bound for V^N, as that part of noise correlated with the market is removed by this market model. This lower bound estimate is represented as $\underline{V}^N$. Relying on the market model residuals to estimate the denominator of the noise ratio statistic, $(V^N + V^I)$ during standard days, would also estimate a lower bound. However, since this sum enters the denominator of the noise ratio statistic in equation (4), our estimate of NR would be upwardly biased. To avoid this problem and ensure that the reported estimate of NR is indeed a lower bound estimate of the fraction of overall volatility attributable to noise, we compute $(V^N + V^I)$ during standard days based on CCEF returns that are only adjusted for exchange rates. Since some of the correlation between noise and the market return will remain, this estimate is an upper bound of overall volatility and is represented as $(\overline{V}^N + \overline{V}^I)$. The estimated NR is therefore a lower bound for the true noise ratio statistic, NR^*, and NR is estimated as follows,

$$NR \equiv \underline{V}^N / (\overline{V}^N + \overline{V}^I) < NR^* . \tag{4b}$$

Maximum likelihood estimates of NR and the underlying estimates of $\underline{V}^N$ and $(\overline{V}^N + \overline{V}^I)$ with their standard errors are presented in Table 2.[7] In the absence of noise, the NR statistic would be zero apart from sampling error. However, the NRs in column 2 are in most

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[6] The constant terms (α's) are uniformly negative with a mean value of -0.0005, and the slope coefficients (β's) range from -0.130 to 1.237 with a mean value of 0.667. The results for the market model and the noise ratios reported below are robust to defining the market return with or without dividends.

[7] Details of the computation of $\overline{V}^I$ are discussed Section 5 of the Appendix.

cases very far from zero with a mean value of 0.767. These results are not driven by outliers and are robust to trimming 1% of the upper and lower tails (2% in total) of the holiday return distribution and then a similar trimming of the standard day distribution. Table 2 documents that there is a substantial amount of noise in the equity returns of CCEFs traded in the United States.

Ten of the estimated NRs are unity. This maximum value reflects a corner solution in which V^I is on the boundary of zero and occurs for those CCEFs whose holiday variance exceeds the standard day variance. Since our model assigns all of the holiday variance to noise and allows noise to also affect the standard day variance, no role remains for information. These extreme situations may well reflect sampling error and, interestingly, do not occur for comparable estimates of UK funds presented in Section V. Most importantly, these tests are not problematic for tests of noise because the holiday variance exceeding the standard day variance clearly rejects the notion that noise is unimportant.

We formally evaluate the null hypothesis that the NR is 0.10 (thus allowing some noise to affect the estimates) against the maximum likelihood values in Table 2.[8] Details of this likelihood ratio test are presented in Section 4 of the Appendix. Hypothesis tests indicate that the NRs are very far from 0.10; the null hypothesis is rejected at the 1% level for 52 of the 54 funds.

Complementary evidence is provided by examining trading volume on holiday and standard days. In this case, the noise ratio statistic, NR^{VOL}, is defined in terms of the first moment of trading volume,

$$NR^{VOL} \equiv VOL^N / (VOL^N + VOL^I)\,, \tag{5}$$

where VOL^N is trading volume due to noise estimated by mean trading volume on holidays and VOL^I is trading volume due to information estimated by mean trading volume on standard days less VOL^N (cf. equations (3) substituting VOL^N and VOL^I for V^N and V^I, respectively). As with the return-based NRs, NR^{VOL} is far from zero. The mean value of the NR^{VOL}'s displayed in column 7 is 0.481. That CCEF's are still briskly traded on foreign holidays suggests that noise matters.

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[8] Under certain distribution assumptions, NR is distributed Cauchy, which has neither a mean nor variance, and therefore NR can not be evaluated with a Wald test.

There are several attractive features of our estimation strategy. The holiday instrument is clearly exogenous and independent of noise in U.S. equity markets, and the factors that contribute to closed-end fund discounts do not vary at a daily frequency. Distortions due to trading location documented by Froot and Dabora (1999) and Hau (2001) do not impact our estimates, which are computed only for the U.S. market on holidays and standard days. The results on average are also not affected by possible distortions due to market illiquidity because any liquidity effects will be reflected in both the numerator and denominator of NR. Mean trading volume on standard days is statistically larger than holiday volume for 14 out of 54 funds at the 5% level of significance and only 8 out of 54 funds at the 1% level. The results in Table 2 rely on few modeling assumptions and thus largely avoid the joint hypothesis quagmire. The only two assumptions are the normality of returns -- to be relaxed in Section III.A -- and the substantial reduction in the flow of information on holidays.

The estimation strategy adopted in this paper rests on the latter identifying assumption. Even if a foreign country is closed for business, the fortunes of its companies might be buffeted by industry or global shocks emanating from outside the country. Industry shocks will not compromise our results because they will be dissipated over the portfolio of stocks constituting the CCEFs. Global shocks will in large part be accounted for by the market model, as the U.S. return will capture both the effects of the U.S. market and that part of the global shock correlated with the U.S. market. The residual global shock is accounted for in two ways. First, insofar as global shocks are large and infrequent, their impact is eliminated by trimming the tails of the distributions. Second, the impact of whatever fundamental global (or industry) information remains will be captured by the double mixture model of returns that allows for information leakage and that will be discussed in Section III.B.

III. Refinements

A. The Scaled-t Model

The variance estimates in Table 2 are based on the assumption that the returns are distributed normal. Given the well documented kurtosis in equity returns (e.g., Campbell, Lo, and MacKinlay, 1997, Table 1.1), variance estimates in the numerator and denominator of NR may be biased. A parsimonious way for addressing this problem is to assume that CCEF returns are distributed Scaled-t.[9] Under this more general assumption, the log likelihood depends on a variance parameter and a "degrees-of-freedom" parameter, $k > 2$. For values of $k \geq 30$, the Scaled-t distribution converges to the normal distribution used in the previous section. Since the Scaled-t distribution introduces higher order even moments, multiple solutions exist for V^N and V^I estimated by the method of moments.[10] To avoid this multiplicity problem, we evaluate the log likelihood function with a grid search over the parameter vector $\Psi = \{V^N, V^I, k\}$. A grid search also allows us to avoid using incorrect estimates defined by a local maximum.[11] The parameter space is determined by the admissible values in equation (3): $V^N = \{(1/m)V_H, (2/m)V_H, \ldots, V_H\}$, $V^I = \{(1/m)V_S, (2/m)V_S, \ldots, V_S\}$, $k = \{3, 4, \ldots, 30\}$, where the parameter

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[9] Focusing on unconditional volatility with the Scaled-t distribution requires fewer modeling choices than GARCH modeling of conditional volatility. Praetz (1972) links a fluctuating return variance to the Scaled-t distribution. The density function for the Scaled-t distribution for standard days is as follows,

$$f(R_t; V^N + V^I, k) = \left\{ \Gamma\left((k+1)/2\right) \Big/ \left(\Gamma\left(k/2\right) \left(\pi(k-2)(V^N + V^I) \right)^{\frac{1}{2}} \right) \right\} \left\{ \left[1 + (R_t^2 / ((k-2)(V^N + V^I))) \right]^{-\left((k+1)/2\right)} \right\}$$

where k is the degrees of freedom parameter and Γ(.) is the gamma function. This density characterizes holiday returns when $V^I = 0$.

[10] We thank Nicholas Kiefer for this observation that led us to maximum likelihood estimation.

[11] The possibility of settling at a local maximum or encountering other estimation problems with an ill-behaved likelihood function (Hamilton, 1994, p. 689) will be exacerbated in the double mixture model discussed in Section III.B.

determining the size of the increments is m = 400.[12] In order to ensure that the estimated NR is a lower bound, the grid search is performed twice with data that are and are not adjusted for the U.S. market; see Section 3 of the Appendix for further discussion. This two step procedure yields estimates of $\underline{V}^N$ and $(\bar{V}^N + \bar{V}^I)$ that determine NR defined in equation (4b).

Table 3 contains maximum likelihood estimates with the Scaled-t distribution. Estimates of k range from 3 to 6 with a mode of 3. These estimates are very far from the value of k = 30 that corresponds to a normality assumption and suggest the importance of accounting for fat-tails in the return distribution.[13] Relative to assuming normality in Table 2, we find that the mean value of NR falls slightly from 0.767 in Table 2 to 0.752 in Table 3. Our conclusion that CCEF returns are largely influenced by noise is insensitive to kurtosis.

B. Double Mixture Model

The above estimates have been based on the assumption that, on foreign holidays, the flow of information relevant for pricing CCEFs is totally eliminated. Nonetheless, it is not unreasonable to assume that -- despite our adjustments for industry and global shocks, exchange rates, and the U.S. market -- some information germane to CCEF returns is transmitted during a holiday. This subsection generalizes the previous models by allowing for the possibility of information "arrival" or "leakage" during holidays.

CCEF returns are modeled as a double mixture.[14] We assume that the probability of information "arrival" or "leakage" during holidays is non-negative and represented by a parameter, $0 \leq \lambda \leq 0.5$, the probability of information arrival. The importance of information (and thus the volatility due to information) for holiday returns is influenced by the value of λ. In

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[12] When we use a finer partition and set m = 1000 for several CCEFs, the point estimates are nearly identical in the sense that the estimate based on the finer partition is always within the "band" created by the coarser partition with m = 400. For example, if V^{N*} is the maximizing value of V^N for the coarser partition and defines a band, $\{((x^*-1)/400)\, V_H,\ (x^*/400)\, V_H\}$ for $x^* = \{1, 400\}$, and if V^{N**} is the maximizing value of V^N for the finer partition, then $((x^*-1)/400)\, V_H \leq V^{**} \leq (x^*/400)\, V_H$.

[13] Similar results for k are obtained by Aparicio and Estrada (2001) for daily aggregate return indices of 13 European equity markets.

[14] See Hamilton (1994, Section 22.3) and McLachlan and Peel (2000, Chapter 1) for overviews of mixture models. Roll (1988, Section IV.A) uses a formulation similar to equation (6).

this case, the holiday return in equation (1) is replaced by the following mixture of noise and information returns,

$$\begin{aligned} R_{H,t} &= R_t^N + \delta_t R_t^I \qquad\qquad \forall t \in \Im^H \text{ where } \Im^H = \{t : t \text{ is a holiday}\}, \\ &= (1-\delta_t) R_t^N + \delta_t (R_t^N + R_t^I), \end{aligned} \tag{6}$$

where the stochastic mixing weight, δ_t, equals 1 if information arrives on a given day, is 0 otherwise, and is distributed Bernoulli with mean λ. The second line in equation (6) highlights how δ_t is a mixing weight for noise and noise plus information. Allowing for information leakage during holidays does not affect the equation for standard days (equation (1) with d=S). The double mixture model combines the returns from holidays and standard days and, for each of these types of days, combines the returns from noise and news,

$$\begin{aligned} R_t &= \rho_t R_{H,t} + (1-\rho_t) R_{S,t} \qquad\qquad \forall t, \\ R_t &= \rho_t ((1-\delta_t) R_t^N + \delta_t (R_t^N + R_t^I)) + (1-\rho_t)(R_t^N + R_t^I). \end{aligned} \tag{7}$$

In general, analyzing this double mixture model would be very difficult. However, in our case, one of the mixing variables (ρ_t) is non-stochastic, and the estimation problem becomes considerably more tractable. The log likelihood function is evaluated with a grid search over the parameter vector $\Psi = \{V^N, V^I, k, \lambda\}$ for the same parameter space as used for the Scaled-t model for the first three parameters. In addition, we allow $\lambda = \{0.1, 0.2, \ldots, 0.5\}$, reflecting the possibility that there is some information flow during holidays. As with the Scaled-t model in Section III.A, the grid search is performed twice with data that are and are not adjusted for the U.S. market to ensure that the estimated NR is a lower bound; see Section 3 of the Appendix for further discussion. This two step procedure yields estimates of $\underline{V}^N$ and $(\overline{V}^N + \overline{V}^I)$ that determine NR defined in equation (4b).

Imposing the Scaled-t distributions for V^N and V^I is important for proper inferences to be drawn from the double mixture model. A possible difficulty with using equations (6) or (7) to quantify noise is that the estimated λ may be influenced by return kurtosis. If the underlying returns were assumed normal, a high value of λ could be due to information flow during holidays or to the kurtotic distribution of returns captured by a mixture of two normals (Press, 1967; McLachlan and Peel, 2000, Section 1.5). By assuming that V^N and V^I are distributed Scaled-t, we allow λ greater scope for capturing information flows.

Parameter estimates for the double mixture model presented in Table 4 confirm the importance of noise. All but four of the NRs are above 0.40. As expected, the role of noise is diminished in this model that allows some of the holiday return volatility to be accounted for by information. The mean NR of 0.673 is lower than the mean NR of 0.752 (Table 3) for the Scaled-t model without leakage.

We formally evaluate the null hypothesis that the NR is 0.10 (thus allowing some noise to affect the estimates) against the maximum likelihood values in Table 4. As indicated by the p-values in column 10, this null hypothesis is rejected at the 1% level for 49 of the 54 CCEFs and at the 5% level for all but one fund. Noise continues to be an important component of return volatility even when there is information leakage during holidays.

IV. Institutional Ownership

The results in Table 4 provide some evidence that the source of noise is not related to characteristics of the non-U.S. markets in which the CCEFs underlying assets are located. Several countries have more than one CCEF. For China, France, and Thailand, the NRs within a country are fairly close together. More dispersion exists for other countries. For example, The Emerging Mexico Fund and the Mexico Equity and Income CCEFs have NRs of 0.606 and 0.729, respectively; these NRs are much higher than the value of 0.381 for the Mexico Fund. Given this variety for Mexican and other CCEFs, country characteristics are unlikely to explain the variation in the NRs.

Noise thus appears to be driven by the trading environment in U.S. equity markets or the nature of CCEF investors. The role of the trading environment will be explored in Section V by drawing comparisons between U.S. and U.K. CCEFs. Here we investigate the role of institutional ownership, which is a salient investor characteristic that presumably identifies financially sophisticated and patient owners less susceptible to the "sentiments" that drive noise trading. We would thus expect the NRs for country c to be negatively related to the percentage of equity of a CCEF for country c held by institutions, $INST_c$.

This hypothesis is evaluated in terms of a linear regression of NR_c on $INST_c$, and the results are presented in Table 5. (Since NR_c appears as the dependent variable, this regression is not subject to a generated regressor problem.) Three NR_c's are used and are taken from column 2 of Tables 2, 3, and 4. $INST_c$ is taken from column 9 of Table 1 (see the Notes to Table 5 for the precise definition of institutional ownership and other details). As shown in panel A, noise is negatively associated with institutional ownership for each estimate of noise. This relation is statistically different from zero at conventional levels of significance for the two NR_c's based on the Scaled-t distribution in column 2; the elasticity of noise with respect to institutional ownership is about -0.17. This result is robust in two dimensions. First, the linear regression model is expanded to included quadratic and cubic terms. The elasticities (not reported) range between -0.07 and -0.20, though they tend to be estimated less precisely than those in Table 5. Second, panel B repeats this exercise but with a sample of CCEFs that are trimmed to reduce the potential effects of outliers, and the results are similar to those for the complete sample.

The associations reported in Table 5 surely do not necessarily imply a causal role for institutional ownership. A plausible alternative interpretation is that institutions prefer less volatile funds and that stock return volatility is positively correlated with the noise volatility estimated here. Gompers and Metrick (2001, Table IV) present a bit of indirect evidence against the alternative interpretation. They report that the relation between stock return volatility and institutional ownership is negative and statistically significant (at the 5% level) in only 13% of their cross-section regressions. Contrary to the alternative interpretation, 46% of their cross-section regressions yield a statistically significant positive relation between institutional ownership and stock return volatility. The alternative interpretation notwithstanding, the results in Table 5 provide some suggestive evidence of an interesting relation between noise and the "smart money."

V. The United Kingdom And The Securities Transaction Tax

Further insights about the importance of noise and its sensitivity to the trading environment can be obtained by repeating the above exercise on CCEFs traded in the United Kingdom (known there as investment trusts). Meaningful comparisons can be drawn because these two equity markets are widely viewed as the most developed in the world. Moreover, only the United Kingdom imposes a securities transaction tax.

Estimated NRs are presented in summary fashion in Table 6 for 32 U.K. CCEFs. (Returns and exchange rate data are taken from Datastream for the period January 1, 1980 to December 31, 2004; the percentage change in the FTSE 100 is used for the market return.) The NRs in columns 2 and 3 are based on the Normal and Scaled-t distributions, respectively, and on the assumption of no leakage of information during holidays. This latter assumption is relaxed in column 4. Two interesting results emerge. First, relative to the U.S. NRs, the average estimates for the United Kingdom are much lower: 0.410 vs. 0.767 for the Normal distribution without leakage, 0.560 vs. 0.752 for the Scaled-t distribution without leakage, and 0.316 vs. 0.673 for the Scaled-t distribution with leakage. Second, the numerous Japanese CCEFs provide further evidence that noise is unrelated to the characteristics of the non-U.K. market in which the underlying assets for the CCEF are located. For the 14 Japanese funds traded in the U.K., the NRs in column 4 range widely -- the standard deviation is 0.202 (relative to a mean of 0.306) and the high and low values are 0.675 and 0.100.

Given the depth and sophistication of the U.S. and U.K. equity markets, the first result concerning relatively less U.K. volatility seems puzzling. However, an important difference between these equity markets is that the United Kingdom has assessed a securities transactions tax since 1891 that is currently the highest among industrialized countries. During our sample period, this tax was assessed at the rate of 2.0% from 1980 to 1983, 1.0% from 1984 to 1985, and 0.5% from 1986 to 2004 on the value of paperless transactions. The impact is substantial, and the transaction tax accounts for about one-half of trading costs for U.K. equities.[15] If this transactions tax is effective in reducing noise trading, we would expect that, for those countries

[15] The data are from the Institute for Fiscal Studies (2007); all dates are for the beginning of the fiscal year. See Campbell and Froot (1995, Section 2.2, especially p. 129, and Table 4-1) and Summers and Summers (1989, Section 2) for detailed discussions of securities transaction taxes in the United Kingdom and other industrialized countries.

in which both U.K. and U.S. CCEFs invest, the NRs for the United Kingdom will be less than those for the United States. There are 11 countries for which CCEFs exist in both the United Kingdom and the United States. The NRs for these matched CCEFs are presented in Table 7, as well as the ratio of the U.K. to the U.S. noise ratios (RNR). Seven of the 11 RNRs are below 0.400. The results in Table 7 do not merely reflect lower relative volatility for U.K. equities (as the U.K. equity market is 19% more volatile than the U.S. market)[16] nor relatively higher capital gains taxes in the United Kingdom that would dampen noise trading.[17] For the most general model based on a Scaled-t distribution and the possibility of information leakage during holidays, the RNRs listed between columns 6 and 7 have a mean value of 0.578.[18]

The evidence presented in Table 7 should only be viewed as preliminary, as the roles of commissions and other frictions impeding noise traders need to be considered in a fuller empirical assessment of the securities transaction tax. Nonetheless, the results are of some interest, especially since they focus on the policy-relevant noise volatility. They offer some support for the "cautious case" for a securities transactions tax proposed by Summers and Summers (1989), and they verify a claim made long ago by Keynes (1936, pp. 159-160),

> That the sins [speculation] of the London Stock Exchange are less than those of Wall Street may be due, not so much to differences in national character, as to the fact that to the average Englishman Throgmorton

[16] The aggregate volatilities are computed as the coefficients of variation in the percentage change in the price indices of the S&P500 and FTSE100. These statistics are merely suggestive, as there can be a great deal of distance between the volatility for the overall equity market (reflecting, in part, the volatility of information about fundamentals) and the volatility due to noise.

[17] During the period 1980 to 2004, the marginal tax rate on short-term capital gains in the United Kingdom (Institute for Fiscal Studies, 2007; all dates are for the beginning of the fiscal year) varied between 30% (1980 to 1987) and 40% (1988 to 2004). The comparable U.S. rate (Office of Tax Policy Research, 2007) exceeded the U.K. rate from 1980 to 1987, was slightly below it in 1988, lower than it from 1989 to 1992, and then virtually equal to it from 1993 to 2004. Moreover, for a large class of U.K. investors, the marginal tax rate is zero because of an annual exemption of net capital gain income. The exemption was £3,000 in 1980 and has increased steadily to £8,200 in 2004.

[18] xx include if we include wtd RNRs
The MEAN RNR is the unweighted or weighted mean of the eleven RNRs in a given column. The weights are computed for a given country (e.g., China) as the sum of the weights for the CCEFs in that country for the U.K. and for the U.S. multiplied by 0.5. The country weights are the sum of the weights for the CCEFs for that country for a given trading area (i.e., the U.K. or the U.S). The CCEF weights are the average market value a given CCEF relative to the average market value for all CCEFs for a given trading area; the averages are computed over the entire sample period.

> Street is, compared with Wall Street to the average American, inaccessible and very expensive. The jobber's "turn", the high brokerage charges and the heavy transfer tax payable to the Exchequer, which attend dealings on the London Stock Exchange, sufficiently diminish the liquidity of the market (although the practice of fortnightly accounts operates the other way) to rule out a large proportion of the transactions characteristic of Wall Street. The introduction of a substantial Government transfer tax on all transactions might prove the most serviceable reform available, with a view to mitigating the predominance of speculation over enterprise in the United States.

The results in Table 7 also support a recent policy initiative by the Chinese finance ministry to curb stock market volatility and a putative bubble by raising the securities transaction tax from 0.1% to 0.3% (Bradsher, 2007).

VI. Discussion

Several papers that have assessed noise and its role in equity markets or that have studied CCEFs are reviewed in this section (though the discussion is not exhaustive). Our approach is related to the well-known study of French and Roll (1986) that exploits a very interesting natural experiment.[19] They compare equity volatilities between trading and non-trading days and find that the hourly return variance during trading days is 71.8 times greater than over weekends (Friday close to Monday close; p. 223). As the authors note, this huge number can be traced to the absence of noise traders or a reduced flow of fundamental information on non-trading days. They use the closure of the New York Stock Exchange on Wednesdays during the second half of 1968 to differentiate between the noise and information explanations. The two-day return volatility from Tuesdays to Thursdays during the Wednesday exchange holidays is 1.15 times greater than the one-day volatility (the latter computed from January 1963 to December 1982). If return volatilities were unaffected by the trading process, this ratio should be 2.00. Clearly, trading raises volatility. While extremely informative, this particular natural experiment is unable to differentiate between two competing and very different hypotheses -- whether the trading process reveals information (which French and Roll favor based on autocorrelation tests) or generates noise.[20] By contrast, our focus on non-U.S. holidays generates a direct estimate of the impact of noise. Information may not be completely turned-off during non-U.S. holidays, and the double mixture model allows for this information leakage.

An alternative means for controlling for fundamentals focuses on the return difference between "Siamese twin" companies, defined as companies whose equity has different trading

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[19] Our approach is also related to Kamstra, Kramer, and Levi (2000, 2003) and Hirshleifer and Shumway (2003), who use the natural experiments of the exogenous shifts to and from daylight saving time and in the amount of daylight and sunshine to identify behavior in equity markets that deviates from the efficient markets benchmark.

[20] French and Roll discriminate between the private information and noise hypotheses by examining weekly return volatility for weeks with and without an exchange holiday. Consistent with the noise hypothesis, weekly return volatility is lower during weeks with exchange holidays, but this difference is not statistically significant. Using the natural experiment of occasional Saturday trading on the Tokyo Stock Exchange, Barclay, Litzenberger, and Warner (1990) find that, for weeks with Saturday trading, the weekly return volatility is unchanged relative to weeks without Saturday trading; this result is consistent with the private information hypothesis. By contrast, the weekly volume for weeks with Saturday trading is higher, which can be interpreted as consistent with the noise hypothesis (per French and Roll) or the private information hypothesis (as argued by Barclay, Litzenberger, and Warner).

and ownership habitats but whose charters specify the division of a common pool of cash flow. Exploiting the key identifying assumption that fundamental information is the same for twin equities, Scruggs (2007) draws-out the implications for noise.[21] His estimates imply relatively low NRs, ranging from 11% to 19% for daily returns.[22] While Siamese twins provide a very clever way to control for fundamentals, location frictions -- such as securities taxes -- that differ between exchanges might cause returns to deviate from fundamentals independent of noise.[23] This problem can be avoided by focusing on ADR's of the twins traded in New York. But a second problem arises in estimating noise from the difference in twin returns, as we must assume that the individual noise components for the twins are uncorrelated (see fn. 22). Given that noise trading is likely to be positively correlated among twin stocks, estimates of NR will be downward biased.

In an interesting paper, Pontiff (1997) documents that the volatility of closed-end fund prices is excessive, being 64% greater than the volatility of fundamentals as represented by NAV's. While there is much useful information in these results, several issues arise linking these volatilities to noise. First, Abraham, Elan, and Marcus (1993) show that the premium in closed-end equity and bond funds have very similar correlations with the market. This coincidence of results may reflect time varying risk premium affecting both equity and bond markets, and it raises questions about interpreting variations in the premium on closed-end equity funds as

[21] Froot and Dabora (1999) also examine twin equities but focus on the relations between trading location and equity returns.

[22] The NR statistic is computed from Scrugg's estimates as follows. Writing the equity returns for each twin as $R_t = R_t^N + R_t^I$ (as in our equation (1) and a simplified version of Scruggs' equation (1)), denoting the twins by a prime or double prime superscript, and taking the difference of the twin returns (θ), we obtain the following relation, $\theta \equiv R_t^{'} - R_t^{''} = R_t^{N'} + R_t^{I'} - R_t^{N''} - R_t^{I''} = R_t^{N'} - R_t^{N''}$, where the R_t^I variables are eliminated by the key identifying assumption that fundamental information is the same for each twin equity ($R_t^{I'} = R_t^{I''}$). The numerator of the NR statistic (equation (4a)) requires a measure of noise, which we define as an average of the noise variances for each twin. Applying the variance operator to the above expression for θ, we obtain the following relation, $(V[R_t^{N'}] + V[R_t^{N''}])/2 = (V[\theta] + 2 * C[R_t^{N'}, R_t^{N''}])/2$, where $C[R_t^{N'}, R_t^{N''}]$ is an unobservable covariance and assumed to be equal to zero for purposes of the computation of NR reported in the text. The denominator of the NR statistic is defined as an average of the return variance for each twin, $(V[R_t^{'}] + V[R_t^{''}])/2$. Values for $V[R_t^{'}]$, $V[R_t^{''}]$, and θ are taken from Scruggs (2007, Table 1).

[23] This concern is noted by Scruggs (2007, pp. 85-86). Other tests reported in his paper indicate that noise volatility, estimated by conditional volatilities, is large.

reflecting noise. Second, if noise affects equity prices in U.S. markets, then presumably noise affects the non-U.S. market as well. Thus, NAV is not a fully satisfactory measure of fundamental value. Third, interactions between fundamentals and noise compromise the use of premia to estimate noise. Consider the following accounting identity relating the price of the CCEF (P) to fundamental value (F, measured by the NAV) and the premium, $\Phi[N, F]$, where the latter depends on both noise (N) and fundamentals and time subscripts have been omitted for notational convenience,

$$P = \Phi[N, F]\, F . \tag{8}$$

Taking logs, totally differentiating with respect to P, N, and F, defining returns (R_X) as the percentage change in X = {P, F, $\Phi[N, F]$ }, computing variances, and rearranging terms, we obtain the following expression,

$$(V[R_P] - V[R_F]) / V[R_F] = \begin{pmatrix} V[R_{\Phi_N}] + \\ V[R_{\Phi_F}] + 2C[R_{\Phi_F}, R_F] + 2C[R_{\Phi_F}, R_{\Phi_N}] + \\ 2C[R_{\Phi_N}, R_F] \end{pmatrix} / V[R_F] \ . \tag{9}$$

where C[.] is the covariance operator and Φ_N and Φ_F refer to partial derivatives of the premium with respect to N and F, respectively. The left-side of equation (9) equals the 64% figure reported by Pontiff. Equation (9) is a useful estimate of the variance due to noise, defined by $V[R_{\Phi_N}]$, provided the remaining four terms in the large parentheses are zero. The results of Campbell and Kyle (1993) and Klibanoff, Lamont, and Wizman (1998) suggest that fundamentals affect the premium, and hence R_{Φ_F} is time-varying. Even if R_{Φ_F} is constant and hence the three terms in the second line in the large parentheses equal zero, the $C[R_{\Phi_N}, R_F]$ term disrupts accurately measuring noise by the variances in market prices and NAV's. The approach presented in the present study avoids these issues in quantifying noise.

Several studies have documented excess volatility in equities relative to various measures of fundamentals. Roll (1984, 1988) finds that the volatilities of orange juice futures and monthly equity returns, respectively, are poorly explained by ex-post fundamentals. In the case of equity returns, about 20% of the variation in daily returns and about 35% of the variation in monthly returns are explained by fundamentals. Cutler, Poterba, and Summers (1989, Table 1) report that contemporaneous macro variables capturing important news items explain about 20% of the variation in monthly aggregate returns (for the period 1926-1985) and about 6% of the variation in annual aggregate returns (1871-1986). Equity prices are much more volatile than the present value of the dividend stream computed with a constant discount rate (LeRoy and Porter (1981), Shiller (1981), and West (1988)), and equity returns are much more volatile than consumption growth (Campbell, 2003). These excess volatility studies suggest a natural and quantitatively important role for noise in filling the wide gap between equity prices (or returns) and fundamentals.

VII. Conclusions

This paper uses the natural experiment provided by foreign holidays to quantify the impact of noise on equity returns in highly developed equity markets. Our baseline results highlight that there is a great deal of noise in the returns of CCEFs in the United States. These results prove robust to kurtosis and the possibility of information leakage during holidays. Noise is negatively related to institutional ownership of U.S. CCEFs and is much less important for U.K. CCEFs. The lower levels of noise for matched U.K. and U.S. CCEFs suggests that the securities tax imposed in the United Kingdom may be effective in reducing equity market noise.

Noise disrupts financial markets, raises the cost of capital, impedes arbitrageurs, complicates stabilization policies, and is central to the issue of stock market efficiency. Merton (1987b, p. 93) states the critical test:

> “The theory [the rational market hypothesis] is not, however, a tautology. It is not consistent with models or empirical facts that imply that either stock prices depend in an important way on factors other than the fundamentals underlying future cash flows and discount rates, …

A quantitatively important role for noise documented in this study thus challenges the efficient market model and points toward the behavioral finance alternative as a framework model for understanding equity markets.

References

Abraham, Abraham, Elan, Don, and Marcus, Alan J., "Does Sentiment Explain Closed-End Fund Discounts?: Evidence from Bond Funds," *Financial Review* 28 (November 1993), 607-616.

Aparicio, Felipe M., and Estrada, Javier, "Empirical Distributions of Stock Returns: European Securities Markets 1990-95," *The European Journal of Finance* 7 (2001), 1-21.

Arnott, Robert D., Hsu, Jason, and Moore, Philip, "Fundamental Indexation," *Financial Analysts Journal* 61 (March/April 2005), 83-99.

Baker, Malcolm, Stein, Jeremy C., and Wurgler, Jeffrey, "When Does The Market Matter? Stock Prices and the Investment of Equity-Dependent Firms," *Quarterly Journal of Economics* 118 (August 2003), 969-1005.

Ball, Turan G., Cakici, Nusret, Yan, Xuemin (Sterling), and Zhang, Zhe, "Does Idiosyncratic Risk Really Matter?," *Journal of Finance* 60 (April 2005), 905-930.

Barclay, Michael J., Litzenberger, Robert H., and Warner, Jerold B., "Private Information, Trading Volume, And Stock Return Variances," *Review Of Financial Studies* 3 (1990), 233-254.

Black, Fischer, "Noise," *Journal of Finance* 41 (July 1986), 529-543. Reprinted as Chapter 1 in Richard H. Thaler (ed.), *Advances In Behavioral Finance* (New York: Russell Sage Foundation, 1993), 3-22.

Bodurtha, James N., Jr., Kim, Dong-Soon, Lee, Charles M.C., "Closed-end Country Funds and U.S. Market Sentiment," *The Review of Financial Studies* 8 (Fall 1995), 879-918.

Bradsher, Keith, "China Triples Tax on Stock Trades," *New York Times* (May 30, 2007).

Brown, David P., and Ferreira, Miguel A., "Information in the Idiosyncratic Volatility of Small Firms," University of Wisconsin (January 2005).

Campbell, John Y., "Consumption-Based Asset Pricing," in George M. Constantinides, Milton Harris, and René M. Stulz (eds.), *Handbook of the Economics of Finance*, Volume 1B (Amsterdam: North-Holland, 2003), 804-887.

Campbell, John Y., and Froot, Kenneth A., "Securities Transaction Taxes: What about International Experiences and Migrating Markets?," in Suzanne Hammond (ed.), *Securities Transaction Taxes: False Hopes and Unintended Consequences* (Chicago: Catalyst Institute, 1995), 110-142.

Campbell, John Y., and Kyle, Albert S., "Smart Money, Noise Trading and Stock Price Behaviour," *Review of Economic Studies* 60 (January 1993), 1-34

Campbell, John Y., Lo, Andrew W., and MacKinlay, A. Craig, *The Econometrics of Financial Markets* (Princeton: Princeton University Press, 1997).

Chen, Nai-Fu, Kan, Raymond, and Miller, Merton H., “Are the Discounts on Closed-End Funds a Sentiment Index?” and “Rejoinder,” *Journal of Finance* 48 (June 1993), 795-800 and 809-810.

Chirinko, Robert S., and Schaller, Huntley, "Business Fixed Investment and 'Bubbles': The Japanese Case," *American Economic Review* 91 (June 2001), 663-680.

Chirinko, Robert S., and Schaller, Huntley, "Fundamentals, Misvaluation, and Investment: The Real Story," University of Illinois at Chicago (June 2008).

Chopra, Navin, Lee, Charles M.C. Shleifer, Andrei, and Thaler, Richard H., “Yes, Discounts on Closed-End Funds Are a Sentiment Index” and “Summing Up,” *Journal of Finance* 48 (June 1993), 801-808 and 811-812.

Cutler, David M., Poterba, James M., and Summers, Lawrence H., "What Moves Stock Prices?," *The Journal of Portfolio Management* (Spring 1989), 4-12. Reprinted as Chapter 5 in Richard H. Thaler (ed.), *Advances In Behavioral Finance* (New York: Russell Sage Foundation, 1993), 133-151.

De Long, J. Bradford, Shleifer, Andrei, Summers, Lawrence H., and Waldmann, Robert J., "Noise Trader Risk in Financial Markets," *Journal of Political Economy* 98 (August 1990), 703-738. Reprinted as Chapter 2 in Richard H. Thaler (ed.), *Advances In Behavioral Finance* (New York: Russell Sage Foundation, 1993), 23-58.

Easley, David, Hvidkjaer, Soeren, and O’Hara, Maureen, “Is Information Risk a Determinant of Asset Returns?,” *Journal of Finance* 57 (October 2002), 2185-2221.

Fama, Eugene, “Efficient Capital Markets: A Review of Theory and Empirical Work,” *Journal of Finance* 25 (1970), 383-417.

Fama, Eugene, “Efficient Capital Markets II,” *Journal of Finance* 46 (1991), 1575-1617.

Fischer, Stanley, and Merton, Robert C., "Macroeconomics and Finance: The Role of the Stock Market," in Karl Brunner and Allan H. Meltzer (eds.), *Essays on Macroeconomic Implications of Financial and Labor Markets and Political Processes*, Carnegie-Rochester Conference Series on Public Policy 21 (Autumn 1984), 57-108.

French, Kenneth R. and Roll, Richard, "Stock Return Variances: The Arrival of Information and the Reaction of Traders," *Journal of Financial Economics* 17 (1986), 5-26. Reprinted as Chapter 8 in Richard H. Thaler (ed.), *Advances In Behavioral Finance* (New York: Russell Sage Foundation, 1993), 219-245.

Froot, Kenneth A., and Dabora, Emil M., "How Are Stock Prices Affected by the Location of Trade?," *Journal of Financial Economics* 53 (August 1999), 189-216.

Gilchrist, Simon, Himmelberg, Charles, Huberman, Gur, “Do Stock Price Bubbles Influence Corporate Investment?” *Journal of Monetary Economics* 52 (May 2005), 805-827.

Gompers, Paul A., and Metrick, Andrew, "Institutional Investors and Equity Prices," *Quarterly Journal of Economics* 116 (February 2001), 229-260.

Goyal, Amit, and Santa-Clara, Pedro, “Idiosyncratic Risk Matters!,” *Journal of Finance* 58 (June 2003), 975-1008.

Hamilton, James D., *Time Series Analysis* (Princeton: Princeton University Press, 1994).

Hardouvelis, G. La Porta, R., Wizman, T.A., "What Moves the Discount on Country Equity Funds?," in Jeffrey Frankel (ed.) *The Internationalization of Equity Markets* (Chicago: University of Chicago Press (for the NBER), 1994), 345-397.

Hau, Harald, “Location Matters: An Examination of Trading Profits,” *Journal of Finance* (October 2001), 1959-1983.

Hirshleifer, David, and Shumway, Tyler, “Good Day Sunshine: Stock Returns and the Weather,” *Journal of Finance* 58 (June 2003), 1009-1032.

Institute for Fiscal Studies, Website (2007), http://www.ifs.org.uk/ff/indextax.php.

Issing, Otmar, “Should central banks burst bubbles?,” *The Wall Street Journal* (February 18, 2004).

Kamstra, Mark J., Kramer, Lisa A., and Levi, Maurice D., “Losing Sleep At The Market: The Daylight Saving Anomaly,” *American Economic Review* 90 (September 2000), 1005-1011.

Kamstra, Mark J., Kramer, Lisa A., and Levi, Maurice D., “Winter Blues: A SAD Stock Market Cycle,” *American Economic Review* 93 (March 2003), 324-343.

Keynes, John Maynard, *The General Theory of Employment, Interest, and Money* (New York: Harcourt Brace, 1936). Reprinted in 1964, First Harbinger Edition.

Klibanoff, Peter, Lamont, Owen A., and Wizman, Thierry A., "Investor Reaction to Salient News in Closed-End Country Funds," *Journal of Finance* 53 (April 1998), 673-699.

Lee, Charles M.C., Shleifer, Andrei, and Thaler, Richard H., "Investor Sentiment and the Closed-End Fund Puzzle," *Journal of Finance* 46 (March 1991), 75-109. Reprinted as Chapter 3 in Richard H. Thaler (ed.), *Advances In Behavioral Finance* (New York: Russell Sage Foundation, 1993), 59-106.

Lei, Vivian, Noussair, Charles N., and Plott, Charles R., "Nonspeculative Bubbles in Experimental Asset Markets: Lack of Common Knowledge vs. Actual Irrationality," *Econometrica* 69 (July 2001), 831-859.

LeRoy, Stephen F., "Efficient Capital Markets and Martingales," *Journal of Economic Literature* 27 (December 1989), 1583-1621.

LeRoy, Stephen F., and Porter, Richard D., "The Present-Value Relation: Tests Based on Implied Variance Bounds," *Econometrica* 49 (May 1981), 555-574.

Levy, Haim, "Equilibrium in an Imperfect Market: A Constraint on the Number of Securities in the Portfolio," *American Economic Review* 68 (September 1978), 643-658.

Malkiel, Burton G., "The Efficient Markets Hypothesis and Its Critics," *Journal of Economic Perspectives* 17 (Winter 2003), 59-82.

Malkiel, Burton G., and Xu, Yexiao, "Idiosyncratic Risk and Security Returns," University of Texas at Dallas (November 2002).

McLachlan, Geoffrey J., and Peel, David, *Finite Mixture Models* (New York: John Wiley & Sons, 2000).

Merton, Robert C., "A Simple Model of Capital Market Equilibrium with Incomplete Information," *Journal of Finance* 42 (July 1987a), 483-510.

Merton, Robert C., "On the Current State of the Stock Market Rationality Hypothesis," in Rudiger Dornbusch, Stanley Fischer, and John Bossons (eds.), *Macroeconomics and Finance: Essays in Honor of Franco Modigliani* (Cambridge: MIT Press, 1987b), 93-124.

Office of Tax Policy Research, University of Michigan, Website (2007), http://www.bus.umich.edu/otpr/otpr/.

Polk, Christopher, and Sapienza, Paola, "The Stock Market and Corporate Investment: a Test of Catering Theory," *Review of Financial Studies* (forthcoming).

Pontiff, Jeffrey, "Excess Volatility and Closed-End Funds," *American Economic Review* 87 (March 1997), 155-169.

Praetz, Peter D., "The Distribution of Share Price Changes," *Journal of Business* 45 (1972), 49-55.

Press, S. James, "A Compound Events Model for Security Prices," *Journal of Business* 40 (July 1967), 317-335.

Roll, Richard, "Orange Juice and Weather," *American Economic Review* 74 (December 1984), 861-880.

Roll, Richard, "R^2," *Journal of Finance* 43 (July 1988), 541-566.

Scruggs, John T., “Noise Trader Risk: Evidence from the Siamese Twins,” *Journal of Financial Markets* 10 (February 2007), 76-105.

Shiller, Robert J., "Do Stock Prices Move Too Much To Be Explained by Subsequent Changes in Dividends?," *American Economic Review* 71 (June 1981), 421-436.

Shiller, Robert J., *Market Volatility* (Cambridge: MIT Press, 1990).

Shiller, Robert J., “From Efficient Markets Theory to Behavioral Finance,” *Journal of Economic Perspectives* 17 (Winter 2003), 83-104.

Shleifer, Andrei, *Inefficient Markets: An Introduction to Behavioral Finance* (Oxford: Oxford University Press, 2000).

Shleifer, Andrei, and Summers, Lawrence H., “The Noise Trader Approach to Finance,” *Journal of Economic Perspectives* 4 (Spring 1990), 19-33.

Siegel, Jeremy J., "The 'Noisy Markets' Hypothesis," *The Wall Street Journal* (June 14, 2006), A14.

Summers, Lawrence H., and Summers, Victoria, P., “When Financial Markets Work Too Well: A Cautious Case for a Securities Transactions Tax,” *Journal of Financial Services Research* 3 (1989), 261-286.

Thaler, Richard H. (ed.), *Advances In Behavioral Finance* (New York: Russell Sage Foundation, 1993).

Thaler, Richard H. (ed.), *Advances In Behavioral Finance, Volume II* (Princeton: Princeton University Press, 2005).

Weaver, Robert S., *International holidays: 204 countries from 1994 through 2015* (Jefferson, North Carolina: McFarland Press, 1995).

West, Kenneth D., “Dividend Innovations and Stock Price Volatility,” *Econometrica* 56 (January 1988), 37-62.

Appendix: The Double Mixture Model:

Derivation Of The Likelihood Function And Likelihood Ratio Test

This appendix derives the likelihood function and likelihood ratio test statistic for the double mixture model of returns. This is the most general of the three models estimated in this paper, and the other two models are derived as special cases under suitable parametric restrictions. We begin with the model of returns.

1. Returns

Returns are defined as a mixture of holiday and standard day returns,

$$R_t = \rho_t R_{H,t} + (1-\rho_t) R_{S,t} \qquad \forall t, \tag{A1}$$

where R_t, $R_{H,t}$, and $R_{S,t}$ are returns for the full sample, holidays, and standard days (non-holidays), respectively, and ρ_t is the mixing variable equal to one for holidays and zero for standard days.

Holiday returns are defined as a mixture of noise (R_t^N) and information (R_t^I) returns,

$$R_{H,t} = R_t^N + \delta_t R_t^I \qquad \forall t \in \Im^H \text{ where } \Im^H = \{t : t = \text{holiday}\}, \tag{A2a}$$

$$R_t^N \sim \text{Scaled} - t(0, V^N, k), \tag{A2b}$$

$$R_t^I \sim \text{Scaled} - t(0, V^I, k), \tag{A2c}$$

$$\delta_t \sim \text{Be}(\lambda), \tag{A2d}$$

where δ_t equals 1 if information arrives at date t and 0 otherwise. This mixing variable is distributed Bernoulli with mean λ. The returns due to noise and information are distributed Scaled-t with zero location parameter, a dispersion parameter equal to V^N or V^I, and k degrees of freedom. Assuming independence between noise returns, information returns, and the arrival of information, we can express the holiday return variance (V_H) as follows,

$$V_H = V^N + \lambda V^I . \qquad \text{(A3)}$$

Returns on standard days are defined as the sum of both noise and information,

$$R_{S,t} = R_t^N + R_t^I \qquad \forall t \in \Im^S \text{ where } \Im^S = \{t : t = \text{standard}\} . \qquad \text{(A4)}$$

The standard day variance (V_S) can be expressed as follows,

$$V_S = V^N + V^I . \qquad \text{(A5)}$$

2. Densities And The Log Likelihood

In our double mixture model, CCEF returns are characterized by a parameter vector, Ψ, containing four unknowns,

$$\Psi = \{V^N, V^I, k, \lambda\} . \qquad \text{(A6)}$$

Based on equation (A1), the density of CCEF returns ($f(R_t;\Psi)$) is a mixture of the densities of holiday and standard day returns,

$$f(R_t;\Psi) = \rho_t f(R_{H,t};\Psi) + (1-\rho_t) f(R_{S,t};\Psi) \qquad \forall t . \qquad \text{(A7)}$$

Holiday returns are also defined as a mixture of returns. With the definition of returns given in equation (A2), the joint density for returns and information leakage is given by the sum of the conditional densities,

$$f(R_{H,t};\Psi) = f(R_t \mid \delta_t = 0) h(\delta_t = 0) + f(R_t \mid \delta_t = 1) h(\delta_t = 1) \qquad \forall t \in \Im^H , \qquad \text{(A8)}$$

where $h(\delta_t)$ is the unconditional density for δ_t. Given the distribution assumptions in equations (A2), equation (A8) becomes,

$$f(R_{H,t};\Psi) = \lambda f(R_t;V^N + V^I,k) + (1-\lambda) f(R_t;V^N,k) \qquad \forall t \in \Im^H. \tag{A9}$$

On standard days, returns are influenced by both noise and information,

$$f(R_{S,t};\Psi) = f(R_t;V^N + V^I,k) \qquad \forall t \in \Im^S. \tag{A10}$$

Combining equations (A7), (A9), and (A10), we have the following double mixture density,

$$\begin{aligned} f(R_t;\Psi) = {} & \rho_t[\lambda f(R_t;V^N + V^I,k) + (1-\lambda) f(R_t;V^N,k)] \qquad \forall t. \\ & + (1-\rho_t) f(R_t;V^N + V^I,k) \end{aligned} \tag{A11}$$

We assume that returns are distributed Scaled-t defined by the following density function,

$$f(R_t;V^N,V^I,k) = \frac{\Gamma\left(\frac{k+1}{2}\right)}{\Gamma\left(\frac{k}{2}\right)\sqrt{\pi(k-2)\left(V^{\#}\right)}}\left[1 + \frac{R_t^2}{(k-2)\left(V^{\#}\right)}\right]^{-(k+1)/2} \qquad \forall t, \tag{A12}$$

$$V^{\#} = \{V^N, V^N + V^I\},$$

where k is the degrees of freedom parameter and $\Gamma(.)$ is the gamma function.

The four unknown parameters in Ψ are estimated using maximum likelihood. The log likelihood function is written as follows,

$$\ln L(\Psi) = \sum_{t=1}^{T} \ln f(R_t;\Psi), \tag{A13}$$

where $f(R_t; \Psi)$ is defined in equation (A11). Since the holiday mixing variable is non-stochastic, we can separate the log likelihood function into two components,

$$\begin{aligned}\ln L(\Psi) = &\sum_{t=1}^{T} \rho_t \ln[\lambda f(R_t; V^N + V^I, k) + (1-\lambda) f(R_t; V^N, k)] \\ &+ \sum_{t=1}^{T} (1-\rho_t) \ln f(R_t; V^N + V^I, k),\end{aligned} \qquad \text{(A14)}$$

With the densities given in equation (A11), the log likelihood function can be maximized with respect to the elements in Ψ (equation (A6)). Note that the first summation is based on holiday returns and the second summation is based on standard day returns.

3. The Empirical Results

All of the empirical results in this paper are based on the log likelihood in equation (A14). The most general results, presented in Table 4 and column 4 of Table 6, are based on estimates of all four parameters in Ψ. The results in Table 3 and column 3 of Table 6 are based on the assumption that there is no leakage of information during holidays; hence, $\Psi = (V^N, V^I, k, \lambda=0)$. In this case, for a given k, the first summation in equation (A14) depends only on V^N. Given this estimate of V^N, the second summation determines V^I. The results in Table 2 and column 2 of Table 6 are based on the assumption that returns are distributed normal. The Scaled-t distribution converges to the normal and provides a good approximation to the normal when $k \geq 30$; hence, $\Psi = (V^N, V^I, k=30, \lambda=0)$. In this case, the first summation in equation (A14) depends only on V^N, the second summation depends only on $(V^N + V^I)$, which effectively determines V^I given an estimate of V^N from the first summation. Thus, we obtain the well known equality between maximum likelihood and method of moments estimates of the variances when returns are normal (cf. equations (3) in the text).

The holiday and standard day return data are subjected to different adjustments for market-wide factors. The different adjustments are made to ensure that the reported NRs are lower bounds relative to the true NRs. For the holiday data, we remove the market effect by estimating for each country a market model of the exchange-rate-adjusted CCEF return against a constant and the return on the S&P 500. The residuals from this regression define the mean zero return series, R_t, for holidays. Note that this estimate is a lower bound for V_N, as that part of noise correlated with the market is removed by the market model. Relying on a similar adjustment with the market model for returns on standard days would also generate a lower bound estimate of the variances. However, since one of these estimated variances enter the denominator of the NR statistic, the estimate of NR would tend to be upwardly biased. To avoid this problem and ensure that the reported estimate of NR is a lower bound, we do not use the market model to remove the market effect for standard day returns. The CCEF returns during standard days are only adjusted for exchange rates. As shown in equation (4b), these different adjustments for market effects for the holiday and standard day data (hereafter referred to as the differential adjustment) yield an estimated NR that is a lower bound relative to the true NR.

The differential adjustment forces a reconsideration of our estimation strategy for each of the three models discussed in the first paragraph of this Section. The key issue is that the differential adjustment results in a difference in the nature of the returns data entering the first and second summations in equation (A14):

1) Baseline Model (Table 2 and column 2 of Table 6): As discussed at the beginning of this Section of the Appendix and the text below equation (A14), V^N is estimated from the holiday data in the first summation, while V^I is effectively estimated from the standard day data in the second summation. (Details of the computation of V^I are discussed Section 5 of the Appendix.) Given this separation, the differential adjustment does not lead to a problem with the estimation of V^N, V^I, and NR.

2) Scaled-t Model (Table 3 and column 3 of Table 6): As discussed at the beginning of this Section of the Appendix and the text below equation (A14), estimates of V^N and V^I are conditioned on k. Since k enters both summations, it is being estimated with data affected by the differential adjustment. To avoid this problem, we estimate the components of NR in two steps. The first step estimates the parameter vector $\Psi(1) = (V^N(1), V^I(1), k(1), \lambda=0)$ with returns data for both holidays and standard days that are adjusted for market effects; these estimates are labeled "excess" in columns 3, 5, and 7 in Table 3. The second step estimates the parameter vector $\Psi(2) = (V^N(2), V^I(2), k(2), \lambda=0)$ with returns data for both holidays and standard days that are not adjusted for market

effects; these estimates are labeled "raw" in columns 4, 6, and 8 in Table 3. The NR statistic is computed as NR = $V^N(1) / (V^N(2) + V^I(2))$. The estimates of k(1) and k(2) are very similar; cf. columns 7 and 8 in Table 3.

3) Double Mixture Model (Table 4 and column 4 of Table 6): The estimation issue related to the differential adjustment is similar to the one raised with the Scaled-t model. As discussed at the beginning of this Section of the Appendix and the text below equation (A14), estimates of V^N and V^I are conditioned on k and λ. Since V^N and V^I enter both summations, they are estimated with data affected by the differential adjustment. To avoid this problem, we estimate the components of NR in two steps. The first step estimates the parameter vector $\Psi(1) = (V^N(1), V^I(1), k(1), \lambda(1))$ with returns data for both holidays and standard days that are adjusted for market effects; these estimates are labeled "excess" in columns 3, 5, and 7 in Table 4. The second step estimates the parameter vector $\Psi(2) = (V^N(2), V^I(2), k(2), \lambda(2)=0)$ with returns data for both holidays and standard days that are not adjusted for market effects; these estimates are labeled "raw" in columns 4, 6, and 8 in Table 4. Since λ(2) does not affect standard day returns and Ψ(2) will be used to estimate (V^N+V^I) for standard days, λ(2) is set to zero in the second step. The NR statistic is computed as NR = $V^N(1) / (V^N(2) + V^I(2))$. The estimates of k(1) and k(2) are very similar; cf. columns 7 and 8 in Table 4. Only the estimate of λ(1) on holidays is relevant to the estimate of NR, and it is the estimate reported in column 9 of Table 4.

4. Likelihood Ratio Test

The noise ratio statistic, NR = $V^N / (V^N+V^I)$, is a ratio of two random variables. Under certain conditions, this ratio is distributed Cauchy, which has neither a mean nor variance, and therefore can not be evaluated directly for hypothesis testing. However, since the parameters are estimated by maximum likelihood, we can test certain hypothesis about NR by comparing the likelihoods associated with unrestricted and restricted models. Restrictions are imposed by relating V^N to V^I such that NR equals a particular value. For example, suppose we want to test a NR equal to Φ. In light of the estimation issues posed by the differential adjustment, we restrict $V^N = \Phi\,(V^N(2) + V^I(2))$.

To test the validity of these restrictions, we use a likelihood ratio test. The maximum likelihood estimates of the unrestricted (Ψ^U) and restricted (Ψ^R) parameter vector are as follows,

$$\Psi^U = \{\hat{V}^N, \hat{V}^I, \hat{k}, \hat{\lambda}\}, \qquad \text{(A15a)}$$

$$\Psi^R = \{V^N = \Phi(\tilde{V}^N + \tilde{V}^I), \tilde{k}, \tilde{\lambda}\}, \qquad \text{(A15b)}$$

where the ^ and ~ superscripts indicate estimated values for the unrestricted and restricted models, respectively. If the restriction is valid, the values of the unrestricted and restricted log likelihood functions should not be too far apart as assessed by the likelihood ratio test statistic, LR(Φ),

$$LR(\Phi) = -2\{\ln L(\Psi^R(\Phi)) - \ln L(\Psi^U)\}, \qquad \text{(A16)}$$

where $LR(\Phi) \sim \chi^2_r$ and r is the number of restrictions (one in this case). The p-values associated with equation (A16) for the null hypothesis that NR = Φ = 0.10 are presented in column 7 of Table 2, column 9 of Table 3, and column 10 of Table 4.

5. *Computation of $\bar{V}^I$ in Section II and Table 2*

In Section II, the numerator ($\underline{V}^N$) and the denominator ($\bar{V}^N + \bar{V}^I$) of the noise ratio statistic in equation (4b) are estimated directly from the data. The estimate of $\bar{V}^I$ in Table 2 is a transformation of these direct estimates, and it remains an open question whether this transformed estimates is an upper bound for V^I. Consider the following identity for the true value of V^I, V^{I*},

$$V^{I*} = V^{I*} + V^{N*} - V^{N*}, \qquad \text{(A17)}$$

where V^{N*} is the true value of V^{N}. The true values are replaced by upper or lower bound estimates

$$\begin{aligned} V^{I*} &= \overline{V}^{I} - \overline{\xi}^{I} + \overline{V}^{N} - \overline{\xi}^{N} - (\underline{V}^{N} + \underline{\xi}^{N}), \\ &= \overline{V}^{I} + \overline{V}^{N} - \underline{V}^{N} - (\overline{\xi}^{I} + \overline{\xi}^{N} + \underline{\xi}^{N}), \end{aligned} \quad \text{(A18)}$$

where the ξ's represent the discrepancies between the upper or lower bound estimates and true values and are all positive. The first three terms on the right side of equation (A18) represents the transformation that estimates V^{I}, and this transformation generates an upper bound estimate for V^{I}.

Table 1: List Of U.S. Country Closed-End Funds

Country (1)	Country Closed-End Fund (CCEF) (2)	Ticker (3)	CUSIP (4)	Start (5)	End (6)	T_H (7)	T_S (8)	INST (9)
Argentina	Argentina Fund	AF	4011210	12/31/1991	12/14/2001	87	2480	0.237
Australia	First Australia Fund	IAF	301110	12/26/1985	12/31/2004	63	4742	0.068
Australia	Kleinwort Benson Australian Fund	KBA	26157B10	12/26/1986	11/19/1999	41	3244	0.072
Austria	Austria Fund	OST	5258710	10/26/1989	4/26/2002	194	2982	0.099
Brazil	Brazil Fund	BZF	10575910	4/15/1988	12/30/2004	213	4013	0.287
Brazil	Brazilian Equity Fund	BZL	10588410	4/15/1992	12/30/2004	143	3069	0.250
Chile	Chile Fund	CH	16883410	10/12/1989	12/30/2004	131	3719	0.288
China	China Fund	CHN	16937310	10/1/1992	12/31/2004	186	2959	0.141
China	Greater China Fund	GCH	39167B10	10/1/1992	12/31/2004	186	2956	0.197
China	Jardine Fleming China Region Fund	JFC	46614T10	10/1/1992	12/31/2004	186	2955	0.091
China	Templeton China World Fund	TCH	88018X10	10/1/1992	8/8/2003	160	2338	0.159
Czech	Czech Republic Fund	CRF	21923Y10	10/28/1994	2/20/1998	54	806	0.085
France	France Fund	FRN	35177610	7/14/1986	12/8/1989	29	864	0.083
France	France Growth Fund	FRF	35177K10	5/24/1990	6/18/2004	97	3459	0.212
Germany	Germany Fund	GER	37414310	11/19/1986	12/30/2004	117	4540	0.057
Germany	New Germany Fund	GF	64446510	4/16/1990	12/30/2004	91	3675	0.157
Germany	Emerging Germany Fund	FRG	26156W10	4/16/1990	2/11/1999	65	2177	N.A.
India	India Fund	IFN	45408910	3/10/1994	12/31/2004	169	2570	0.282
India	India Growth Fund	IGF	45409010	8/15/1988	5/22/2003	226	3187	0.229
India	Morgan Stanley India Investment Fund	IIF	61745C10	3/10/1994	12/31/2004	169	2567	0.274
India	Jardine Fleming India Fund	JFI	78657R10	3/10/1994	6/27/2003	155	2192	0.378
Indonesia	Indonesia Fund	IF	45577810	3/27/1990	11/12/2001	142	2811	0.132
Indonesia	Jakarta Growth Fund	JGF	47012010	4/26/1990	4/11/2000	123	2405	0.088
Ireland	Irish Investment Fund	IRL	64567310	4/16/1990	12/31/2004	108	3613	0.193
Israel	First Israel Fund	ISL	32063L10	3/5/1993	12/30/2004	229	2843	0.196
Italy	Italy Fund	ITA	46539510	4/25/1986	2/13/2003	95	4187	0.189
Japan	Japan Equity Fund	JEQ	47105710	9/15/1992	12/30/2004	167	2953	0.083
Japan	Japan Fund	JPN	47107010	1/14/1980	8/13/1987	97	1828	0.065
Japan	Japan OTC Equity Fund	JOF	47109U10	3/20/1990	12/30/2004	200	3533	0.128
Korea	Fidelity Advisor Korea Fund	FAK	31580410	12/29/1994	6/30/2000	68	1367	0.108
Korea	Korea Fund	KF	50063410	10/7/1987	12/30/2004	223	4914	0.253
Korea	Korean Investment Fund	KIF	50063710	3/10/1992	11/26/2001	119	2347	0.200
Korea	Korea Equity Fund	KEF	50063B10	12/29/1993	12/30/2004	142	2654	0.251
Malaysia	Malaysia Fund	MF	56090510	5/12/1987	12/31/2004	196	4257	0.086
Mexico	Emerging Mexico Fund	MEF	29089110	10/12/1990	4/6/1999	73	2076	0.111
Mexico	Mexico Equity & Income Fund	MXE	59283410	10/12/1990	12/31/2004	107	3519	0.201
Mexico	MEXICO FUND INC	MXF	59283510	2/5/1990	12/31/2004	111	5834	0.198
Pakistan	Pakistan Investment Fund	PKF	69584410	1/10/1994	6/22/2001	160	1736	0.229

Philippines	First Philippine Fund	FPF	33610010	4/9/1990	6/18/2003	104	3327	0.131
Portugal	Portugal Fund	PGF	74337620	12/1/1989	12/22/2000	129	2687	N.A.
Russia	Templeton Russia Fund	TRF	88022F10	11/6/1995	12/31/2004	126	2216	0.081
Singapore	Singapore Fund	SGF	82929L10	8/9/1990	12/31/2004	98	3543	0.128
South Africa	Southern Africa Fund	SOA	84215710	8/9/1990	11/18/2004	79	2285	0.322
South Africa	New South Africa Fund	NSA	64880R10	8/9/1990	6/1/1999	30	958	0.198
South Africa	ASA LTD	ASA	G3156P10	8/9/1990	12/31/2004	81	2312	0.216
Spain	Growth Fund of Spain	GSP	39987710	3/19/1990	12/7/1998	111	2116	0.137
Spain	Spain Fund	SNF	84633010	6/29/1988	12/30/2004	183	3987	0.115
Switzerland	Swiss Helvetia Fund	SWZ	87087510	12/24/1987	12/30/2004	113	4269	0.188
Taiwan	Taiwan Equity Fund	TYW	87403110	7/29/1994	5/5/2000	103	1362	0.202
Taiwan	Taiwan Fund	TWN	87403610	9/28/1987	12/31/2004	274	4276	0.220
Thailand	Thai Fund	TTF	88290410	3/2/1988	12/30/2004	230	4027	0.055
Thailand	Thai Capital Fund	TC	88290520	7/9/1990	2/6/2001	149	2556	0.068
Turkey	Turkish Investment Fund	TKF	90014510	4/3/1992	12/29/2004	146	3655	0.169
UK	United Kingdom Fund	UKM	91076610	8/31/1987	4/26/1999	46	2914	0.085

Notes to Table 1

Data are drawn from the CRSP database on WRDS beginning 1/1/80 (or that the date the CCEF was first listed) and ending 12/31/04 (or the date that the CCEF was delisted). Closed-end funds are identified by the second digit in SHARE CODE and, from this set, CCEFs are identified by the fund name. CCEFs that cover more than one country (with one exception noted below) are excluded. Several changes occurred during the sample period: First Australia became Aberdeen Australia after 5/1/2001; Czech Fund renamed Central European Value Fund after 2/20/1998; and Jardine Fleming India became the Saffron Fund after 6/27/2003 and coverage expands to Bangladesh, Pakistan, and Sri Lanka, which constitute 10% of total fund assets. Thai Capital Fund has price = 0 from 2/7/2001 - 12/26/2003. T_H and T_S are the number of holidays and standard days, respectively. INST is the percentage of outstanding equity owned by institutions; see the notes to Table 5 for details. N.A. indicates data are not available.

Table 2: Estimates Of Noise, Information, And The Noise Ratio Returns Distributed Normal And Trading Volume U.S. CCEFs

	Returns						Volume
Country Closed End Fund	**NR**	$\underline{V}^N$	**SE(** $\underline{V}^N$ **)**	$\bar{V}^I$	**SE(** $\bar{V}^I$ **)**	**p-value**	**NR-VOL**
(1)	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**	**(7)**	**(8)**
Argentina Fund	0.515	2.392	0.363	2.251	0.386	0.000	0.468
First Australia	0.744	2.630	0.469	0.905	0.474	0.000	0.528
Kleinwort Benson Australian Fund	0.870	1.514	0.334	0.226	0.337	0.000	0.566
Austria Fund	0.103	3.356	0.341	29.085	0.907	0.378	0.412
Brazil Fund	0.953	5.885	0.570	0.290	0.587	0.000	0.487
Brazilan Equity Fund	0.602	4.624	0.547	3.060	0.581	0.000	0.500
Chile Fund	0.662	3.088	0.382	1.575	0.397	0.000	0.434
China Fund	0.898	3.856	0.400	0.438	0.415	0.000	0.525
Greater China Fund	0.775	3.815	0.396	1.107	0.416	0.000	0.519
Jardine Fleming China Region	0.930	3.955	0.410	0.297	0.425	0.000	0.516
Templeton China World Fund	0.672	2.431	0.272	1.186	0.292	0.000	0.451
Czech Republic Fund	1.000	3.442	0.663	0.000	0.684	0.000	0.443
France Fund	0.322	2.022	0.531	4.265	0.611	0.013	0.407
France Growth Fund	0.751	1.982	0.285	0.656	0.292	0.000	0.430
Germany Fund	0.828	3.746	0.490	0.779	0.499	0.000	0.511
New Germany Fund	0.897	2.659	0.394	0.305	0.400	0.000	0.442
Emerging Germany Fund	1.000	3.526	0.619	0.000	0.628	0.000	0.498
India Fund	0.583	2.807	0.305	2.004	0.334	0.000	0.495
India Growth Fund	0.827	3.947	0.240	0.828	0.268	0.000	0.459
Morgan Stanley India	0.640	3.180	0.346	1.786	0.373	0.000	0.510
Jardine Fleming India	0.543	2.625	0.298	2.21	0.332	0.000	0.475
Indonesia Fund	0.906	10.065	1.195	1.044	1.231	0.000	0.497
Jakarta Growth Fund	0.719	6.496	0.828	2.533	0.868	0.000	0.501
Irish Investment Fund	1.000	4.477	0.609	0.000	0.618	0.000	0.522
First Israel Fund	0.685	2.028	0.190	0.933	0.205	0.000	0.411
Italy Fund	0.525	2.525	0.366	2.282	0.381	0.000	0.426
Japan Equity Fund	0.888	3.640	0.398	0.461	0.412	0.000	0.498
Japan Fund	1.000	4.097	0.588	0.000	0.604	0.000	0.481
Japan OTC Equity Fund	0.724	4.372	0.437	1.668	0.460	0.000	0.489
Fidelity Korea Fund	0.798	5.376	0.922	1.363	0.957	0.000	0.540
Korea Fund	0.772	6.509	0.616	1.921	0.639	0.000	0.518
Korean Investment Fund	0.705	4.759	0.617	1.995	0.648	0.000	0.508
Korea Equity Fund	0.994	5.880	0.698	0.033	0.716	0.000	0.537
Malaysia Fund	0.930	6.762	0.683	0.506	0.701	0.000	0.478
Emerging Mexico Fund	0.615	3.999	0.662	2.498	0.692	0.000	0.455
Mexico Equity and Income	0.576	2.609	0.357	1.923	0.373	0.000	0.440
Mexico Fund	0.311	2.485	0.334	5.498	0.365	0.000	0.509
Pakistan Investment Fund	0.758	6.175	0.690	1.969	0.744	0.000	0.502

First Philippine Fund	1.000	6.557	0.909	0.000	0.923	0.000	0.468
Portugal Fund	0.705	2.535	0.316	1.063	0.331	0.000	0.496
Templeton Russia Fund	0.938	9.173	1.156	0.602	1.192	0.000	0.460
Singapore Fund	0.522	1.981	0.283	1.817	0.297	0.000	0.488
Southern Africa Fund	0.990	2.547	0.405	0.025	0.412	0.000	0.478
New South Africa Fund	0.400	1.052	0.272	1.576	0.297	0.006	0.387
ASA Ltd	1.000	6.133	0.964	0.000	0.980	0.000	0.500
Growth Fund of Spain	0.890	2.409	0.323	0.298	0.334	0.000	0.459
Spain Fund	1.000	5.000	0.523	0.000	0.535	0.000	0.462
Swiss Helvetia Fund	0.581	1.904	0.253	1.372	0.263	0.000	0.504
Taiwan Equity Fund	0.944	3.278	0.457	0.195	0.476	0.000	0.423
Taiwan Fund	1.000	9.999	0.854	0.000	0.881	0.000	0.515
Thai Fund	0.786	5.978	0.557	1.628	0.583	0.000	0.480
Thai Capital Fund	1.000	8.820	1.022	0.000	1.051	0.000	0.486
Turkish Investment Fund	0.651	6.206	0.726	3.329	0.760	0.000	0.489
United Kingdom Fund	1.000	4.445	0.927	0.000	0.934	0.000	0.489
Mean	0.767						0.481
Median	0.781						0.489

Notes to Table 2

See Notes to Table 1 for details about the data. The noise ratio for returns (NR) in column 2 is computed according to equation (4b) for U.S. CCEFs. Returns are assumed to be distributed normal and are adjusted for exchange rates and the U.S. market return as discussed in Section II. $\underline{V}^N$ and $\bar{V}^I$ are maximum likelihood estimates, which are equivalent to method of moments estimates under the assumptions used to construct the estimates in this table. Standard errors (SE) are computed as follows: $SE[\underline{V}^N] = \{(2/T_H)* V_H^2\}^{(1/2)}$ and $SE[\bar{V}^I] = \{(2/T_H)*V_H^2 + (2/T_S)*V_S^2\}^{(1/2)}$. The values of $\underline{V}^N$, $\bar{V}^I$, and the SE's reported in Table 2 are multiplied by 10^4. Column 7 contains the p-value for evaluating the null hypothesis that $\underline{V}^N = 0.10*(\bar{V}^N + \bar{V}^I)$ from a one-sided t-test (i.e., NR is equal to or less than 10%). The noise ratio for volume (NR^{VOL}) in column 8 is computed according to equation (5) for U.S. CCEFs.

Table 3: Estimates Of Noise, Information, And The Noise Ratio
Returns Distributed Scaled-t
U.S. CCEFs

Country Closed End Fund (CCEF)	**NR**	$\underline{V}^N$		$\overline{V}^I$		**k**		**p-value**
		Excess	***Raw***	***Excess***	***Raw***	***Excess***	***Raw***	
(1)	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**	**(7)**	**(8)**	**(9)**
Argentina Fund	0.611	3.134	3.168	1.532	1.962	4	4	0.000
First Australia	0.965	3.077	3.072	0.035	0.117	4	4	0.000
Kleinwort Benson Australian Fund	0.959	1.620	1.460	0.052	0.230	6	6	0.000
Austria Fund	0.782	4.531	4.813	0.973	0.980	3	3	0.000
Brazil Fund	0.598	5.355	7.619	0.679	1.334	4	3	0.000
Brazilan Equity Fund	0.694	5.503	5.621	1.767	2.304	4	4	0.000
Chile Fund	0.648	2.933	2.778	1.352	1.746	3	3	0.000
China Fund	0.802	4.049	3.837	0.558	1.215	4	4	0.000
Greater China Fund	0.807	4.464	4.131	0.443	1.403	4	4	0.000
Jardine Fleming China Region	0.812	4.074	4.680	0.128	0.340	5	4	0.000
Templeton China World Fund	0.654	2.942	3.243	0.760	1.252	5	4	0.000
Czech Republic Fund	0.991	2.926	2.924	0.029	0.030	4	4	0.000
France Fund	0.619	3.215	3.973	0.692	1.216	4	3	0.000
France Growth Fund	0.647	2.002	2.219	0.554	0.874	5	5	0.000
Germany Fund	0.747	4.158	4.739	0.860	0.827	3	3	0.000
New Germany Fund	0.800	2.845	3.055	0.030	0.503	4	4	0.000
Emerging Germany Fund	0.810	3.209	2.998	0.032	0.963	5	4	0.000
India Fund	0.578	3.116	3.244	1.684	2.151	6	5	0.000
India Growth Fund	0.893	5.013	4.969	0.621	0.646	3	3	0.000
Morgan Stanley India	0.720	3.784	4.023	1.241	1.235	5	5	0.000
Jardine Fleming India	0.487	2.704	3.150	2.176	2.403	5	4	0.000
Indonesia Fund	0.963	11.977	11.150	0.111	1.283	3	3	0.000
Jakarta Growth Fund	0.904	9.549	8.587	0.993	1.982	3	3	0.000
Irish Investment Fund	0.708	2.283	3.200	0.023	0.026	4	3	0.000
First Israel Fund	0.711	2.211	2.390	0.740	0.720	4	4	0.000
Italy Fund	0.798	4.266	4.585	0.817	0.765	3	3	0.000
Japan Equity Fund	0.737	3.822	4.939	0.205	0.244	5	4	0.000
Japan Fund	0.940	2.991	3.145	0.034	0.036	4	4	0.000
Japan OTC Equity Fund	0.654	5.115	5.998	1.027	1.821	5	4	0.000
Fidelity Korea Fund	0.523	4.999	6.444	2.089	3.111	4	3	0.000
Korea Fund	0.585	5.012	5.821	1.096	2.739	4	3	0.000
Korean Investment Fund	0.561	5.282	6.825	1.553	2.583	4	3	0.000
Korea Equity Fund	0.864	5.938	6.807	0.059	0.067	6	5	0.000
Malaysia Fund	0.935	7.506	7.943	0.654	0.084	3	3	0.000
Emerging Mexico Fund	0.632	4.598	4.598	1.754	2.676	4	4	0.000
Mexico Equity and Income	0.729	4.043	4.371	1.224	1.174	3	3	0.000
Mexico Fund	0.381	3.703	4.286	5.029	5.438	3	3	0.000

Pakistan Investment Fund	0.696	5.990	6.184	2.525	2.420	4	4	0.000
First Philippine Fund	0.929	5.311	5.551	0.264	0.169	4	4	0.000
Portugal Fund	0.907	3.878	4.161	0.180	0.115	3	3	0.000
Templeton Russia Fund	0.659	9.264	10.946	3.030	3.103	3	3	0.000
Singapore Fund	0.566	2.476	2.892	1.557	1.484	4	4	0.000
Southern Africa Fund	0.702	2.114	1.953	0.798	1.061	3	3	0.000
New South Africa Fund	0.629	1.651	1.891	1.183	0.734	3	3	0.000
ASA Ltd	1.000	6.440	6.334	0.050	0.050	3	3	0.000
Growth Fund of Spain	0.692	2.336	2.563	0.352	0.812	5	4	0.000
Spain Fund	0.945	5.350	5.611	0.050	0.053	3	3	0.000
Swiss Helvetia Fund	0.879	2.570	2.447	0.098	0.478	3	3	0.000
Taiwan Equity Fund	0.679	2.787	3.252	0.799	0.850	5	4	0.000
Taiwan Fund	0.806	6.300	6.849	1.000	0.964	3	3	0.000
Thai Fund	0.832	8.428	7.556	1.445	2.569	3	3	0.000
Thai Capital Fund	0.915	9.084	7.887	0.528	2.036	3	3	0.000
Turkish Investment Fund	0.647	6.765	6.695	2.384	3.755	5	4	0.000
United Kingdom Fund	0.839	2.445	2.885	0.025	0.030	5	4	0.000
Mean	0.752							
Median	0.733							

Notes to Table 3

See Notes to Table 1 for details about the data. "Excess" in columns 3, 5, and 7 refers to estimates based on returns that have been adjusted for exchange rate and U.S. market movements. "Raw" in columns 4, 6, and 8 refers to estimates based on returns that have been adjusted only for exchange rate movements. The noise ratio for returns (NR) in column 2 is computed according to equation (4b), $NR \equiv \underline{V}^{N,Excess} / (\bar{V}^{N,Raw} + \bar{V}^{I,Raw})$. Returns are assumed to be distributed Scaled-t as described in fn. 9. The parameters $\{ V^{N}, V^{I}, k \}$ for both excess and raw returns are estimated by maximum likelihood based on a grid search run separately for both excess and raw returns. (See Section 3 of the Appendix for further discussion.) The parameter space for the grid search is determined by the admissible values in equation (3): $V^{N,X} = \{(1/m)V_H, (2/m)V_H, \ldots, V_H\}$, $V^{I,X} = \{(1/m)V_S, (2/m)V_S, \ldots, V_S\}$, $k = \{3, 4, \ldots, 30\}$, where the parameter determining the size of the increments is $m = 400$ and $X = \{Excess, Raw\}$. When k equals 30, the Scaled-t distribution approximates a normal distribution. The values of $\underline{V}^{N}$ and $\bar{V}^{I}$ reported in Table 3 are multiplied by 10^4. Column 9 contains the p-value for evaluating the null hypothesis that NR = 0.10. The LR statistic is computed by restricting $\underline{V}^{N,Excess} = 0.10 * (\bar{V}^{N,Raw} + \bar{V}^{I,Raw})$, the latter sum having been estimated previously from the Raw data. For robustness, a parallel LR test was performed where $(\bar{V}^{N,Raw} + \bar{V}^{I,Raw}) = \underline{V}^{N,Excess} / 0.10$, with the latter variance having been estimated previously from the Excess data. The results of the LR test are not qualitatively different between methods, so the first test (restricting excess returns based on estimates from raw returns) is reported in column 9.

Table 4: Estimates Of Noise, Information, And The Noise Ratio Returns Distributed Scaled-t Information Affects Holiday Returns ($\lambda \geq 0$) U.S. CCEFs

Country Closed End Fund	NR	$\underline{V}^N$		$\overline{V}^I$		k		λ	p-value
		Excess	Raw	Excess	Raw	Excess	Raw		
(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Argentina Fund	0.611	3.134	3.168	1.532	1.962	4	4	0.0	0.000
First Australia	0.959	3.090	0.013	0.018	3.208	4	4	0.5	0.000
Kleinwort Benson Australian Fund	0.009	0.015	0.014	1.654	1.678	6	6	0.5	0.000
Austria Fund	0.424	2.652	0.037	2.920	6.210	3	3	0.5	0.004
Brazil Fund	0.496	4.414	5.885	1.667	3.018	4	3	0.5	0.000
Brazilan Equity Fund	0.694	5.503	5.621	1.767	2.304	4	4	0.0	0.000
Chile Fund	0.648	2.933	2.778	1.352	1.746	3	3	0.0	0.000
China Fund	0.753	3.818	3.271	0.816	1.798	4	4	0.3	0.000
Greater China Fund	0.796	4.388	3.436	0.541	2.076	4	4	0.3	0.000
Jardine Fleming China Region	0.797	4.015	4.574	0.191	0.462	5	4	0.2	0.000
Templeton China World Fund	0.654	2.942	3.243	0.760	1.252	5	4	0.0	0.000
Czech Republic Fund	0.956	2.909	0.018	0.014	3.024	4	4	0.5	0.000
France Fund	0.619	3.215	3.973	0.692	1.216	4	3	0.0	0.001
France Growth Fund	0.647	2.002	2.219	0.554	0.874	5	5	0.0	0.000
Germany Fund	0.527	2.960	0.041	2.036	5.570	3	3	0.5	0.002
New Germany Fund	0.803	2.859	2.692	0.015	0.870	4	4	0.5	0.000
Emerging Germany Fund	0.807	3.262	0.015	0.016	4.027	5	4	0.5	0.000
India Fund	0.578	3.116	3.244	1.684	2.151	6	5	0.0	0.000
India Growth Fund	0.711	4.303	0.042	1.289	6.009	3	3	0.5	0.000
Morgan Stanley India	0.720	3.784	4.023	1.241	1.235	5	5	0.0	0.000
Jardine Fleming India	0.367	2.022	2.294	2.853	3.221	5	4	0.3	0.002
Indonesia Fund	0.936	11.776	0.104	0.111	12.483	3	3	0.5	0.000
Jakarta Growth Fund	0.904	9.549	8.587	0.993	1.982	3	3	0.0	0.000
Irish Investment Fund	0.737	2.350	3.177	0.012	0.013	4	3	0.5	0.000
First Israel Fund	0.701	2.170	2.251	0.800	0.845	4	4	0.1	0.000
Italy Fund	0.798	4.266	4.585	0.817	0.765	3	3	0.0	0.000
Japan Equity Fund	0.720	3.749	4.764	0.287	0.439	5	4	0.5	0.000
Japan Fund	0.914	3.011	0.020	0.017	3.274	4	4	0.5	0.000
Japan OTC Equity Fund	0.654	5.115	5.998	1.027	1.821	5	4	0.0	0.000
Fidelity Korea Fund	0.161	1.559	0.053	5.593	9.623	4	3	0.5	0.402
Korea Fund	0.432	3.710	3.577	2.445	5.005	4	3	0.5	0.000
Korean Investment Fund	0.561	5.282	6.825	1.553	2.583	4	3	0.0	0.000
Korea Equity Fund	0.876	5.968	6.715	0.030	0.100	6	5	0.5	0.000
Malaysia Fund	0.935	7.506	7.943	0.654	0.084	3	3	0.0	0.000
Emerging Mexico Fund	0.606	4.439	4.345	1.884	2.982	4	4	0.1	0.003
Mexico Equity and Income	0.729	4.043	4.371	1.224	1.174	3	3	0.0	0.000
Mexico Fund	0.381	3.703	4.286	5.029	5.438	3	3	0.0	0.006

Pakistan Investment Fund	0.405	3.643	0.062	4.968	8.928	4	4	0.5	0.048
First Philippine Fund	0.909	5.180	5.417	0.370	0.281	4	4	0.4	0.000
Portugal Fund	0.907	3.878	4.161	0.180	0.115	3	3	0.0	0.000
Templeton Russia Fund	0.436	6.146	7.758	6.158	6.322	3	3	0.5	0.000
Singapore Fund	0.566	2.476	2.892	1.557	1.484	4	4	0.0	0.018
Southern Africa Fund	0.464	1.401	1.085	1.518	1.932	3	3	0.5	0.002
New South Africa Fund	0.629	1.651	1.891	1.183	0.734	3	3	0.0	0.019
ASA Ltd	0.993	6.409	6.430	0.025	0.025	3	3	0.5	0.000
Growth Fund of Spain	0.456	1.566	0.024	1.110	3.410	5	4	0.5	0.011
Spain Fund	0.940	5.425	0.026	0.025	5.747	3	3	0.5	0.000
Swiss Helvetia Fund	0.852	2.494	2.157	0.164	0.771	3	3	0.4	0.000
Taiwan Equity Fund	0.404	1.672	0.034	1.911	4.103	5	4	0.5	0.005
Taiwan Fund	0.605	4.700	5.327	2.600	2.447	3	3	0.5	0.000
Thai Fund	0.832	8.428	7.556	1.445	2.569	3	3	0.0	0.000
Thai Capital Fund	0.851	8.731	0.080	0.830	10.182	3	3	0.5	0.000
Turkish Investment Fund	0.647	6.765	6.695	2.384	3.755	5	4	0.0	0.000
United Kingdom Fund	0.830	2.423	0.027	0.038	2.890	5	4	0.5	0.000
Mean	0.673								
Median	0.698								

Notes to Table 4

See Notes to Table 1 for details about the data. "Excess" in columns 3, 5, and 7 refers to estimates based on returns that have been adjusted for exchange rate and U.S. market movements. "Raw" in columns 4, 6, and 8 refers to estimates based on returns that have been adjusted only for exchange rate movements. The noise ratio for returns (NR) in column 2 is computed according to equation (4b), $NR \equiv \underline{V}^{N,Excess} / (\overline{V}^{N,Raw} + \overline{V}^{I,Raw})$. Returns are assumed to be distributed Scaled-t as described in fn. 9. The parameters $\{V^N, V^I, k, \lambda\}$ for both excess and raw returns are estimated by maximum likelihood based on a grid search run separately for both excess and raw returns. (See Section 3 of the Appendix for further discussion.) The parameter space for the grid search is determined by the admissible values in equation (3): $V^{N,X} = \{(1/m)V_H, (2/m)V_H,\ldots, V_H\}$, $V^{I,X} = \{(1/m)V_S, (2/m)V_S,\ldots, V_S\}$, $k = \{3, 4,\ldots, 30\}$, $\lambda = \{0.1, 0.1,\ldots,0.5\}$, where the parameter determining the size of the increments is m = 400 and X={Excess, Raw}. When k equals 30, the Scaled-t distribution approximates a normal distribution. The values of V^N and V^I reported in Table 4 are multiplied by 10^4. Column 9 gives estimates for λ, the information leakage parameter. This variable is estimated using excess returns only. Column 10 contains the p-value for evaluating the null hypothesis that NR = 0.10. The LR statistic is computed by restricting $\underline{V}^{N,Excess} = 0.10 * (\overline{V}^{N,Raw} + \overline{V}^{I,Raw})$, the latter sum having been estimated previously from the Raw data. For robustness, a parallel LR test was performed where $(\overline{V}^{N,Raw} + \overline{V}^{I,Raw}) = \underline{V}^{N,Excess} / 0.10$, with the latter variance having been estimated previously from the Excess data. The results of the LR test are not qualitatively different between methods, so the first test (restricting excess returns based on estimates from raw returns) is reported in column 10.

Table 5: OLS Estimates of the Elasticity of Noise with Respect to Institutional Ownership, U.S. CCEFs

Noise Ratio (1)	Elasticity (2)	α (3)	β (4)	Adj. R^2 (5)
A. Complete Sample				
Normal	-0.097	0.838	-0.443	0.007
λ = 0	(0.082)	(0.069)	(0.378)	
Scaled-t	-0.187**	0.887	-0.838	0.186
λ = 0	(0.053)	(0.043)	(0.236)	
Scaled-t	-0.163*	0.779	-0.655	0.039
λ ≠ 0	(0.093)	(0.069)	(0.374)	
B. Trimmed Sample				
Normal	-0.107	0.842	-0.501	0.005
λ = 0	(0.095)	(0.079)	(0.447)	
Scaled-t	-0.179**	0.889	-0.832	0.155
λ = 0	(0.058)	(0.047)	(0.268)	
Scaled-t	-0.193**	0.841	-0.840	0.094
λ ≠ 0	(0.080)	(0.061)	(0.346)	

Notes to Table 5

Estimates based on $NR_c = \alpha + \beta * INST_c + e_c$, where c indexes the CCEFs. The elasticity in column 2 equals β * MEAN[$INST_c$] / MEAN[NR_c]. ** and * indicate statistical significance of the estimated elasticities at the 5% and 10% levels, respectively. The NR_c's in the first, second, and third rows of panels A and B are taken from column 2 of Tables 2, 3, and 4, respectively. $INST_c$ is the percentage of outstanding equity owned by institutions (a bank, insurance company, investment companies (mutual funds) and their managers, independent investment advisor (brokerage companies), or a college/university endowment). Only institutions that manage over $100 million are required to file, and filers may omit small holdings under 10,000 shares or $200,000. The data are drawn from the CDA/Spectrum database on WRDS. $INST_c$ equals the average institutional ownership stated as a percentage of outstanding shares. It is computed as the mean of $INST_{c,t}$, where $INST_{c,t}$ is institutional ownership for CCEF c at time t and is defined as the sum over all institutions (i) of shares held by institutions at the end of the year ($SHARES_{i,c,t}$) divided by the number of shares outstanding in millions at the end of the year ($SHROUT1_{c,t}$ * 1,000,000). The results in panel A are based on a sample of 52 out of the 54 CCEFs listed in Table 1; $INST_c$ data are not available for the Emerging Germany Fund and the Portugal Fund. The results in Panel B are based on a sample of 48 CCEFs that has been trimmed to reduce the possible effect of outliers -- two CCEFs (Kleinwort Benson Australian Fund and Fidelity Korea Fund) with relatively low NR_c's (from Table 4) and two CCEFs (Southern Africa Fund and Jardine Fleming India Fund) with relatively high $INST_c$'s.

Table 6: Estimates Of Noise, Information, And The Noise Ratio Returns Distributed Normal, Scaled-t, Scaled-t With Information Affecting Holiday Returns ($\lambda \geq 0$) U.K. CCEFs

Country Closed-End Fund	Normal	Scaled-t	Scaled-t
	$\lambda = 0$	$\lambda = 0$	$\lambda \neq 0$
(1)	(2)	(3)	(4)
Canadian General Investments Ltd	0.202	0.028	0.018
JPMF Chinese	0.393	0.855	0.855
German Smaller Companies	0.532	0.968	0.049
JPMF Indian	0.497	0.661	0.326
New India IT	0.555	1.000	1.000
Edinburgh Java	0.670	1.000	0.153
Gartmore Irish Growth Fund	0.400	1.000	0.932
Baillie Gifford Japan	0.283	0.397	0.100
Baillie Gifford Shin Nippon	0.357	0.571	0.177
Edinburgh Japan	0.301	0.571	0.571
Fidelity Japanese Value	0.305	0.399	0.125
Gartmore Select Japanese	0.260	0.427	0.165
GT Japan	0.386	0.675	0.675
Henderson Japan	0.386	0.454	0.454
Invesco Japan Discovery	0.387	0.594	0.331
Invesco Tokyo	0.371	0.375	0.175
JPMF Japanese	0.102	0.614	0.614
JPMF Japanese Smaller Companies	0.338	0.679	0.153
Martin Currie Japan	0.338	0.285	0.187
Perpetual Japan	0.331	0.424	0.139
Schroder Japanese Growth	0.353	0.462	0.419
Korea-Europe Fund	0.614	0.395	0.056
Schroder Korea	0.609	0.563	0.113
JF Philippine	0.150	0.150	0.150
JPMF Russian	0.786	0.476	0.068
Old Mutual S. Africa	0.403	1.000	1.000
Aberdeen New Thai	0.391	0.851	0.250
Edinburgh US Tracker	0.521	0.262	0.006
Foreign & Colonial US Smaller	0.785	0.441	0.231
JPMF American	0.116	0.250	0.006
JPMF US Discovery	0.620	0.528	0.031
US Smaller Companies	0.386	0.576	0.576
Mean	0.410	0.560	0.316
Median	0.386	0.545	0.176

<u>Notes to Table 6</u>

See Notes to Table 1 for details about the data. The table entries in columns 2, 3, and 4 are comparable to those in column 2 of Tables 2, 3, and 4, respectively, except that the entries in Table 6 are computed for U.K. CCEFs. In column 2, returns are assumed to be distributed normal, and there is no leakage of information during holidays. In column 3, returns are assumed to be distributed Scaled-t, and there is no leakage of information during holidays. In column 4, returns are assumed to be distributed Scaled-t, and there is the possibility of leakage of information during holidays. To eliminate the upward bias for NR due to return series adjusted by the market model, the noise ratio statistic in columns 3 and 4 are computed according to equation (4b), $NR \equiv \underline{V}^{N,Excess} / (\bar{V}^{N,Raw} + \bar{V}^{I,Raw})$, where "Excess" refers to estimates based on returns that have been adjusted for exchange rate and U.S. market movements and "Raw" refers to estimates based on returns that have been adjusted only for exchange rate movements. All parameters are estimated by maximum likelihood based on a grid search. See the notes to Tables 2, 3, and 4 for details.

Table 7: Noise Ratio Means And The Ratio Of The Noise Ratios U.S. And U.K. CCEFs For The Same Country

Country	Normal, λ = 0		Scaled-t, λ = 0		Scaled-t, λ ≠ 0	
	U.S.	U.K.	U.S.	U.K.	U.S.	U.K.
(1)	(2)	(3)	(4)	(5)	(6)	(7)
China	0.819	0.393	0.769	0.855	0.750	0.855
	RNR = 0.480		RNR = 1.112		RNR = 1.140	
Germany	0.908	0.532	0.786	0.968	0.712	0.049
	RNR = 0.586		RNR = 1.232		RNR = 0.069	
India	0.648	0.526	0.669	0.831	0.594	0.663
	RNR = 0.812		RNR = 1.242		RNR = 1.116	
Indonesia	0.813	0.670	0.933	1.000	0.920	0.153
	RNR = 0.824		RNR = 1.072		RNR = 0.166	
Ireland	1.000	0.400	0.708	1.000	0.737	0.932
	RNR = 0.400		RNR = 1.412		RNR = 1.265	
Japan	0.871	0.321	0.777	0.495	0.763	0.306
	RNR = 0.369		RNR = 0.637		RNR = 0.401	
Korea	0.758	0.611	0.670	0.479	0.507	0.084
	RNR = 0.806		RNR = 0.715		RNR = 0.166	
Philippines	1.000	0.150	0.929	0.150	0.909	0.150
	RNR = 0.150		RNR = 0.161		RNR = 0.165	
Russia	0.938	0.786	0.659	0.476	0.437	0.068
	RNR = 0.838		RNR = 0.722		RNR = 0.155	
South Africa	0.797	0.403	0.780	1.000	0.695	1.000
	RNR = 0.506		RNR = 1.282		RNR = 1.438	
Thailand	0.893	0.391	0.874	0.851	0.841	0.250
	RNR = 0.438		RNR = 0.974		RNR = 0.297	

Notes to Table 7

The table entries (apart from the RNRs) are unweighted means of noise ratios (NR) for CCEFs by the country in which the assets are invested. The figures in columns 2, 4, and 6 are taken from column 2 of Tables 2, 3, and 4, respectively; the figures in columns 3, 5, and 7 from columns 2, 3, and 4, respectively, of Table 6. RNR equals the ratio of the U.K. entry to the U.S. entry for a given country.