# Empirical Investigation of the Link Between Systematic Risk and Precision of Information

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*Abstract:* This paper investigates the link between systematic risk and other firm's characteristics around earnings and dividend announcements. The analysis uses a sample of 212 securities drawn from the London Stock Exchange for the period January 1990 to December 1998.

The analytical work of Kim and Verrecchia (henceforth KV, 1991a, 1997) predict that the variance of price change is increasing in the precision of the announcement but decreasing in the amount of pre-announcement information. The paper shows analytically that the implication of KV's predictions on volatility carry over to systematic risk of securities as well. Their predictions are tested using empirical surrogates for the quality of pre-announcement information and the precision of news releases based on fundamental values of sample firms. After controlling for contemporaneous correlation, changes in the degree of operating and financial leverage, and other firm characteristics that are thought to be associated with changes in systematic risk, we find that the level of systematic risk around earnings and dividends announcements is negatively related to the amount of pre-announcement information. Specifically, all else being equal, we document relatively small shifts in beta around anticipated announcements of large firms. The opposite is true for small firms.

We also document a positive association between the proxy for the precision of the announced news and the level of systematic risk around earnings and dividends release dates. Overall, the evidence presented in this paper is consistent with the predictions of KV.

## INTRODUCTION

KV's analytical work predicts that the variance of stock returns around information-intensive periods will be directly proportional to the quality of news disclosed and negatively related to the quality of the pre-announcement information. It is quite possible that the insightful intuition and predictions of KV's work extends to systematic risk (betas) of securities as well. Indeed, if KV's predictions carry over to systematic risk of securities, we would expect firms with a lot of pre-announcement information to experience smaller shifts in beta around public news releases than firms with relatively little pre-announcement information. Further, firms that release precise (imprecise) public news may be expected to undergo large (small) shifts in beta around news disclosure dates.

Investigation of the link between systematic risk of securities and the quality of information will certainly enrich our understanding of risk-return trade-off around information intensive periods.This knowledge, in turn, will undoubtedly be useful in the area of asset pricing and investors trading behaviour around public news. In this paper we combine KV's (1991a, 1997) insightful intuition and the assumption of the single index model to formally demonstrate that on average, the quality of information disclosed is related to the level of systematic risk around information-intensive periods. Specifically, it is shown that systematic risk of a typical security is predicted to be positively (negatively) related to the quality of information released (the quality of pre-announcement information).

Of course, systematic risk reaction to public news releases is a complex phenomenon that can not be simply explained by the precision of the news released and the quality of the pre-announcement information. A comprehensive study of systematic risk changes around news release dates will also consider, *inter alia*, factors such as changes in earnings (see for example, Ball and Kothari 1991), changes in the degree of operating and financial leverage (see for example, Lev 1974; Mandelker and Rhee, 1984) and other

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factors such as the pre-announcement levels of market risk (see for example, Blume 1971, 1975). The remainder of this paper is organised as follows. Section 1 presents a brief summary of some literature on news releases that has direct implication and relevance to the current paper. Section 2 describes data and sample selection procedure, while section 3 presents the methodology and the empirical tests to be performed. Empirical results are discussed in section 4. Section 5 examines the sensitivity of empirical results to time-variation in systematic risk. The last section offers concluding remarks.

## RELATED WORK AND RESEARCH QUESTIONS

The analytical model proposed by KV (1991a) and the extensions covered in KV (1991b, 1994, 1997) has implications for both the trading volume reactions and risk shifts around news release dates. However, the trading volume reaction to public news releases is not analysed in this paper because Atiase and Bamber (1994) and Bamber et al. (1997) have carefully examined the volume aspect. While Pope and Inyangete (1992) empirically examined differential variability of securities' returns around earnings and dividends announcements between small and large UK firms driven by the amount of pre-disclosure information. In contrast, there has been no empirical work testing the association between systematic risk and quality (precision) of a public announcement based on a theoretical model.[1] Indeed, the current investigation furthers this line of research by examining how robust the KV's (1991a, 1997) predictions are with respect to different metrics of risk.

Most of the earlier work on systematic risk has for the most part explicitly or implicitly relied on the rather implausible assumption that beta of individual securities is constant over time, except possibly during information intensive periods [see for example, Kalay and Lowestein (1985), Ball and Kothari (1991)]. This practice is not consistent with the evidence of time-variation in betas documented by, among others, Sunder (1980), Schwert and Seguin (1990), Kim (1993), Koutmos et al. (1994), Pope and Warrington (1996), and Faff *et al.* (2000). We explore this line of research by allowing for temporal instability of beta estimates in the empirical analysis of KV's (1991a) predictions.

Another objective of this paper is to examine the robustness of KV's (1991a, 1997) predictions with respect to other variables that are capable of affecting systematic risk. Specifically, our empirical design allows for differences in operating and financial leverage across stocks. Theoretically, systematic risk may be affected by changes in operating and financial leverage.[2] To the extent that risk originating from operating and financial leverage can not be diversified away, changes in these ratios surrounding earnings and dividends announcements dates will be positively related to changes in systematic risk.

Fourth, a natural by-product of the current research is to give light to some anomalous issues regarding temporary beta shifts around news events. For instance, while studies like Kalay and Lowestein (1985), Ball and Kothari (1991), appear to suggest that systematic risk is likely to increase at the time of company specific public news releases, this evidence of upward pulse in betas does not appear to be conclusive. Eades, Hess, and Kim (1985), for example, found no evidence of beta shifts around dividend announcements.

## DATA AND SAMPLE SELECTION

FTSE-350 constituents stocks for the period ending 31st December 1998 are initially chosen for analysis. These are relatively large, actively traded and closely watched stocks. Several investment services predict both the date and content of these firms' forthcoming earnings and dividend announcements. The main criterion

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[1] The notable exception is Hsieh *et al.* (1999), who investigate the link between systematic risk of a sample of US stocks and their proxy for inverse precision. Unfortunately, their work lacks analytical explanation or intuition that justifies a negative relationship between *beta* and their proxy for the inverse precision. Such theoretical justification appears in Mohamed (2001), Mohamed and Yadav (2002), and Hillier *et al.* (2003)

[2] See for example, Hamada (1972), Lev (1974), and Mandelker and Rhee (1984).

used to select a company in the final sample is data availability. The final sample consists of 212 firms from 10 broad industries with a total of 4617 earnings and dividend announcements. The selection process involves a trade-off between the benefits of mitigating nontrading problems and the costs of having a final sample that is not representative of the market as a whole. To the extent that nontrading is the most severe problem, the sample selection criterion is justified.[3]

Adjusted closing middle market prices obtained from Data Stream Service covering the same period are used to generate continuous return series. Returns calculated from closing middle market prices contain no noise due to the bid-ask bounce [see for example, Jones et al. (1994), Abhyankar et al. (1997), Mohamed (2002)]. The returns for holidays for each of the sample firm are removed from estimation process, and this procedure reduces the sample size to 2277 daily observations for each of the sample firm and the FT All Share Index.

The dataset used to create the other firm specific information proxies and leverage ratios include market value of outstanding shares, earnings, dividends, preference capital, borrowings due within one year, longterm-borrowing, equity capital and reserves, the net fixed assets, and the total assets of a firm. This information was collected from Datastream. The detailed procedures on how to estimate information proxies appear under relevant sections.

## METHODOLOGY AND EMPIRICAL TESTS

### Specification of the Return Generating Process

Following Scholes and Williams (1977), Dimson (1979),[4] Cohen *et al* (1983, 1986), and Andersen (1989) the return generating process is first modelled using the aggregated coefficients method[5] with one lag and one lead:

$$R_{jt}=\alpha_{0j}+\sum_{k=-1}^{1}\beta_{jk}R_{m(t+k)}+\alpha_{1j}D_A+\beta_{Aj}(D_A*R_{mt})+\varepsilon_{jt} \qquad (1$$

where Rjt is the rate of return to a security *j* in period *t*, $R_{mt}$ is the rate of return to the market portfolio in period *t* (FTSE All Share Index is used as a proxy for the market portfolio), $D_A$ is an announcement dummy which assumes the value of one on the earnings or dividend announcement date and on the preceding and post-announcement day and has the value of zero otherwise,[6] $\alpha_{0j}, \alpha_{1j}$, are regression parameters, $\beta_{Aj}$ is an announcement beta, and $\beta_{jk}$ for k=-1, 0, and 1 are the observed lag, intertemporal and lead security betas, and the residuals series, $\{\varepsilon_{jt}\}$, measures the extent to which the actual return to security j in period t differs from the expected rate of return.

The inclusion of lagged, contemporaneous and lead terms of $R_{mt}$ in (1) is expected to "damp out" any spurious auto-and cross-correlation that might be caused by lagged price adjustments[7] and nonsynchronous trading. It is important not to ignore such a possibility because the presence of such effects may mask the true underlying process governing returns behaviour. In particular, the main "cost" of nonsynchroneity and poor price adjustment lies in the biasdness they place on the market model betas, the correlation structure, and volatility of returns. For instance, in the presence of non-synchronous trading, the ordinary least

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[3] See Mohamed (2003) for detailed characteristic of the data used.

[4] It is important to note that Dimson's (1979) estimator is not specified correctly. Fowler and Rorke (1983) show that Dimson's estimator is not consistent with that of Scholes and Williams (1977). The corrected version appears in Fowler and Rorke's paper.

[5] This is essentially a multiple regression using lead, lagged and dummy variables as additional regressors

[6] The day -1 is chosen to take into account the possibility that announcement effects (leakage of information) may begin to influence prices before the formal announcement. It is possible for this effect to begin several days before the formal announcement but for simplicity, such effects are assumed to begin one day before a public announcement. The day +1 account for the possibility of post-announcement effects caused by the announcement. This paper implicitly assumes that the announcement effect does not extend beyond day +1.

[7] Examples of empirical work which apply such adjustments include, among others, Keim (1983), Basu (1983), Blume and Stambaugh (1983), Reinganum (1990).

squares estimators of beta are biased and inconsistent. Specifically, beta estimates for relatively frequently (infrequently) traded stocks are generally upward (downward) biased.

There are two reasons why higher lags and leads for market returns were ignored. First, the use of fairly liquid and actively traded securities may reduce the effect of spurious auto-and cross-correlation. This is because the problem is expected to be more pronounced on estimated returns for thinly traded stocks, mostly those of small firms, for which nontrading[8] and poor price adjustment is generally a larger problem. Specification (1) essentially uses Dimson's (1979)-adjustment procedure, and can be thought of as a standard ("naïve") model against which more elaborate models should be judged. The model uses total rather than excess returns because when daily data is used, the excess returns are almost indistinguishable from the total returns. This is because the daily risk-free rate is of the order of only about 0.02%.[9] The adjustment for a possible bias in beta is performed via the extended Fowler and Rorke (1983) procedure. This procedure ignores the variables DA and DA*Rmt in (1) and obtains the following system of equations:

$$\begin{aligned} \beta_j^{-1} &= \beta_{j-1} + \beta_{j0}\rho_{1m} \\ \beta_j^{0} &= \beta_{j-1}\rho_{1m} + \beta_{j0} + \beta_{j+1}\rho_{1m} \qquad (2) \\ \beta_j^{+1} &= \beta_{j0}\rho_{1m} + \beta_{j+1} \end{aligned}$$

where $\rho_{1m}$ is the first order serial correlation coefficient, $\beta_{j-1}$, $\beta_{j0}$, and $\beta_{j+1}$ are obtained from the [Dimson (1979)] multiple regression (1), and $\beta_j^{-1}$, $\beta_j^{0}$, and $\beta_j^{+1}$ are the Scholes and Williams (1977) coefficients from simple regressions.

The set of equations in (2) implicitly assume stationarity of the returns series Rjt and Rmt, and that all serial correlation coefficients are zero except $\rho_{1m}$.

From (2) we arrive at:

$$\sum_{k=-1}^{1} \beta_j^k = \beta_{j-1}(1+\rho_{1m}) + \beta_{j0}(1+2\rho_{1m}) + \beta_{j+1}(1+\rho_{1m})$$

If we divide $\sum_{k=-1}^{1} \beta_j^k$ by $(1+2\rho_{1m})$ we obtain the Scholes and Williams consistent estimator of beta, $\text{plim}\hat{\beta}_j$. This consistent estimator is equivalent to the modified Dimson (1979) estimator (see Fowler and Rorke, 1983). That is:

$$\text{plim}\hat{\beta}_j = \frac{1}{\vartheta}[\,\omega^* \beta_{j-1} + \vartheta\beta_{j0} + \omega^* \beta_{j+1}] \qquad (3)$$

where $\omega = (1+\rho_{1m})$ and $\vartheta = (1+2\rho_{1m})$.

## Systematic Risk and the Quality of Information

The analytical expression that links systematic risk around public announcements, the quality of public announcement, and the quality of the pre-announcement information is formerly developed in Mohamed (2001) and Hillier *et al.* (2003). For clarity a brief theoretical explanation underpinning the analysis is provided.

In the KV (1991a) economy, securities markets are assumed to be composed of a countably infinite number of traders who are risk-averse and three time periods referred to as periods 1, 2, and 3. The model endows agents with a high degree of rationality. For example, traders are capable of learning from prices and the activities of other traders. Trading is assumed to take place in periods 1 and 2, and consumption occurs in period 3. There are two assets in the economy; a riskless bond and a risky asset that represents ownership in the firm. One unit of riskless bond guarantees a payment of one unit of consumption good in period 3. The risky asset has an uncertain liquidating value denoted by a random variable u, and is also realised in period 3. At the beginning of period 1 agents are assumed to have identical beliefs about u, which are represented by a normal distribution with mean and precision (inverse of variance) h.

For reasons mentioned earlier, the analytical relationship between the announcement period beta and the precision of the public announcement

[8] It is not fair to direct all problems of spurious auto-and cross-correlation to nontrading. Lo and Mackinlay (1990b), for example, argue that their nontrading model can not explain some levels of autocorrelation found in the data. Foerster and Keim (1992) document that periods when nontrading is highest do not correspond to periods during which daily autocorrelations are the largest.

[9] See for example, Bodie *et al.* (1999), p.283.

relative to the average posterior beliefs of traders is restated as:[10]

$$\beta_{ja} = \frac{n}{\theta}\left\{ R^{-1}_{mta} \left( \left[ R'_{jL} + \frac{\eta_2}{P_1} \right] - \frac{\theta}{n} \left[ \alpha_{ja} + \xi_{ja} \right] \right) \right\} \tag{4}$$

The relationship expressed in (4) clearly generates insights into the dynamics of systematic risk around information-intensive periods. For instance, all else being equal, the relationship predicts that systematic risk of a security at the time of public announcement is, on average, increasing in the precision of the announcement, n, and inversely related to the quality of the pre-announcement private and public information. In accordance with the analytical work of KV (1991a), the (average) quality of pre-announcement information set is captured by θ.[11] To the extent that the amount of pre-announcement information for firms is increasing in firm size, then equation (4) predicts that the public announcements of small firms should be accompanied by relatively strong price reaction.

### Control Variables

The control variables considered in this paper are constructed using the operating and financial leverage ratios, the magnitude of pre-announcement betas as measured by the modified Dimson (1979) adjustment procedure, and the average change in the magnitude of earnings per share (EPS).

The operating and financial leverage ratios are considered because of the following reasons. First, high levels of borrowing translate into riskier earnings before interest and taxes. Although this kind of financial leverage does not affect the risk or the expected return on the firm's total assets, it does push up the risk of common stocks.

Second, other things being equal, we would intuitively associate higher fixed production costs with high risk. The EBIT of a firm with high operating leverage, for instance, is relatively more sensitive to changes in the sales revenue than the EBIT of a corresponding firm with low operating leverage. This type of leverage increases as fixed costs rise and as variable costs fall. Empirical research confirms that firms with high operating leverage actually do have high systematic (beta) risk [see for example, Lev (1974), and Mandelker and Rhee (1984)]. In this paper we use an aggregate measure of leverage derived analytically from Mandelker and Rhee (1984) results.[12]

We also control for the magnitude of non-announcement level of beta. This is important for any empirical study that seeks to investigate the behaviour of beta around information intensive periods. One justification for such a practice is based on the evidence that betas tend to evolve (regress) toward a grand mean of 1.0 over time [see Blume (1971, 1975)]. A simple numerical example may help clarify this phenomenon. A security with a beta of, say, 1.8 is more likely to undergo a decrease in market risk than a security whose systematic risk is 0.7. For this reason, controlling for pre-announcement level of beta may enhance the power of our empirical tests.[13]

The last control variable considered in this paper is the magnitude of earnings change. This variable is used because firms with high accounting or cashflow betas,[14] such as cyclical

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[10] A full derivation of the model is found in Hillier *et al.*(2003).

[11] $\beta_{ja}$ is announcement period beta, P1(P2) is the price of the risky asset in the first (second) trading period, $\eta_2$ is the noise (error) in the public announcement in period 2, $R'_{jL}$ is an estimate of (liquidating) expected return in period 3, and $\alpha'_{ja}$ $R_{mta}$, and $\xi_{ja}$ denote the announcement period values for market model alpha, market return, and the residual returns for security j respectively..

[12] See Mohamed (2001) and Mohamed and Yadav (2002).

[13] For more information about this issue see Kross *et al.* (1994), and Hsieh *et al.* (1999).

[14] Accounting (cash flow) betas are similar to the usual (stock) betas except that changes in book earnings (cashflow) are used in the place of rates of returns to securities. Specifically, analysts estimate the sensitivity of each firm's earnings changes to changes in the aggregate earnings of all other firms in the market.

firms whose revenues and earnings are strongly tied to the performance of the broader economy, and therefore they should also have high stock betas. Empirical evidence by Beaver et al. (1970), Beaver and Manegold (1975), and Ball and Kothari (1991) confirm that there is a positive association between earnings changes and systematic risk changes. Thus, firms whose earnings changes are systematically positive are more likely to experience an increase in market risk while those experiencing negative earnings change are likely to undergo a decrease in beta around earnings announcements.

### Empirical Relationship Between Risk Shift and Firm-Specific Characteristics

The empirical validity of the predictions of model (4) is tested using the following regression equation:[15]

$$R\beta_{ja} = \mu_0 + \mu_1 \text{LSIZE}_j + \mu_2 \text{PRN1}_j + \mu_3 \text{MCEPS}_j + \mu_4 D_j + \zeta_j \qquad (5)$$

where $R\beta_{ja} = \beta_{ja}/\beta_j$ is an announcement period beta, j is the pre-announcement period beta for firm j, $\mu_0, \mu_1 \ldots \mu_5$ are regression parameters, Dj is the proxy for the net change in the degree of leverage,[16] and $L(\theta)$ is the operator which stands for the natural logarithm of $\theta$, and $\zeta_j$ is the error term for firm j. The variables LSIZE, MCEPS, and PRN denote the natural logarithm of the mean market value of security j (in millions of pounds) for the period January 1990 to December 1998, the magnitude of the average change in net earnings per share adjusted for rights and scrip issues, and the empirical surrogate for the unobservable precision of earnings and dividends announcements.[17] In this paper we use SIZE as a proxy for the amount (quality) of pre-announcement information [See for example, Atiase (1985), Kim et al. (1997), Mohamed and Yadav (2002), and Mohamed (2003)].

The random error $\zeta_j$, captures other factors that affect the value of systematic risk around announcement periods but not introduced in model (5). The degree of operating leverage (DOL) and financial leverage (DFL) used to estimate Dj are respectively defined as and . In this definition, the variable TADFj denotes the sum of preference capital, borrowings due within one year, and longterm-borrowing, amount of long-term debt, Ej denotes equity capital and reserves, NFAj is the net fixed assets, and TAj is the total assets in firm j. The use of the said proxies is motivated by data availability. Implicit in our computations is that DOL and DFL are good instrumental (proxy) variables for the "true" degrees of operating and financial leverages as defined by Mandelker and Rhee (1984).

There are many other variables that determine the magnitude of price (and hence systematic risk) reaction to news releases but do not appear in model (5). One of those variables is the degree of information asymmetry. Model (5) implicitly makes two assumptions about our test procedure. First, it is assumed that there is a homogeneous degree of information asymmetry around dividend and earnings announcements for all firms analysed in this study.[18] Second, although anticipated information events may stimulate relatively more information gathering, it is assumed that this process preserve the tendency for firms with more public information (assumed to be large firms) to have lower degrees of information asymmetry around announcement periods. The importance of this assumption is that the appearance of the natural logarithm of market value of equity, (LSIZEj), in model (5) also

[15] The functional form of model (5) is justified in Mohamed (2001) and Mohamed (2003).

[16] In practice, it is difficult to differentiate degrees of operating and financial leverages around announcements from those of non-announcement periods. We therefore estimate the value of $D_{t+1}$ by using [$DFL_{(t+1)}DOL_{(t+1)}$ - $DFL_t DOL_t$], where the subscript (t+1) denotes the period January 1990 to December 1998 and t denotes the period January 1988 to December 1989. The DFL and DOL are average values for the period under consideration.

[17] For analytical derivation of the precision (quality) of information and its empirical proxy, see Mohamed and Yadav (2002) and Mohamed (2003)

[18] Korajczyk *et al.* (1991) and Brooks (1996) appear to suggest that there is a reduction in asymmetric information on earnings announcement days, while Kim and Verrecchia (1994) and Gajewski (1999) argue that earnings announcements should be followed by increased information asymmetry.

controls for differences in asymmetric information that are due to differences in the amount of public information available to traders. This view is justified by the empirical work of Guo and Mech (2000).

## Hypothesis to be Tested

The hypotheses to be tested derive their legitimacy from the analytical predictions of KV (1991a, 1997) and some other previous empirical evidence on the behaviour of systematic risk around news release dates. The empirical work does not, however, attempt to conduct any formal hypotheses regarding abnormal returns or the control variables used because these are not the primary focus of the current study.

Previous empirical evidence by, among others, Kalay and Lowenstein (1985), and Ball and Kothari (1991) appear to suggest that systematic risk is likely to increase around news release dates. Consistent with this view, the appropriate null and the alternative hypothesis for risk shifts is:

$H_{01}$: There is no upward risk shift around earnings and dividend announcements: $\beta_A \leq 0$.
$H_{A1}$: There is an upward risk shift around earnings and dividend announcements: $\beta_A > 0$.

The other hypotheses are based on Kim and Verrecchia (1991a, 1997) analytical work. These hypothesis are formally presented as:[19]

$H_{02}$: The relationship between risk shift and the amount of prior information around earnings and dividend announcements is not inverse: $\mu_1 \geq 0$.

$H_{A2}$: The upward risk shift around the time of earnings and dividend announcements is inversely proportional to the amount of pre-announcement information: $\mu_1 < 0$.

$H_{03}$: There is no positive relationship between risk shift and the precision of the information content of announcement around the time of earnings and dividend disclosures: $\mu_2 \leq 0$.

$H_{A3}$: The risk shift around the time of earnings and dividend disclosures is directly proportional to the precision of the announcement: $\mu_2 > 0$.

# EMPIRICAL RESULTS

## Preliminary Analysis

For brevity, we provide preliminary correlation analysis between systematic risk, the empirical surrogate for the precision (quality) news released, and the proxy for the amount of pre-announcement information. It was found that the proxy for pre-announcement information (LSIZE) is negatively correlated with $R\beta_{ja}$ with small but significant (at conventional levels) correlation coefficient. The implied correlation between the surrogate of the pre-announcement information and $R\beta_{ja}$ is -0.180 (p-value = 0.009).

On the other hand, the empirical proxy for the precision of information, PRN, is also significantly correlated (correlation coefficient = 0.152, p-value = 0.027) with $R\beta_{ja}$ but the magnitude of correlation is also relatively small. Further more, the sign of the above correlation coefficients are consistent with the predictions of the analytical work of KV (1991a, 1997). The other variables MCEPSj and Dj are weakly correlated with the relative risk, $R\beta_{ja}$, around earnings and dividends announcement dates.

Overall, the preliminary evidence suggests that, with the exception of the proxy for the amount of the pre-announcement information and the surrogate for the empirical proxy of the precision of public news release, the remaining variables may not have much power in explaining the magnitude of systematic risk around earnings and dividends announcements. This evidence seems to provide support for KV (1991a, 1997) analytical predictions. However, it is too early to draw any strong and

[19] See Hsieh *et al.* (1999) for justification of hypotheses 2 and 3.

meaningful conclusion until the interactions between variables is taken care of in a multivariate framework.

**OLS Parameter Estimates for the Single Index Model**

The Ordinary Least Squares Method is used to estimate parameters of an aggregated coefficient market model. A summary of parameter estimates and other selected characteristics such as the number of parameters having positive or negative sign are also presented under table 1. In addition, the table also provide p-values corresponding to some hypotheses of interest. The table shows that the OLS estimate for the intercept, $\alpha_{0j}$, has a mean (median) of -0.004(-0.003) and a standard deviation of 0.046. Although the intercept for specification (1) is small in magnitude, its cross-sectional mean is statistically different from zero at the 5 percent level of significance.

The abnormal return, $\alpha_{1j}$, around announcement periods has a mean (median) of 0.099(0.043) percent and a standard deviation of 0.477 percent. Out of the 212 estimates for $\alpha_{1j}$, about 59(57) percent are significantly (insignificantly) positive, and 30(66) percent are significantly (insignificantly) negative. It is interesting to note that although the average abnormal return is small in magnitude, it is predicted to be generally positive and statistically different from zero at the 5 percent level of significance. Overall, these results are consistent with the findings of Ball and Kothari (1991) and Hsieh *et al.* (1999) who document significant positive abnormal returns of 0.159 percent in the event window [-1,0.+1] and 0.001 percent in the event window [-1,0] respectively.

The third and the sixth column in Table 1 show the estimates of beta using model (1) and the market model when lead and lag terms of the market index are excluded from the estimation process respectively. The market model beta is reported for comparability purposes only and does not enter in the main section of hypotheses testing. The cross-sectional average value of the pre-announcement beta estimated according to Dimson's (1979) procedure [after Fowler and Rorke (1983) correction] ranges between 0.122 and 1.646 with a mean (median) of 0.829(0.799) and a standard deviation of 0.283. The corresponding values for the market model beta are 0.090 and 1.573 with a mean (median) of 0.762(0.743) and a standard deviation of 0.315. The market model betas are generally lower than betas that have been adjusted for nonsynchronous trading but the difference is relatively small. This evidence suggests that while it is a standard practice to adjust the pre-announcement beta by using Dimson's (1979) procedure, the nonsynchronous trading is not generally a major problem for the sample of stocks being analysed. This result is hardly surprising since the stocks selected for analysis are relatively large, actively traded and closely watched by investment analysts.

The p-values and the alternative hypothesis (in parentheses) are shown for the null hypothesis that the average value of beta is equal to their grand mean (expected value) of unity. The test results reject the null hypothesis at all conventional levels of significance (p-value =0.000). Given the validity of the assumptions of the classical linear regression models, these results would appear to suggest that a typical security analysed in this study is not equally risky as the market as a whole.

The results of "shift" in systematic risk around the three-days' announcement window ($\beta_{Aj}$), indicate that the mean (median) shift in beta is 0.151(0.118) with a standard deviation of 0.565. The shift in beta on these event days vary from -1.615 to 2.514. The preliminary analysis so far suggests that around the time of dividends and earnings release an average firm in the sample is expected to have an announcement beta that is 18.2 percent higher than the value of the pre-announcement beta.

Although we can reject the null hypothesis ($H_{01}$) (p-value = 0.000) of no increase in systematic risk around the time of earnings and dividends announcements, this evidence only tells part of the story. A much better picture emerges when one examines the number of securities that experienced significant (insignificant) positive or negative increases in risk. Table 1 reveals that out of 212 firms, about 22 (35) percent of the firms had experienced small but significant (insignificant) positive shift in beta around earnings and dividends news releases while 11(32) percent experienced significant (insignificant) reductions in their levels of pre-announcement betas. All these tests are one-tailed and are conducted at the 5 percent level of significance.

The cross-sectional average estimate of the adjusted R-Squared for specification (1) in table 1 ranges between 1.0 and 43.2 percent with a mean (median) of 15.4 (13.0) percent and a standard deviation of 9.4 percent. Overall, the evidence appears to suggest that, on average, the market is able to explain only 15.4 percent of the changes in stock returns and the remaining 84.6 percent of variation in returns is unaccounted for by model (1).

### Analysis of the Association Between Beta Shifts and Information Proxies

Parameter estimates for model (5) are obtained by using the OLS estimation method. The dependent variable for each panel of Table 2 is the relative systematic risk $R\beta_{ja}$. Panel A reports results for the case where the dependent variable $R\beta_{ja}$ is constructed using non-standardised betas, while estimates corresponding to standardised betas appear under panel B. White (1980b) and Newey-West (1987a) standard errors are also reported. The maximum lag used to compute the Newey-West standard errors is 4. The figures in parentheses are p-values corresponding to a null hypothesis $H_0$: PES = 0 against the alternative $H_0$: PES$\neq$ 0.[20]

We begin with results for which the non-standardised announcement beta ($\beta_{ja}$) and the pre-announcement beta ($\beta_j$) are used to construct the dependent variable $R\beta_{ja}$. The coefficients of our proxy for the amount of pre-announcement information, LSIZE, for panel A is -0.110. The parameter estimate has the predicted sign and is statistically significant at the 5 percent level. The significance of the coefficient of LSIZE implies that the null hypothesis $H_{02}$ is rejected. To the extent that LSIZE is a legitimate surrogate for the quality of pre-announcement information, this evidence confirms the predictions of equation (4) and hence consistent with the implications of the analytical work of KV (1991a, 1997). Specifically, the level of systematic risk at the time of public news releases appears to be less for firms with a lot of pre-announcement information compared to firms with relatively little prior information.

For the parameter associated with the proxy for the precision of the announced information, the estimates of $\mu_2$ for panel A is 0.730. Again, estimate of $\mu_2$ is positive and significant at the 10 percent level or lower. Thus, the evidence presented in panel A suggests that the announcement period beta is larger following a precise (quality) news release than for cases when firms announce imprecise information. Clearly, this evidence is not consistent with the null hypothesis $H_{03}$, and therefore confirms KV's (1991a, 1997) theoretical predictions. Similar results are observed under panel B.

### SENSITIVITY OF EMPIRICAL RESULTS TO TIME-VARIATION IN BETA

So far our analysis has implicitly assumed that systematic risk of sample stocks is a constant over time. It may well be argued that our results are due to our failure to account for the time-variation in common stocks betas. Ignoring temporal variation in betas is not consistent with the evidence of time-variation in systematic risk documented by, among others, Sunder (1980), Schwert and Seguin (1990), Kim (1993), Koutmos et al. (1994), Pope and Warrington (1996), and Faff et al. (2000). We examined the robustness of empirical results to time variation in systematic risk by using betas estimated from Schwert and Seguin (1990) heteroscedastic market model:

$$R_{jt} = \alpha_j + \beta_j R_{mt} + \delta_j \left( \frac{R_{mt}}{\hat{\sigma}^2_{mt}} \right) + \varepsilon_{jt} \tag{6}$$

where $\beta_j$ is a constant, $\frac{\delta_j}{\hat{\sigma}^2_{mt}}$ is the time-varying term, and $\hat{\sigma}^2_{mt}$ is the estimate of the variance of market returns for period t.

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[20] To conserve on space, we do not test any hypotheses about the significance or signs of parameters associated with control variables.

We estimate $\sigma^2_{mt}$ and $\sigma^2_{It}$ by using the MA(1)-Glosten *et al.* (1993) GJR-GARCH (1,1) conditional heteroscedasticity specification similar to that used by Koutmos *et al.* (1994), except that we make use of asymmetric rather than the standard GARCH (1,1) model.

Our results (not reported here) suggests that even after accounting for time variation in beta, the evidence offers additional support in favour of the alternative hypotheses $H_{A2}$ and $H_{A3,}$ and corroborates the conclusions reached in the theoretical work of KV (1991a & 1997).

## CONCLUSION

The purpose of this study was to investigate systematic risk shifts around earnings and dividend announcements and its association with the amount of pre-announcement information and the precision of news disclosed.

Three main research hypotheses were tested. The first hypothesis focused on the possibility of upward systematic risk shift around earnings and dividends announcements. The second hypothesis contends that there is an inverse relationship between the level of systematic risk and the amount of pre-announcement information around earnings and dividends news releases. The last hypothesis advocates that the level of systematic risk at the time of earnings and dividends announcements is directly related to the precision of the announced news.

We find that there is a partial support for the first hypothesis. This finding remains true even when one corrects for estimation biases caused by ignoring time-variations in beta. Specifically, part of the evidence is consistent with Ball and Kothari (1991), and Kalay and Lowenstein (1985) that there is an increase in risk during scheduled news announcement. However, some stocks had either experienced decreases in beta or no change at all.

After controlling for the average change in earnings per share over the sample period, the level of pre-announcement beta, and the degree of operating and financial leverage, we find that there is generally a strong negative statistically significant relationship between relative systematic risk, $R\beta_{ja}$, around news event dates and the proxy for the amount of prior information. In generally, this association remains true whether one constructs $R\beta_{ja}$ using non-standardised betas or standardised beta estimates. To the extent that LSIZE (PRN1) is a legitimate surrogate for the theoretical construct of the quality of the pre-announcement information (the precision of the disclosed news) implied by KV, our evidence confirms their analytical predictions. That is, for a given level of precision and other firm's characteristics, firms with a lot of prior information experience relatively smaller price responses at the time of earnings and dividends announcements. This phenomenon, in turn, manifests itself into relatively small shifts in beta around news events. The opposite is generally true for firms with relatively small amount of prior information. That is, changes in beta should be greater for firms with relatively little amount of pre-announcement information.

The above result is hardly surprising when compared with the empirical work of Richardson (1984), who reports that firms experiencing higher shifts in risk in the week of their annual earnings typically are of smaller size. It is also due to the evidence that price reaction to news announcement is generally positively correlated with the amount of surprise in the earnings (dividends) news (see Ball and Kothari, 1994). In fact, analysts' forecast, management's' earnings forecasts, rumours, news about expenses and/or sales, and insider dealings (trades) generate a lot of information about firms. In addition, more analysts and press following potentially results in systematically more information impounded in stock prices (see Richardson, 1984). To the extent that the amount of prior information is proportional to the firm size, we would expect that most of the information content of forthcoming announcements for large firms to be already anticipated by the market. This partial anticipation of information content of announcements may cause earnings (dividends) releases for large firms with a lot of prior information to be associated with relatively weak price reaction compared to releases of firms (usually small) with little prior information.

We also find that relative systematic (market) risk, $R\beta_{ja}$, around news event dates is generally positively related to the empirical proxy for the precision (quality) of earnings and dividends announcements. This result is in line with the predictions of hypothesis $H_{A3}$. Overall, empirical results are fairly robust to the procedure used to estimate systematic risk.

**Table 1:** *Parameter Estimates Based on a Single Index (Market) Model*

| Parameter → | $\alpha_{0j}$ | $\beta_{pj}$ | $\alpha_{1j}$ | $\beta_{Aj}$ | $\beta_{mktj}$ | Adjusted $R^2$ |
|---|---|---|---|---|---|---|
| Mean | -0.004 | 0.829 | 0.099 | 0.151 | 0.762 | 0.154 |
| p-value | 0.200(≠ 0) | 0.000(≠ 1) | 0.003(≠ 0) | 0.000(>0) | 0.000(≠ 1) | 0.000(≠ 0) |
| Median | -0.003 | 0.799 | 0.043 | 0.118 | 0.743 | 0.130 |
| Standard Deviation | 0.046 | 0.283 | 0.477 | 0.565 | 0.315 | 0.094 |
| SGP(IGP) | 20(77) | - | 59(57) | 47(75) | - | - |
| SGN(IGN) | 18(97) | - | 30(66) | 23(67) | - | - |

Table 1 presents summary statistics (cross-sectional) of parameter estimates for a sample of 212 firms over the period January 1, 1990 to December 31, 1998. Parameter estimates are based on the following model:

$$R_{jt} = \alpha_{0j} + \sum_{k=-1}^{1} \beta_{jk} R_{m(t+k)} + \alpha_{1j} D_A + \beta_{Aj}(D_A * R_{mt}) + \varepsilon_{jt}$$

where Rjt is the rate of return to a security j in period t, $R_{mt}$ is the rate of return to the market portfolio in period t, DA is an announcement dummy which assumes the value of one on the dates [-1, 0, +1] and has the value of zero otherwise(day t= 0 denotes earnings or dividend announcement date), $\alpha_{0j}$,$\alpha_{1j}$ (abnormal return), are regression parameters, $\beta_{Aj}$ is an announcement beta, and $\beta_{jk}$ for k= -1, 0, and 1 are the observed lag, intertemporal and lead security betas, and the residuals series, { jt}, measures the extent to which the actual return to security i in period t differs from the expected rate of return.

The symbols SGP, IGP, SGN, and IGN denote significantly positive, insignificant but positive, significantly negative, and insignificant but negative at 5 percent significant level respectively.

The estimate $\beta_{pj}$ denotes a pre-announcement beta estimated according to Dimson's (1979) procedure after Fowler and Rorke (1983) correction, while $\beta_{mktj}$ denotes a ("market model") beta estimate when leads and lag terms of the market index are excluded from the above model. The row labelled p-value shows the alternative hypothesis(in parentheses) and the corresponding probability(attained of the test) of obtaining a value as large in magnitude(extreme or more extreme) as the t-statistic corresponding to null hypotheses $H_0$: $\alpha_{0j} = 0$, $H_0$:$\beta_{pj} = 1$, $H_0$: $\alpha_{1j} = 0$, $H_0$:$\beta_{Aj} \leq 0$, and $H_0$:$\beta_{mktj} = 1$ respectively. For a reasonably large sample the null is rejected when p-value< the chosen level of significance.

**Table 2:** *Parameter Estimates based on the Regressions of Changes in Relative Risk around Firms' Earnings and Dividends Announcements on Firm Specific Variables*

Panel A: *Dependent Variable is $R\beta ja$ using "Market" Betas from Model I*

| PR | $\mu_0$ | $\mu_1$ | $\mu_2$ | $\mu_3$ | $\mu_4$ | |
|---|---|---|---|---|---|---|
| PS | UN | NG | PO | PO | PO | Adj.$R^2$(%) |
| | | | | | | - |
| PES | 1.797 | -0.110 | 0.730 | -0.004 | -0.417 | 4.00 |
| W-se | 0.398 | 0.052 | 0.313 | 0.010 | 0.491 | |
| | (0.000) | (0.018) | (0.010) | (0.712) | (0.397) | |
| NW-se | 0.484 | 0.063 | 0.292 | 0.011 | 0.492 | |
| | (0.000) | (0.041) | (0.006) | (0.732) | (0.398) | |

Panel B: *Dependent Variable is $R\beta ja$ using Standardised "Market" Betas from Model I*

| | | | | | | |
|---|---|---|---|---|---|---|
| PES | 1.176 | -0.024 | 0.174 | -0.001 | -0.089 | 4.20 |
| W-se | 0.087 | 0.011 | 0.068 | 0.002 | 0.102 | |
| | (0.000) | (0.018) | (0.006) | (0.510) | (0.385) | |
| NW-se | 0.107 | 0.014 | 0.064 | 0.002 | 0.104 | |
| | (0.000) | (0.042) | (0.004) | (0.538) | (0.394) | |

This table reports regression results for the model used to conduct hypothesis tests of the empirical relationship between systematic risk shifts and firm-specific characteristics. The model is specified as:

$$R\beta_{ja} = \mu_0 + \mu_1 LSIZE_j + \mu_2 PRNI_j + \mu_3 MCEPS_j + \mu_4 D_j + \zeta_j$$

where $R\beta_{ja} = \beta_{ja}/\beta_j$, $\beta_{ja}$ is an announcement period beta, $\beta_j$ is the pre-announcement period beta for firm j, $\mu_0, \mu_1, \ldots, \mu_4$ are regression parameters, $D_j$ is the proxy for the net change in the degree of leverage, and $L(\theta)$ is the operator which stands for the natural logarithm of $\theta$, $\zeta_j$ is the error term for firm j, and all the other independent variables are as defined in the paper.
The symbol PR, PS, PES, W-se, and NW-se denote parameter of interest, predicted sign of the parameter, parameter estimate, White (1980) standard error, and Newey-West (1987) standard error respectively. The maximum lag used in the computations of Newey-West standard errors is 4. The symbol UN, NG, and PO are used to denote unknown, negative, and positive signs respectively. The figures in parentheses are p-values corresponding to a null hypothesis $H_0$: PES = 0 against the alternative $H_0$: PES $\neq$ 0, for all parameters, except for $\mu_1$ and $\mu_2$, for which one-tailed p-values are reported in line with the structure of hypothesis 2 and hypothesis 3. For a reasonably large sample the null is rejected when p-value< the chosen level of significance.

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