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Earnings Quality, Insider Trading, and Cost of Capital

DAVID ABOODY,* JOHN HUGHES,* AND JING LIU†

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ABSTRACT

Previous research argues that earnings quality, measured as the unsigned abnormal accruals, proxies for information asymmetries that affect cost of capital. We examine this argument directly in two stages. In the first stage, we estimate firms' exposure to an earnings quality factor in the context of a Fama-French three-factor model augmented by the return on a factor-mimicking portfolio that is long in low earnings quality firms and short in high earnings quality firms. In the second stage, we examine whether the earnings quality factor is priced and whether insider trading is more profitable for firms with higher exposure to that factor. Generally speaking, we find evidence consistent with pricing of the earnings quality factor and insiders trading more profitably in firms with higher exposure to that factor.

1. *Introduction*

In this paper, we examine two closely related issues regarding the cost of capital effects of asymmetric information: (1) whether the systematic component of asymmetric information is priced, and (2) whether privately informed traders earn greater profits when trading stocks with higher exposure to an asymmetric information risk factor.

*The authors are from the Anderson Graduate School of Management at UCLA; †Part of the paper was completed while Liu was visiting the Cheung Kong Graduate School of Business.
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There is a growing empirical literature on both the cost of capital effects of asymmetric information (e.g., Botosan [1997], Botosan and Plumlee [2002], Botosan, Plumlee, and Xie [2004], Healy, Hutton, and Palepu [1999], Francis et al. [2005], and Easley, Hvidkjaer, and O'Hara [2002]) and the association of insider trading with asymmetric information (e.g., Rozeff and Zaman [1998], Aboody and Lev [2000], Frankel and Li [2004], Piotroski and Roulstone [2005], and Ke, Huddart, and Petroni [2003]).[1] However, apart from a study by Bhattacharya and Daouk [2002] that investigates a correlation between enforced insider trading laws and cost of capital across countries, it remains to be determined if the prospect of privately informed trading is what drives a cost of capital effect of asymmetric information. It is plausible that the risk posed by privately informed trading may be fully diversifiable by uninformed traders, implying no cost of capital effect (e.g., Hughes, Liu, and Liu(2005]).

This paper differs from prior literature on insider trading (e.g., Frankel and Li [2004], Aboody and Lev [2000]) in that we examine the relationship between insider trading and the systematic (priced) component of information risk, while the existing literature examines the relationship between insider trading and firm-specific characteristics. This aspect is subtle but important. For example, size is a firm-specific characteristic (idiosyncratic) and therefore cannot be a risk factor according to modern finance theories such as the Capital Asset Pricing Model or the Arbitrage Pricing Theory (APT). To construct a risk factor related to size, one has to use factor-mimicking portfolios à la Fama and French [1993]. Finding evidence that insiders exploit certain firm-specific information characteristics has no implications for the firm's cost of capital, since this risk could be fully diversifiable. To relate insider trading to priced information risk (cost of capital), we must first isolate that risk, which is conventionally done via a factor model such as that of Fama and French [1993].

Theoretical support for the exploitation of private information by informed traders as the explanation for a cost of capital effect of asymmetric information comes from Amihud and Mendelson [1986], Admati [1985], Dow and Gorton [1995], and Easley and O'Hara [2004]. Other theories that do not rely on privately informed trading are based on incomplete information (e.g., Merton [1986], Shapiro [2002]) and estimation risk (e.g., Barry and Brown [1985], Clarkson and Thompson [1990], and Clarkson, Guedes, and Thompson [1996]). Each of the above theories exploits the idea that the common knowledge assumption about the mean-variance matrix of asset payoffs adopted in neoclassical asset pricing theories may not hold, allowing for the prospect that investors will be differentially informed about the asset payoffs leading to a cost of capital effect if diversification is incomplete. A distinguishing characteristic of theories that rely on private information is

[1] Other studies by Seyhun [1998] and Lakonishok and Lee [2001] document profitability of insider trading without linking such profitability to specific measures of asymmetric information.

the exploitation of that information by informed traders. While we cannot rule out the possibility of multiple explanations for a cost of capital effect of earnings quality, a finding that insiders profit from trading on the priced component of earnings quality is consistent with an explanation based on asymmetric information.

Our measure for identifying firms for which privately informed trading is likely to be more pronounced, and, hence, pose a greater asymmetric information risk to uninformed traders, is earnings quality defined as unsigned abnormal accounting accruals.[2] We chose earnings quality as our measure of information asymmetry for several reasons: the findings of Francis et al. [2005] suggest that earnings quality is priced; relative to the cash flow component of earnings, accounting accruals are more prone to management discretion and manipulation, implying less private information may be preempted by earnings announcements; earnings quality is a ubiquitous construct that applies to all publicly traded firms; and unlike the probability of informed trading measure used by Easley, Hvidkjaer, and O'Hara [2002], an indirect measure of a firm's information asymmetry derived from trade data, earnings quality is a relatively more direct measure of a firm's information environment derived from fundamental accounting data contained in its financial statements. While we believe that it is reasonable to use unsigned abnormal accruals to proxy for information asymmetry, we are aware that abnormal accruals can be highly correlated with growth (Zhang [2005]). Hence, there could exist alternative interpretations of our empirical results.[3]

To be consistent with the idea of diversification in neoclassical asset pricing theory, following Francis et al. [2005] and Easley, Hvidkjaer, and O'Hara [2002], we estimate cost of capital using a factor model based on APT from ex post regressions. As in those studies, we use three factors recommended by Fama and French [1993] augmented by a fourth factor based on earnings quality to capture asymmetric information risk. If asymmetric information risk as measured by an earnings quality factor–mimicking portfolio is priced, then this risk should command a positive risk premium, and greater exposure to this risk as estimated from factor loadings should imply greater profits to privately informed trading.

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[2] Specifically, we employ four measures related to the unsigned value of a firm's discretionary accruals as defined in Francis, et al. [2005]. Two proxies are based on the modified Jones' model (Dechow, Sloan, and Sweeney [1995]) and two are based on the Dechow and Dichev [2002] model. Other proxies appearing in the literature include the Association for Investment and Research's assessment of corporate disclosure practices (Botosan [1997], Botosan and Plumlee [2002]), dispersion in analysts' earnings forecasts (Botosan, Plumlee, and Xie [2004]), and analysts' coverage (Healy, Hutton, and Palepu [1999]).

[3] For example, if growth proxies for some priced risk factor, then the pricing result we document could be due to risk factors not related to information asymmetries. Further relating the cost of capital effect to informed trading profits helps to distinguish between these competing explanations, since it is difficult to establish a link between non–information based risk factors and informed trading profits.

Lastly, we employ trades by corporate insiders (officers, directors, and principal stockholders) subsequently publicly disclosed to the Securities and Exchange Commission (SEC) as our measure of privately informed trading. While insiders are representative of informed traders, we envision that a similar information advantage extends to institutions and other professional traders, implying a sufficient impact on market order flow to constitute a great enough potential risk to induce detectable risk premiums. Since these trade data are the measure of private information we use in assessing insider trading profitability, the natural window for purposes of estimating such profitability is from the trade date to the SEC filing date.[4] Accordingly, insider's trading profits are measured by abnormal returns from the date of trade to one day after filing reports of those trades to the SEC after controlling for all risk factors including the asymmetric information risk factor.

Since both the frequency of insider trades and the profitability to each trade contribute to the total profits earned by the insiders, we examine both aspects and find both are positively correlated with the firm's exposure to the earnings quality (asymmetric information) risk factor. On the buy side, based on results for a hedge portfolio that is long in the highest quintile of asymmetric information risk factor loading firms and short in the lowest quintile, the difference in abnormal returns to insiders ranges from 1.339% to 3.344% for an average holding period of 27 days. Sell side results have the correct signs, but are smaller in magnitude and weaker in statistical significance, a finding consistent with the popularity of stock-based compensation and the prospect that insider sell transactions are more likely than buy transactions to be motivated by diversification (e.g., Ofek and Yermack [2000]) and consumption.

The results for the pricing effect of earnings quality are relatively weak but consistent. We find that while the quintile portfolios with highest exposure to the asymmetric information factor have statistically significant positive intercepts (i.e., Jensen's alphas) after controlling for the Fama and French [1993] risk factors, the quintile portfolios with the lowest exposure have alphas that are insignificantly different from zero. Results for the hedge portfolio have correct signs, with estimates ranging from 0.992% per month to 1.178% per month, but weak statistical significance levels, potentially due to noise in the system from the low-exposure firms, or the diversification effect (e.g., Hughes, Liu, and Liu [2005]).

The combination of these findings is supportive of the conclusions that asymmetric information risk, as measured by the firm's exposure to an earnings quality factor–mimicking portfolio, has measurable pricing effects, and this risk is positively associated with expected insider trading profits.

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[4] Note that studies that explore insider trading profitability over longer windows (e.g., Piotroski and Roulstone [2005] and Ke, Huddart, and Petroni [2003]) condition their tests on the assumption that information revealed ex post was known to insiders ex ante. In contrast, we condition on just the portion of insiders' private information contained in trades reported to the SEC. Market efficiency in the semi-strong form dictates that one cannot predict abnormal returns to this information beyond the filing date.

The remainder of this paper is organized as follows. In section 2, we describe how we measure earnings quality. In section 3, we describe the sample data employed and provide descriptive statistics. Our empirical analyses of the cost of capital effects of asymmetric information risk and its association with insider trading are contained in section 4. We conclude in section 5 with a summary of our results and a discussion of how they relate to extant theory.

2. *Earnings Quality*

We employ two approaches for deriving measures of earnings quality that are also used by Francis et al. [2005]. Following Dechow, Sloan, and Sweeney [1995], one is based on estimates of abnormal accruals and the other, following Dechow and Dichev [2002], is based on the extent to which working capital accruals map into cash flow realizations. Both approaches rely on accounting fundamentals to separate accruals into nondiscretionary (normal) and discretionary (abnormal) components. Earnings quality is defined as the absolute value of the abnormal component. The larger the absolute value, the lower is earnings quality. Specifically, total accruals ($TA_{j,t}$), total current accruals ($TCA_{j,t}$), and cash flow from operations ($CFO_{j,t}$) for firm j and year t are calculated as shown below:

$$TA_{j,t} = (\Delta CA_{j,t} - \Delta CL_{j,t} - \Delta CASH_{j,t} + \Delta STDEBT_{j,t} - DEPN_{j,t})$$

$$TCA_{j,t} = (\Delta CA_{j,t} - \Delta CL_{j,t} - \Delta CASH_{j,t} + \Delta STDEBT_{j,t})$$

$$CFO_{j,t} = NIBE_{j,t} - TA_{j,t}$$

where:

$\Delta CA_{j,t}$ = firm j's change in current assets (Compustat #4) in year t,
$\Delta CL_{j,t}$ = firm j's change in current liabilities (Compustat #5) in year t,
$\Delta CASH_{j,t}$ = firm j's change in cash (Compustat #1) in year t,
$\Delta STDEBT_{j,t}$ = firm j's change in short-term debt (Compustat #34) in year t,
$DEPN_{j,t}$ = firm j's depreciation and amortization expense (Compustat #14) in year t, and
$NIBE_{j,t}$ = firm j's net income before extraordinary items (Compustat #18) in year t.

To estimate abnormal accruals ($AA_{j,t}$) for firm j in year t, we perform the following cross-sectional regression for each of Fama and French's [1997] 48 industry groups containing at least 20 firms in each year:

$$\frac{TA_{j,t}}{Asset_{j,t-1}} = k_{1,t}\frac{1}{Asset_{j,t-1}} + k_{2,t}\frac{\Delta REV_{j,t}}{Asset_{j,t-1}} + k_{3,t}\frac{PPE_{j,t}}{Asset_{j,t-1}} + \varepsilon_{j,t} \quad (1)$$

where:

$\Delta REV_{j,t}$ = firm j's change in revenues (Compustat #12) in year t,
$PPE_{j,t}$ = firm j's gross value of property, plant, and equipment (Compustat #7) in year t, and we have deflated by firm j's total assets in year t−1 ($Assets_{j,t-1,}$ Compustat #6).

We then use the industry-year-specific parameter estimates from (1) to estimate firm-specific normal accruals ($NA_{j,t}$) for firm j in year t as a percent of lagged total assets; that is,

$$NA_{j,t} = \hat{k}_{1,t}\frac{1}{Asset_{j,t-1}} + \hat{k}_{2,t}\frac{(\Delta REV_{j,t} - \Delta AR_{j,t})}{Asset_{j,t-1}} + \hat{k}_{3,t}\frac{PPE_{j,t}}{Asset_{j,t-1}}$$

where:

$AR_{j,t}$ = firm j's change in accounts receivable (Compustat #2) in year t.[5]

In turn, abnormal accruals ($AA_{j,t}$) for firm j in year t are

$$AA_{j,t} = \frac{TA_{j,t}}{Asset_{j,t-1}} - NA_{j,t}.$$

The absolute value of abnormal accruals ($|AA_{j,t}|$) is the first earnings quality measure (EQ1), with larger values indicating lower earnings quality.

Similar to total accruals, we estimate abnormal current accruals ($ACA,_{jt}$) using the following variation of (1):

$$\frac{TCA_{j,t}}{Asset_{j,t-1}} = \gamma_{1,\tau}\frac{1}{Asset_{j,t-1}} + \gamma_{2,\tau}\frac{\Delta REV_{j,t}}{Asset_{j,t-1}} + v_{j,t}. \quad (2)$$

We use the parameter estimates from equation (2) to calculate each firm's normal current accruals as a percent of lagged assets,

$$NCA_{j,t} = \hat{\gamma}_{1,t}\frac{1}{Asset_{j,t-1}} + \hat{\gamma}_{2,t}\frac{(\Delta REV_{j,t} - \Delta AR_{j,t})}{Asset_{j,t-1}},$$

and then calculate the abnormal component for firm j in year t as:

$$ACA_{j,t} = \frac{TCA_{j,t}}{Asset_{j,t-1}} - NCA_{j,t}.$$

Our second earnings quality measure (EQ2) is the absolute value of the abnormal current accruals ($|ACA_{j,t}|$) calculated as shown above. Similar to EQ1, larger values of EQ2 indicate poorer earnings quality.

[5] Note that a prominent feature of the modified Jones model à la Dechow, Sloan, and Sweeney [1995] is that changes in accounts receivables are included in revenues in the estimation of the model parameters, but are deducted from revenues in the event year. Dechow, Sloan, and Sweeney [1995] find this model exhibits the most power in detecting earnings management among other models they consider.

The two remaining measures of earnings quality proceed from estimates of total current accruals based on cash flows from operations, again employing Fama and French's [1997] 48 industries for each year with at least 20 firms:

$$\frac{TCA_{j,t}}{Aveasset_{j,t}} = \theta_{0,j} + \theta_{1,j}\frac{CFO_{j,t-1}}{Aveasset_{j,t}} + \theta_{2,j}\frac{CFO_{j,t}}{Aveasset_{j,t}} + \theta_{3,j}\frac{CFO_{j,t+1}}{Aveasset_{j,t}} + v_{j,t} \quad (3)$$

where:

$Aveasset_{j,t}$ = firm j's average total assets over years t and $t-1$.

Our third earnings quality measure (EQ3) is the absolute value of the firm's residual ($|\hat{v}_{j,t}|$) from equation (3), and our fourth earnings quality measure (EQ4) is the time-series standard deviation of these firm-specific residuals ($\sigma(\hat{v}_{j,t,})$) calculated using a minimum of five residual observations. Consistent with the construction of the other measures, larger absolute residuals and larger standard deviations of residuals are interpreted as lower earnings quality.

3. *Sample and Descriptive Statistics*

We gather accounting and cash flow data from Compustat and price data from the Center for Research in Security Prices (CRSP). The insider trading data are obtained from the CDA/Investnet part of Thompson/First Call. The database contains all buy and sell transactions made by corporate insiders and reported to the SEC from January, 1985, through November, 2003. Corporate insiders are defined by the 1934 Securities and Exchange Act as corporate officers, directors, and owners of 10% or more of any equity class of securities.

We estimate each of the four earnings quality measures annually over the years 1985 to 2003, yielding 989,530 firm-month observations for EQ1 and EQ2, 910,477 observations for EQ3, and 614,981 observations for EQ4.[6] Table 1 contains descriptive statistics of the variables used in our analysis.

It is evident from panel A of table 1 that sample distributions of EQ1 and EQ2 are highly skewed, with means being substantially higher than medians. This suggests the presence of outliers in the positive extremes, caused partially by deflation using total assets when total assets for some firms are close to zero.[7] Consistent with Francis et al. [2005], to reduce the influence of outliers in the calculation of correlations, we trimmed the sample at 1% and 99% in table 1, panel B and panel C. In all other parts of our study, outlier treatments are not necessary since all analyses are done at a portfolio level. Inspecting insider buying and selling, we find that insiders sell more

[6] EQ measures are calculated on an annual basis. These annual measures are then matched with monthly returns to obtain firm-month observations.

[7] In section 4.3, we discuss robustness checks on this issue.

TABLE 1
Summary Statistics

Panel A: Univariate distribution

	N	Mean	Median	Std. Dev.	Q1	Q3
$EQ1_{it}$	989,530	0.110	0.053	0.341	0.023	0.113
$EQ2_{it}$	989,530	0.089	0.044	0.316	0.017	0.098
$EQ3_{it}$	910,477	0.073	0.041	0.131	0.017	0.088
$EQ4_{it}$	614,981	0.073	0.054	0.073	0.032	0.091
MV_{it}	989,530	1109	81.68	752.1	20.47	385.9
RET_{it}	989,530	0.009	−0.004	0.206	−0.088	0.078
BUY_{it}	51,959	3.690	2.000	9.069	1.000	3.000
$SELL_{it}$	85,075	8.736	2.000	76.93	6.000	1.000

Panel B: Correlation matrix, Pearson above diagonal and Spearman below

	EQ1	EQ2	EQ3	EQ4
EQ1		0.64	0.28	0.32
EQ2	0.75		0.18	0.16
EQ3	0.49	0.53		0.56
EQ4	0.42	0.43	0.53	

Panel C: Mean and median coefficient estimates across 9,966 firms[a]

	Rm – RF		SMB		HML		EQ	
	Mean	Median	Mean	Median	Mean	Median	Mean	Median
EQ1	0.91	0.85	0.68	0.70	0.30	0.25	0.49	0.08
EQ2	0.92	0.83	0.62	0.62	0.32	0.22	0.52	0.09
EQ3	0.88	0.81	0.59	0.60	0.32	0.23	0.51	0.10
EQ4	0.89	0.81	0.52	0.56	0.31	0.22	0.38	0.08

The sample consists of 989,530 observations from 1985 to 2003. EQ1 and EQ2 are earnings quality measures based on the modified Jones model (Dechow, Sloan, and Sweeney [1995]). $EQ1_{it}$ is the absolute value of the abnormal total accruals and $EQ2_{it}$ is the absolute value of the abnormal current accruals. EQ3 and EQ4 are earnings quality measures based on the Dechow and Dichev [2002] model. $EQ3_{it}$ is based on cross-sectional regressions and $EQ4_{it}$ is based on time-series regressions. MV_{it} is firm i's market capitalization in millions at the end of calendar month t. RET_{it} is monthly calendar return for firm i. A firm is classified as BUY if insiders' purchases exceed insiders' sales and SELL if insiders' sales exceed insiders' purchases.

[a]Values in panel C calculated using the following four-factor model:

$$R_{j,t} - R_{f,t} = \alpha_j + \beta_j(R_{m,t} - R_{f,t}) + \delta_j SMB_t + \sigma_j HML_t + \phi_j EQ_t + \varepsilon_{j,t}$$

where: $R_{p,t}$ is portfolio stock return; $R_{f,t}$ is the risk-free rate, measured as the one-month treasury bill rate; $R_{m,t}$ is the market portfolio return, measured using the CRSP value-weighted index; SMB_t and HML_t are the Fama and French [1993] size and market-to-book factor returns, respectively; EQ_t is the hedge return going long in the low earnings quality firms and going short in the high earnings quality firms. The return window is monthly, and factor loadings are estimated using a time-series regression based on 227 months of data, from January, 1985 to November, 2003.

frequently than they buy, and the ratio of the sell to buy transactions is approximately 2 to 1. This is reasonable, because insiders, often compensated in stocks, have incentives to sell their company shares in order to diversify their personal wealth (e.g., Ofek and Yermack [2000]) and consume.

Moving to the correlation matrix, we find the first two earnings quality measures based on the modified Jones model (Dechow, Sloan, and Sweeney [1995]) are highly correlated, with a Spearman rank order (SRO) correlation coefficient at 0.75. However, the two measures based on Dechow and Dichev's [2002] model are not as highly correlated (SRO = 0.53), nor are they highly correlated with EQ1 and EQ2. This suggests that EQ1 and EQ2 are potentially capturing similar information constructs about the firm, but that EQ3 and EQ4 are capturing somewhat distinct constructs.

As a benchmark, similar to Francis et al. [2005], we estimate a four-factor asset-pricing model by adding an earnings quality factor to the Fama and French [1993] three-factor model. In particular, we perform time-series regressions for the following model:

$$R_{j,t} - R_{f,t} = \alpha_j + \beta_j (R_{m,t} - R_{f,t}) + \delta_j SMB_t + \sigma_j HML_t + \phi_j EQ_t + \varepsilon_{j,t} \quad (4)$$

where:

$R_{j,t}$ = firm j's stock return;
$R_{f,t}$ = the risk-free rate, measured as the one-month treasury bill rate;
$R_{m,t}$ = the market portfolio return, measured using the CRSP value-weighted index;
SMB_t and HML_t = the Fama and French [1993] size and market-to-book factor returns, respectively; and
EQ_t = the hedge portfolio return going long in the extreme quintile of low earnings quality firms and going short in the extreme quintile of high earnings quality firms.[8]

In order to make sure that observations pertaining to all variables on the right-hand side of equation (4) are publicly available ex ante, we construct the EQ factor–mimicking portfolio using the realizations of earning quality measures in the preceding fiscal year. We run time-series regressions on each firm using data from January, 1985, to November, 2003.

The mean and median factor loadings are reported in panel C of table 1. The results are broadly consistent with those reported in Francis et al. [2005]. The consistently positive factor loading on the earnings quality factor–mimicking portfolio indicates that, like the Fama and French [1993] factors, asymmetric information risk seems to be priced by the market. However, since positive loadings do not in themselves imply a nonzero risk premium, we move to the next section to investigate whether the asymmetric information risk, as captured by our earnings quality proxies, is priced.

4. *Empirical Analyses*

4.1 EARNINGS QUALITY, INSIDER TRADING, AND ABNORMAL RETURNS

In this section, we establish that the earnings quality measures are effective proxies for information asymmetry that serves as a source of trading advantage exploited by corporate insiders. Because the purpose of this section is

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[8] We obtained the Fama and French [1993] factor-mimicking portfolio returns from Ken French's Web site: http://mba.tuck.dartmouth.edu/pages/faculty/ken.french/data_library.html.

to verify whether the EQ measures proxy for information used by corporate insiders, we calculate the EQ measures contemporaneously based on annual financial statements, and these measures are only partially observable to the investing public from previously released quarterly financial reports. Corporate insiders are hypothesized to know more about these measures than the general public. In section 4.2 and 4.3, where we investigate the systematic component of information risk, we only use publicly available information and construct the factor-mimicking portfolios based on EQ measures of the preceding fiscal years.

To begin, in each year, we partition firms into quintile portfolios based on contemporaneous estimates of earnings quality and construct a hedge portfolio that is long in the lowest quintile of earnings quality firms and short in the highest quintile. Similar to prior studies on insider trading (e.g., Aboody and Lev [2000], Piotroski and Roulstone [2005], and Ke, Huddart, and Petroni [2003]), we use the earnings quality measures as proxies for information available to insiders, but not to outside investors. To the extent that these measures are effective proxies for information asymmetry, we should be able to detect differential insider trading profits across the quintile portfolios, that is, insider trading profits should be negatively correlated with earnings quality measures.

We examine insider trading by further partitioning each quintile portfolio based on earnings quality into those firms in which monthly insider trades are net buys and net sells and estimating abnormal returns using the Fama and French [1993] three-factor model applied to excess portfolio returns:[9]

$$R_{p,t} - R_{f,t} = \alpha_p + \beta_p(R_{m,t} - R_{f,t}) + \delta_p SMB_t + \sigma_p HML_t + \varepsilon_{p,t}. \qquad (5)$$

To measure insider profits, in each month, we compute a firm-specific mean return from the transaction dates of insiders' trades to one day after the filing date of those transactions with the SEC. These firm-specific transaction-to-reporting returns are averages over all the individual insider trades that occurred during the month. We take the average of trades because in many firm-months insiders have multiple trades. In a firm-month in which insiders trade at opposite directions, we still take the average transaction-to-reporting returns, and the average return is classified as a buy (sell) if the net transaction is a buy (sell) in number of shares. Firm-months with no trades are excluded from the analysis.[10] We group insider trading according to the transaction dates rather than filing dates, for example, if two trades occurred in the same calendar month, and one trade is filed within that month and the other trade is filed in the subsequent month, we assign both trades to the former month. This grouping causes some inevitable calendar time mismatching between components of portfolio returns, as well

[9] We classify a firm as a net buyer if the number of shares bought by all insiders during the month exceeds the number of shares sold. The number of transactions, number of shares, and total value of transactions are highly correlated. Replication of the tests with dollar value of transactions yields very similar results to those derived from the number of shares.

[10] We examine the no-trade sample as a benchmark. Results are discussed in section 4.3.

as mismatching between portfolio returns and Fama-French factor returns. We believe potential biases from such mismatching are minimal since the mean interval between transactions and filing dates is only 27 days. Such mismatching may be viewed as a source of noise, and, therefore, likely to reduce the power of tests, implying a bias against finding significant results.

We run time-series regressions for these portfolios based on 227 months of data, from January, 1985 to November, 2003, to estimate the intercept and factor loadings. Because insider buying and selling are measured monthly, in effect our portfolios are rebalanced every month. In addition to the monthly rebalancing due to insider trading, the portfolios are also rebalanced by the end of each fiscal year to reflect changes in earnings quality. Inasmuch as the estimated intercepts for the five-quintile portfolios are generally (with minor exceptions) monotone for all four earnings quality measures, to save space, we only report the results for the extreme portfolios and the hedge portfolio in table 2.

The factor loadings on the three Fama and French [1993] risk factors are qualitatively similar to the benchmark case reported in panel C of table 1. However, they differ in magnitude. Some differences are expected since table 2 is based on portfolio regressions while panel C of table 1 is based on firm-by-firm regressions. The most notable and systematic differences in portfolio loadings occur for the hedge portfolios. These loadings on the three Fama and French [1993] factors are, for the most part, closer to zero, as we would expect given the long and short trading strategy of the hedge portfolios. Our primary interest is with Jensen's alphas (intercepts), α_p, as a measure of excess returns for each combination of insider transaction type, earnings quality portfolio, and earnings quality measure, after controlling for the Fama and French [1993] risk factors.

As expected, the estimated intercepts from time-series regressions of the difference in return between low earnings quality firms and those with high earnings quality on the three systematic factors are all positive when insiders bought shares. For the hedge portfolio, the estimated intercept for EQ1 is 1.342% (t-statistic = 2.16), for EQ2 is 2.408% (t-statistic = 3.00), for EQ3 is 1.282% (t-statistic = 1.88), and for EQ4 is 0.912% (t-statistic = 1.04).[11] While the results for EQ1 and EQ2 are statistically significant at conventional levels, the result for EQ3 is only marginally significant, and the result for EQ4 is insignificant, suggesting the first three measures are better proxies for information advantages experienced by insiders. Another plausible explanation for the insignificant results for EQ4 is the material reduction in sample size that occurs because the estimation of EQ4 requires five years of data.[12]

[11] The t-statistics are conventional measures based on the assumption of zero serial dependence. In all tests here, and later in table 4, we also calculate standard errors using the Newey-West procedure to adjust for potential serial dependence. The results are quite close due to negligible serial dependence in stock returns at the monthly frequency.

[12] Note that the number of observations per regression is considerably reduced from those indicated for our full original sample by partitioning on earnings quality and direction of insider trades.

TABLE 2

Insider Exploitation of Asymmetric Information as Characterized by Earnings Quality Measures

	α	$R_{m,t} - R_{f,t}$	SMB_t	HML_t	Adjusted R^2
EQ1 BUY					
HQ	1.268	0.710	0.669	0.493	0.41
	3.41	*8.86*	*5.88*	*4.90*	
LQ	2.282	1.150	1.139	0.454	0.50
	3.79	*8.99*	*6.52*	*3.05*	
LQ-HQ	1.342	0.249	0.415	−0.148	0.10
	2.16	*1.90*	*2.34*	*−0.99*	
EQ1 SELL					
HQ	1.169	0.543	0.693	0.214	0.47
	4.76	*9.68*	*8.71*	*2.98*	
LQ	0.933	0.657	1.173	0.037	0.51
	2.54	*7.28*	*10.2*	*0.36*	
LQ-HQ	−0.316	0.072	0.422	−0.196	0.16
	−0.99	*0.91*	*4.20*	*−2.22*	
EQ2 BUY					
HQ	1.241	0.622	0.617	0.460	0.35
	3.41	*7.77*	*5.48*	*4.61*	
LQ	3.044	0.992	1.171	0.472	0.29
	3.71	*5.61*	*4.83*	*2.28*	
LQ-HQ	2.408	0.278	0.474	−0.044	0.05
	3.00	*1.63*	*2.01*	*−0.22*	
EQ2 SELL					
HQ	1.134	0.549	0.713	0.208	0.47
	4.60	*9.74*	*8.92*	*2.89*	
LQ	1.033	0.675	1.084	−0.050	0.52
	2.98	*7.94*	*9.89*	*−0.52*	
LQ-HQ	−0.180	0.098	0.381	−0.253	0.15
	−0.55	*1.21*	*3.65*	*−2.75*	
EQ3 BUY					
HQ	1.741	0.570	0.511	0.458	0.23
	3.82	*5.69*	*3.66*	*3.74*	
LQ	2.973	0.816	1.091	0.512	0.35
	4.26	*5.76*	*5.47*	*2.94*	
LQ-HQ	1.282	0.194	0.569	0.057	0.06
	1.88	*1.24*	*2.61*	*0.31*	
EQ3 SELL					
HQ	1.116	0.478	0.614	0.237	0.43
	4.86	*9.13*	*8.30*	*3.52*	
LQ	1.349	0.680	1.049	−0.023	0.43
	3.39	*7.40*	*8.18*	*−0.20*	
LQ-HQ	−0.108	0.250	0.458	−0.255	0.21
	−0.31	*3.06*	*4.08*	*−2.53*	
EQ4 BUY					
HQ	0.644	0.464	0.360	0.472	0.22
	1.47	*4.92*	*2.71*	*4.26*	
LQ	1.948	1.304	1.454	0.419	0.55
	2.85	*8.80*	*7.15*	*2.53*	
LQ-HQ	0.912	0.746	1.149	−0.097	0.32
	1.04	*3.88*	*4.55*	*−0.48*	

TABLE 2 — *Continued*

	α	$R_{m,t} - R_{f,t}$	SMB_t	HML_t	Adjusted R^2
EQ4 SELL					
HQ	1.098	0.568	0.314	0.303	0.42
	5.22	*11.2*	*4.64*	*4.99*	
LQ	0.952	0.642	1.229	0.093	0.43
	2.14	*5.85*	*8.83*	*0.78*	
LQ-HQ	−0.218	0.193	0.963	−0.182	0.32
	−0.50	*1.81*	*7.13*	*−1.58*	

Time-series regression results are obtained using the Fama-French three-factor model:

$$R_{p,t} - R_{f,t} = \alpha_p + \beta_p\left(R_{m,t} - R_{f,t}\right) + \delta_p SMB_t + \sigma_p HML_t + \varepsilon_{j,t}$$

where: $R_{p,t}$ is the portfolio stock return, the return interval is between the transaction date and one day after the SEC filing date—27 days on average; $R_{f,t}$ is the risk-free rate, measured as the one-month treasury bill rate; $R_{m,t}$ is the market portfolio return, measured using the CRSP value-weighted index; SMB_t and HML_t are the Fama and French [1993] size and market-to-book factor returns, respectively. The return window is monthly, and factor loadings are estimated using a time-series regression based on 227 months of data, from January, 1985 to November, 2003. *t*-statistics are under the coefficient estimates and in *italics*. We report results for insider buying and selling for three portfolios based on four measures of earnings quality: the quintile of firms with highest earnings quality (HQ), the quintile of firms with lowest earnings quality (LQ), and a hedge portfolio where we buy LQ quintile and sell HQ quintile. Insiders are classified as BUY if they have net purchases during the month and SELL if they have net sales during the month. Only months where at least 15 firms are present are included. See section 3 for the definitions of the four earnings quality measures.

Moving to insider sell transactions, we observe much weaker results. For example, regardless of the earnings quality measure used, there are no discernable negative abnormal returns for the high or low earnings quality portfolios. Results for hedge portfolios are slightly stronger; while the abnormal returns for the hedge portfolios that are long in low quality firms and short in high quality firms are negative for all EQ measures, none of the estimates are statistically significant. The estimated intercepts for the hedge portfolios are −0.316% for EQ1 (*t*-statistic = −0.99), −0.180% for EQ2 (*t*-statistic = −0.55), −0.108 for EQ3 (*t*-statistic = −0.31), and −0.218 for EQ4 (*t*-statistic = −0.50).

To summarize results in table 2, we find evidence (primarily from buy transactions) that insiders trade more profitably in low earnings quality firms than in high earnings quality firms. In addition to their statistical significance, the gains to insider buys are economically significant. The annualized gains to insiders in buy transactions are 16.10% for EQ1, 28.89% for EQ2, 15.38% for EQ3, and 10.94% for EQ4. The weaker results for sell transactions are not unexpected. Because corporate insiders usually have disproportional financial capital and human capital invested in their firms, they have strong incentives to sell their company stocks and diversify. Moreover, to the extent their compensation is in the form of stock, they have incentive to sell in order to consume. The selling driven by diversification and consumption motives plausibly overshadows the selling driven by private information, resulting in weaker results documented in table 2.

Having established that earnings quality measures are effective proxies for information asymmetries exploited by corporate insiders, in section 4.2 we examine whether information asymmetries have pricing implications. The fact that insiders exploit their information advantages in a multi-asset setting does not guarantee any pricing implications, because the price effects could be fully diversifiable (e.g., Hughes, Liu, and Liu [2004]). In order to assess pricing implications, we focus on the systematic component of asymmetric information risk and investigate whether the firm's exposure to the systematic component of asymmetric information risk, as measured by the loadings on the earnings quality factor–mimicking portfolio estimated using the four-factor model introduced in section 3, is priced, that is, reflected in expected returns. Finding a priced component to asymmetric information risk, in section 4.3 we investigate whether insiders exploit information advantages reflected in that component by earning abnormal returns after controlling for the Fama and French [1993] risk factors as well as for the newly constructed asymmetric information risk factor. We hypothesize that insiders' abnormal trading gains are positively correlated with the firm's exposure to the systematic component of asymmetric information risk.

4.2 PREMIUM ON ASYMMETRIC INFORMATION RISK

In order to measure a firm's exposure to the systematic component of asymmetric information risk, we again add an asymmetric information risk factor to the Fama and French [1993] three-factor model mimicked by a portfolio that is long in the quintile of stocks with the lowest earnings quality and short in the quintile with the highest quality as in equation (4). To obtain an estimate for each firm-month, we use the past 36 months of data to estimate factor loadings through monthly time-series regressions. As previously noted, the estimated regression coefficient, or loading, on the earnings quality factor–mimicking portfolio, ϕ_j, measures the firm's exposure to the systematic component of asymmetric information risk. This measure is different from measures based on earnings quality per se in that the nonpriced idiosyncratic component is filtered out.

To verify that the systematic component of asymmetric information risk is priced, we form quintile portfolios according to a firm's exposure to that component (factor loading) in each month, and run the Fama and French [1993] three-factor model to estimate Jensen's alphas. Because factor loadings are estimated every month, the portfolios are in effect rebalanced every month. We report results for measures of exposure based on each of the four earnings quality proxies: the quintile of firms with the highest $\phi_j\,(H\phi)$, the quintile of firms with the lowest $\phi_j\,(L\phi)$, and a hedge portfolio where we sell firms in the $L\phi$ quintile and buy those in the $H\phi$ quintile. Our results are contained in table 3.

The results in table 3 are broadly consistent with the notion that the systematic component of information risk commands a positive risk premium. While the low-exposure firms have alphas not significantly different from zero, the high-exposure firms all have statistically significant positive alphas. As a result, the hedge portfolio that is long in $H\phi$ firms and short in $L\phi$ firms has positive alphas across the four EQ measures, implying a positive risk premium on the EQ factor in each case.

TABLE 3

Premiums on Exposure to the Systematic Component of Asymmetric Information Risk

		α	$R_{m,t} - R_{f,t}$	SMB_t	HML_t	Adjusted R^2
EQ1	$L\phi$	0.159	0.893	0.887	0.053	0.64
		0.58	*14.5*	*10.0*	*0.66*	
	$H\phi$	1.190	1.150	1.744	0.358	0.50
		2.24	*9.51*	*10.1*	*2.30*	
	$H\phi - L\phi$	1.031	0.257	0.857	0.306	0.05
		1.37	*1.50*	*3.50*	*1.38*	
EQ2	$L\phi$	0.182	0.914	0.824	0.126	0.63
		0.67	*14.9*	*9.38*	*1.59*	
	$H\phi$	1.197	1.142	1.756	0.345	0.50
		2.22	*9.32*	*10.0*	*2.18*	
	$H\phi - L\phi$	1.016	0.228	0.932	0.218	0.06
		1.33	*1.31*	*3.76*	*0.97*	
EQ3	$L\phi$	0.189	0.892	0.795	0.156	0.65
		0.76	*15.7*	*9.80*	*2.13*	
	$H\phi$	1.181	1.154	1.792	0.350	0.49
		2.15	*9.22*	*10.0*	*2.17*	
	$H\phi - L\phi$	0.992	0.261	0.996	0.194	0.07
		1.30	*1.50*	*4.01*	*0.86*	
EQ4	$L\phi$	0.133	0.881	0.736	0.239	0.65
		0.56	*16.4*	*9.60*	*3.45*	
	$H\phi$	1.311	1.105	1.837	0.246	0.49
		2.34	*8.67*	*10.1*	*1.50*	
	$H\phi - L\phi$	1.178	0.224	1.102	0.007	0.09
		1.53	*1.28*	*4.40*	*0.03*	

A firm's exposure to asymmetric information risk is estimated by running the following four-factor model:

$$R_{j,t} - R_{f,t} = \alpha_j + \beta_j \left(R_{m,t} - R_{f,t}\right) + \delta_j SMB_t + \sigma_j HML_t + \phi_j EQ_t + \varepsilon_{j,t}$$

where: $R_{p,t}$ is the portfolio stock return; $R_{f,t}$ is the risk-free rate, measured as the one-month treasury bill rate; $R_{m,t}$ is the market portfolio return, measured using the CRSP value-weighted index; SMB_t and HML_t are the Fama and French [1993] size and market-to-book factor returns, respectively; and EQ_t is the hedge return going long in the low earnings quality firms and going short in the high earnings quality firms. The return window is monthly, and factor loadings are estimated for every 36 months using a time-series regression. The regression coefficient, or factor loading, ϕ_j, measures a firm's exposure to asymmetric information risk. We then sort firms into quintiles using the estimated factor loadings and report results for time-series regressions using the Fama and French [1993] three-factor model:

$$R_{p,t} - R_{f,t} = \alpha_p + \beta_p \left(R_{m,t} - R_{f,t}\right) + \delta_p SMB_t + \sigma_p HML_t + \varepsilon_{j,t}$$

where the return interval is monthly. The variables in the three-factor model are defined the same as in the four-factor model above. Factor loadings are estimated using a time-series regression based on 227 months of data, from January 1985 to November, 2003. *t*-statistics are under the coefficient estimates and in *italics*.

The estimates are all economically significant, with values of 1.031% for EQ1, 1.016% for EQ2, 0.992% for EQ3, and 1.178% for EQ4. The statistical significance for the hedge portfolio is less impressive—all *t*-statistic estimates for the hedge portfolio are below two. Combining the strong results for the $H\phi$ firms and the weaker but consistent results for the hedge portfolios, we conclude that it is likely that the noise contained in the observations, primarily from the $L\phi$ firms,[13] reduced the significance level for the hedge portfolio results.

This finding also illustrates the importance of examining the premium on the information risk factor explicitly. Francis et al. [2005] found that

[13] Estimates of alpha have larger standard errors for $L\phi$ firms than for $H\phi$ firms.

firms have highly significant exposures to the information risk factors and thus concluded that the risk factors are priced. Our results show that the evidence is in fact weaker than one might have surmised from factor loading estimates alone. Of course, this difference in conclusion is only at a quantitative level. In the next section, we show that the alpha spread between $H\phi$ and $L\phi$ firms is positively correlated with insider trading profits, a finding that further supports the notion that the systematic component of the asymmetric information risk factor is priced.

4.3 SYSTEMATIC ASYMMETRIC INFORMATION RISK, INSIDER TRADING AND ABNORMAL RETURNS

As indicated earlier, if privately informed trading lies behind the pricing of asymmetric information risk, then insiders should earn abnormal trading profits after controlling for all risk factors, including the systematic component of asymmetric information risk. Furthermore, we expect insiders' trading profits to be positively correlated with firms' loadings on the earnings quality factor-mimicking portfolio.[14] To demonstrate these points, we now turn to table 4.

Table 4 is similar to table 2 except that just the systematic (priced) component of the firm's asymmetric information risk is the focus of our investigation. We estimate a firm's exposure to the systematic component of asymmetric information risk in the same way as in table 3. For each month, we first sort firms into buy and sell portfolios by checking whether the insiders are net buyers or net sellers. Then, within each buy or sell portfolio, we further partition firms into five quintiles according to their exposure to the asymmetric information risk factor, as measured by loadings on that factor. Finally, we run the four-factor model given by equation (4) for each portfolio, generating regression intercepts (Jensen's alphas) that now represent abnormal trading profits to insiders after controlling for all four risk factors. We report results for the bottom and top quintiles, as well as for a hedge portfolio that is long in the top quintile and short in the bottom quintile.

We find strong evidence that insiders profit by exploiting the systematic portion of their private information for buy transactions. Judging by the hedge portfolio results, all estimates of abnormal returns (Jensen's alphas) are highly statistically significant with the exception of the result based on EQ4. Their significance is similarly strong in the economic sense. For example, insiders make 3.344%, 1.827%, 1.676%, and 1.339% monthly abnormal profits (after controlling for known risk factors) for hedge portfolio returns based on EQ1, EQ2, EQ3, and EQ4 measures of earnings quality, respectively. Comparing this set of results with those reported in table 2,

[14] We note that there is considerable noise contained in insider trading data. In theory, a thorough test would require consideration of additional phenomena such as other motives for insider trading and insiders' private portfolio holdings of the firm's stock and other assets, etc. Unfortunately, data on these additional issues are generally not available.

TABLE 4
Insider Exploitation of Firms' Exposure to the Systematic (Priced) Component of Asymmetric Information Risk

	α	$R_{m,t} - R_{f,t}$	SMB_t	HML_t	EQ_t	Adjusted R^2
EQ1 BUY						
$L\phi$	2.008	0.825	1.164	0.169	−0.708	0.36
	4.06	*7.48*	*6.54*	*1.26*	*−4.36*	
$H\phi$	5.176	0.698	−0.016	0.994	2.284	0.53
	7.31	*3.96*	*−0.06*	*5.11*	*9.59*	
$H\phi - L\phi$	3.344	−0.230	−1.213	0.836	3.089	0.47
	3.82	*−1.07*	*−3.87*	*3.62*	*10.8*	
EQ1 SELL						
$L\phi$	1.612	0.627	1.100	−0.151	−0.497	0.46
	5.16	*8.55*	*9.00*	*−1.62*	*−4.41*	
$H\phi$	1.778	0.414	0.219	0.334	1.700	0.65
	4.81	*4.62*	*1.51*	*3.05*	*12.8*	
$H\phi - L\phi$	−0.093	−0.141	−0.853	0.496	2.179	0.48
	−0.07	*−1.30*	*−4.83*	*3.74*	*13.5*	
EQ2 BUY						
$L\phi$	2.335	0.830	1.232	0.243	−0.806	0.35
	4.59	*7.35*	*6.46*	*1.77*	*−4.83*	
$H\phi$	4.311	0.893	−0.050	1.111	2.253	0.63
	6.97	*5.98*	*−0.22*	*6.65*	*11.2*	
$H\phi - L\phi$	1.827	−0.235	−1.446	0.839	3.204	0.54
	2.26	*−1.18*	*−4.78*	*3.91*	*12.2*	
EQ2 SELL						
$L\phi$	1.512	0.626	1.104	−0.114	−0.515	0.45
	4.96	*8.82*	*8.89*	*−1.25*	*−4.68*	
$H\phi$	1.559	0.497	0.265	0.435	1.662	0.66
	4.20	*5.28*	*1.72*	*3.96*	*12.5*	
$H\phi - L\phi$	−0.019	−0.173	−0.937	0.551	2.216	0.51
	−0.04	*−1.52*	*−5.00*	*4.17*	*13.8*	
EQ3 BUY						
$L\phi$	2.758	0.786	1.399	0.359	−0.793	0.29
	4.84	*6.05*	*6.42*	*2.34*	*−4.52*	
$H\phi$	4.401	0.655	−0.202	0.857	2.194	0.62
	6.97	*4.35*	*−0.83*	*5.03*	*11.3*	
$H\phi - L\phi$	1.676	−0.386	−1.659	0.442	3.115	0.58
	2.15	*−2.12*	*−5.72*	*2.21*	*13.6*	
EQ3 SELL						
$L\phi$	1.912	0.563	1.035	−0.114	−0.418	0.38
	5.87	*7.17*	*7.65*	*−1.19*	*−3.84*	
$H\phi$	1.348	0.429	−0.032	0.370	1.698	0.66
	3.56	*4.64*	*−0.20*	*3.33*	*13.5*	
$H\phi - L\phi$	−0.656	−0.118	−1.060	0.499	2.117	0.53
	−1.48	*−1.08*	*−5.77*	*3.86*	*14.3*	
EQ4 BUY						
$L\phi$	3.528	0.809	1.283	0.335	−0.475	0.20
	6.21	*6.06*	*5.30*	*1.99*	*−3.86*	
$H\phi$	4.867	0.581	−0.413	0.723	1.372	0.46
	7.65	*3.89*	*−1.53*	*3.84*	*9.94*	
$H\phi - L\phi$	1.339	−0.228	−1.696	0.388	1.847	0.32
	1.56	*−1.13*	*−4.65*	*1.53*	*9.93*	

TABLE 4 — *Continued*

	α	$R_{m,t} - R_{f,t}$	SMB_t	HML_t	EQ_t	Adjusted R^2
EQ4 SELL						
$L\phi$	1.396	0.590	1.013	0.009	−0.240	0.39
	4.56	*8.21*	*7.77*	*0.10*	*−3.62*	
$H\phi$	1.049	0.563	−0.128	0.304	1.132	0.65
	2.60	*5.94*	*−0.74*	*2.55*	*12.9*	
$H\phi - L\phi$	−0.349	−0.028	−1.140	0.295	1.373	0.49
	−0.75	*−0.25*	*−5.74*	*2.14*	*13.6*	

A firm's exposure to asymmetric information risk is estimated by the following four-factor model:

$$R_{j,t} - R_{f,t} = \alpha_j + \beta_j \left(R_{m,t} - R_{f,t}\right) + \delta_j SMB_t + \sigma_j HML_t + \phi_j EQ_t + \varepsilon_{j,t}$$

where: $R_{p,t}$ is the portfolio stock return; $R_{f,t}$ is the risk-free rate, measured as the one-month treasury bill rate; $R_{m,t}$ is the market portfolio return, measured using the CRSP value-weighted index; SMB_t and HML_t are the Fama and French [1993] size and market-to-book factor returns, respectively; and EQ_t is the hedge return going long in the low earnings quality firms and going short in the high earnings quality firms. The return window is monthly and factor loadings are estimated for every 36 months using time-series regressions. The regression coefficient, or factor loading, ϕ_j, measures a firm's exposure to asymmetric information risk. We then sort firms into quintiles using the estimated factor loadings and report results for insider buying and selling based on each of four measures of exposure for three portfolios: the quintile of firms with the lowest ϕ_j ($L\phi$), the quintile of firms with the highest ϕ_j ($H\phi$), and a hedge portfolio where we buy $H\phi$ quintile and sell $L\phi$ quintile. Insiders are classified as BUY if they have net purchases during the month and SELL if they have net sales during the month. We apply the Fama and French [1993] three-factor model to each portfolio:

$$R_{p,t} - R_{f,t} = \alpha_p + \beta_p \left(R_{m,t} - R_{f,t}\right) + \delta_p SMB_t + \sigma_p HML_t + \phi_j EQ_t + \varepsilon_{j,t}$$

where the return interval is between the transaction date and one day after the SEC filing date—27 days on average. The regression variables are defined the same as in the four-factor model above. Factor loadings are estimated using a time-series regression based on 227 months of data, from January, 1985 to November, 2003. t-statistics are under the coefficient estimates and in *italics*.

we note that both the abnormal returns and the significance levels are generally higher in table 4. The fact that the results are stronger in table 4 suggests that insider trading is associated with the systematic component of the asymmetric information risk, and parsing out the systematic component via a factor model is important empirically. Results are weaker for insider sell transactions. Although all four EQ measures produce EQ factor loadings with the correct signs and significance at conventional levels, none of the Jensen's alphas are significant. These results suggest that insiders' buy transactions are more likely to have been based on private information than their sell transactions.

The above results are based on a sample in which there are insider trades within the firm-months. However, for the majority of the firm-months in our data set, there are no insider trades. To benchmark our analysis, we repeat the same analysis as in table 4 on the sample with no insider trades. We expect the Jensen's alphas to be close to zero for this subsample since the risk factors we control for include the EQ factor. The results confirm our expectation; the point estimates for Jensen's alphas are very close to zero, with estimates of 0.016%, 0.012%, 0.006%, and 0.011% for EQ1, EQ2, EQ3, and EQ4, respectively.

Another robustness issue follows from the skewness of earnings quality measures reported in table 1. Inasmuch as the earnings management literature generally concludes that all abnormal accruals measures are noisy, it is reasonable to question whether our results are an artifact of outliers. To alleviate concerns, we construct new samples, by excluding firms with total assets below $50 million or trimming the sample at the top 5% and bottom 5% of EQ, and repeat the analyses in table 3 and table 4. We find qualitatively similar results in all these sensitivity tests.

In sum, the results from our tests of an association between a firm's exposure to the systematic component of asymmetric information risk as measured by loadings on an earnings quality factor–mimicking portfolio and gains to insider trading provide supporting evidence that asymmetric information risk premiums are a consequence of insider trading. The weakness of the findings on the sell side of insider trading is consistent with the greater likelihood of mixed motives for insiders to sell.

4.4 TRADING FREQUENCIES

Establishing a relationship between abnormal returns to insiders from the date they trade to the date those trades are made public and a firm's exposure to asymmetric information risk is of first-order importance, in the sense that failing to find such a relationship precludes further inquiry on the prospect that insider trading is driving the risk premium on our earnings quality factor. Having found evidence of a relationship, it remains to be shown whether the higher abnormal returns associated with high exposure to the earnings quality factor are accompanied by a higher frequency of insider trades, implying that insiders are, indeed, exploiting their information advantage to realize greater profits. Accordingly, we extend our analysis to incorporate relative trading frequencies.

Specifically, we examine the connection between insider trading frequencies and the firms' exposure to the earnings quality factor by comparing numbers of monthly insider trades for portfolios sorted on firms' exposure to the EQ factor. Table 5 provides strong evidence that insiders trade more frequently for high-exposure firms ($H\phi$) than for low-exposure firms ($L\phi$): the differences are highly significant for all four EQ measures as well as for both buy and sell transactions. Because of the concern that statistical significance might be overstated due to cross-sectional correlation, we also apply an alternative method similar to that of Fama and Macbeth [1973] to measure the sample statistics and associated significance levels: instead of measuring the global pooled statistics, we first measure the means and medians for each month, then treat these numbers as a time series and measure the mean of the mean series and the median of the median series. Significance levels are then measured based on these two time series of data. We employ both a parametric t-test of differences in means, and a nonparametric Wilcoxon rank-sum test. The results are largely consistent with the pooled results, but weaker in significance levels, as one would expect from such a test design.

TABLE 5

Insider Trading Frequencies

	BUY	SELL
EQ1		
$L\phi$	3.576	9.800
	2.000	*2.000*
$H\phi$	3.690	13.496
	2.000	*2.000*
p-value	0.001	0.001
	0.001	*0.001*
EQ2		
$L\phi$	3.527	9.752
	2.000	*2.000*
$H\phi$	3.763	13.33
	2.000	*3.000*
p-value	0.001	0.001
	0.001	*0.001*
EQ3		
$L\phi$	3.481	3.838
	2.000	*2.000*
$H\phi$	9.497	10.41
	2.000	*2.000*
p-value	0.001	0.001
	0.001	*0.001*
EQ4		
$L\phi$	3.526	3.815
	2.000	*2.000*
$H\phi$	9.240	9.311
	2.000	*2.000*
p-value	0.001	0.001
	0.001	*0.001*

A firm's exposure to asymmetric information risk is estimated by the following four-factor model:

$$R_{j,t} - R_{f,t} = \alpha_j + \beta_j \left(R_{m,t} - R_{f,t}\right) + \delta_j SMB_t + \sigma_j HML_t + \phi_j EQ_t + \varepsilon_{j,t}$$

where: $R_{p,t}$ is the portfolio stock return; $R_{f,t}$ is the risk-free rate, measured as the one-month treasury bill rate; $R_{m,t}$ is the market portfolio return, measured using the CRSP value-weighted index; SMB_t and HML_t are the Fama and French [1993] size and market-to-book factor returns, respectively; and EQ_t is the hedge return going long in the low earnings quality firms and going short in the high earnings quality firms. The return window is monthly and factor loadings are estimated for every 36 months using time-series regressions. The regression coefficient, or factor loading, ϕ_j, measures a firm's exposure to asymmetric information risk. We then sort firms into quintiles using the estimated factor loadings and report results for monthly insider buying and selling frequencies based on each of four measures of exposure for two extreme portfolios: the quintile of firms with the lowest $\phi_j (L\phi)$ and the quintile of firms with the highest $\phi_j (H\phi)$. Insiders are classified as BUY if they have net purchases during the month and SELL if they have net sales during the month. For each portfolio, we report the mean frequency and median frequency (in *italics*), as well as p-values for the significance tests for the mean and median, with the null hypothesis being that there is no difference between the $L\phi$ portfolio and the $H\phi$ portfolio.

5. *Concluding Remarks*

The research questions addressed in this paper are: (1) whether the systematic component of earnings quality (a proxy for information asymmetry) is priced, and (2) whether privately informed traders earn greater profits when trading stocks with higher exposure to an earnings quality risk factor. Our results are consistent with an affirmative answer to both questions.

Measures of earnings quality employed in this paper include two based on the modified Jones model (Dechow, Sloan, and Sweeney [1995]) and two based on the Dechow and Dichev [2002] model. In the context of a multi-factor model of asset pricing, we use these proxies to form earnings quality factor-mimicking portfolios that, in turn, allow us to filter out the nonpriced component of earnings quality. Exposure to the priced component of earnings quality is measured by loadings on the earnings quality factor in time-series regressions of a Fama-French three-factor model augmented by the earnings quality factor. We find that asymmetric information risk premiums are significantly positive for firms with high factor loadings, insignificantly different from zero for firms with low factor loadings, and positive but not significant for the hedge portfolio that is long in the high-exposure firms and short in the low-exposure firms. These results suggest that the systematic component of earnings quality is priced, but the evidence is relatively weak, a result not available from portfolio loadings evidence as reported in prior literature (Francis et al. [2005]).

Partitioning sample firms by the direction of insider (corporate officers, directors, and owners of 10% or more of any equity class of securities) trades subsequently reported to the SEC, as well as by loadings on the earnings quality factor, we find strong evidence that insiders trade more profitably in firms with high factor loadings. The combination of the findings is supportive of the conclusions that asymmetric information risk as proxied by earnings quality is priced, and insider trading is an important element in establishing this effect.

Given the strong results on insider trading, the relatively weak result for the pricing of earnings quality is intriguing. One plausible explanation is that the latter tests have low power due to a large amount of noise in the system. Another equally plausible explanation is that diversification eliminates some pricing effects of asymmetric information risk. The results of Hughes, Liu, and Liu [2005] suggest that the cross-sectional effect of asymmetric information on cost of capital may be fully diversified away in a pure exchange economy with a large number of assets. It seems intuitive that, in large economies, information must directly affect asset payoffs through production and investment decisions in order for asymmetric information to enter as a separate factor in the determination of cost of capital. An explicit model to incorporate such ideas is not available in the literature and warrants future research.

Reversing the perspective, our results also suggest that there is some truth to the notion, popular both in the accounting literature and on Wall Street, that earnings numbers have different "qualities." The fact that unsigned abnormal accruals as proxies for earnings quality can generate our results implies that accruals are an important characteristic of a firm's information environment, in particular, an indication of asymmetric information risk emanating from privately informed trading.

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