The Sarbanes-Oxley Act of 2002: Implications for Market Efficiency and Analysts’ Performance

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Abstract

The Sarbanes-Oxley Act, enacted as a response to the multiple cases of corporate fraud during the years 2001-2002, is considered to be one of the most important reforms in the corporate disclosure policy with potentially far-reaching implications for the stock markets. In this paper we study these implications for two major "consumers" of information released by companies, namely investors and analysts. We find that following the reform the US stock market is characterized by a significantly higher speed of adjustment to new information, suggesting that information is more rapidly incorporated in the stock prices by investors, thus, leading to an increase in informational market efficiency. On the other hand, we find strong evidence of the analysts’ forecasts becoming more overpessimistic in the post-reform period, suggesting that following the corporate scandals of 2001-2002 analysts became more cautious. An increase in the analysts’ overpessimism is especially pronounced during turbulent periods and when "good news" is released. We interpret these findings as the evidence that following the bankruptcy of reknown firms with blown-up earnings, analysts treat the disclosed information more carefully.

1 Introduction

Transparent corporate disclosure of information is undoubtfully one of the corner stones needed for an efficient functioning of stock markets. Financial statements, reports, and announcements serve as a useful valuation tool both for analysts, who submit their forecasts and investment recommendations based on this data, and for the investors who use this information to price financial instruments and to choose optimal portfolio strategies. In this context the Sarbanes-Oxley Act (hereafter SOX act) is unanimously described as one of the most far-reaching and significant changes in the disclosure obligations of publicly traded companies (Smith, 2002; Ribstein, 2003). The purpose of this paper is to study the reaction of the stock market participants, the investors and the analysts, to this reform.
The Sarbanes-Oxley Act, signed into law on 30 July 2002, was enacted as a response to a series of severe corporate scandals that have shaken the confidence of the public in the stock markets in particular and in the financial institutions in general. On 16 October 2001 Enron Corporation announced a 35 million decrease in stated earnings and a 1.2$ billion loss of shareholder equity due to accountancy misreporting, a statement followed by the November announcement that an additional 500$ millions in earnings were overstated. On 2 December 2001 Enron declared itself bankrupt, making history as the largest bankruptcy in the corporate history of the United States. However, this was just the beginning. The bankruptcy of Global Crossing in January 2002 and Adelphia Communications in March 2002 due to inflated earnings, Xerox, admitting more than 6$ billion in overstated earnings on June 2002, and Worldcome, filing for bankruptcy on July 2002 - these are just a few landmarks that uncovered severe institutional and regulatory problems in the financial reporting of firms. The costs of these misreportings led scandals were severe: Standard and Poor’s 500 index lost about 20 percent of its value during the year 2002 (see Figure 1). This decline has been accompanied by an upward shift in the Implied Volatility Index (VIX), indicating a sharp increase in overall market uncertainty (see Figure 2 ) and a downward shift in the consumer confidence index (Hevesi, 2003).

As a response to these multiple cases of corporate fraud the SOX act concerns the following key aspects (for the comprehensive reviews see Smith, 2002 and Ribstein, 2003 among others)

1 **New disclosure requirements.** The companies are required to provide an internal control report as part of their annual report. Also, the quarterly and annual reports must disclose all off-balance sheet transactions and other material changes in their financial conditions on a "rapid and current basis". Insider trading reporting rules are substantially tightened as well.

2 **Securities analysts regulation.** A new regulation ensures the analysts’ independence from their firms’ investment banking activities.
3 Internal monitoring and gatekeeper regulation. More responsibility for the independent board audit committee for hiring and overseeing auditors; prohibits fraudulently influencing and/or misleading auditors; reducing ties between auditors and audited companies.

As the SOX act is aimed to increase the transparency of the information disclosure by the firms and, in this way, to restore the confidence of the public in the stock markets, an important question is whether this goal has been achieved. In particular, if the information disclosure did become more transparent and news is revealed on a more frequent and reliable basis, one would expect the stock prices to respond more rapidly to news regarding the financial and economic conditions of the companies. However, although the SOX act’s primary goal was to increase the information disclosure, to the best of our knowledge, no study examines whether and, if so, to what extent the speed of adjustment to the new information of stock markets has changed after the Act has been signed into law. The first purpose of this paper is, therefore, to fill this gap by studying whether the informational efficiency of the US stock markets has changed following the SOX act.

Also, since corporate news is an important source of information for analysts, an increase in the information disclosure should increase the accuracy of the analysts’ forecasts. However, if the Act has not succeeded in restoring public confidence, shaken by the corporate scandals, the picture might be reverse. In particular, we may expect both investors and analysts becoming more cautious in interpreting the news, which would result in stock prices reflecting more slowly the information on the one hand and the analysts’ forecasts becoming over pessimistic on the other hand. The second purpose of this paper is, therefore, to study the implications of the SOX for the accuracy of the analysts forecasts.

In terms of methodology we estimate the partial adjustment model with noise of Amihud and Mendelson (1987), which we apply to all the stocks listed on NYSE/AMEX during the last decade. Within this framework we test for the presence of structural breaks in the speed of adjustment coefficients after the SOX act has been signed into law. Next, we apply
nonparametric tests to the analysts forecasts and actual earnings data of the abovementioned firms to study whether the corporate scandals and the SOX act have influenced the performance of the analysts and, if so, in which direction.

We find strong support in favor of markets becoming more informationally efficient in the post SOX period, with average speed of adjustment exhibiting a substantial increase. On the other hand, we also find strong evidence of the analysts’ forecasts becoming less accurate, and, in particular, more "overpessimistic" in the post-SOX era. Though we also find some weak evidence in favor of the SOX act affecting the analysts’ performance in the recent two years, overall the positive impact of the legislation appears to be overwhelmed by the impact of the corporate scandals which have distorted the confidence of the analysts in the information provided by the firms.

The remainder of this paper is organized as follows. In Section 2 we briefly overview some relevant literature on the stock market informational efficiency and the performance of the firms’ analysts. Section 3 describes the data. In Section 4 we formulate our research questions and discuss the methodology applied in this paper. In Section 5 we present and discuss our estimation results. Finally, in Section 6 we present our concluding remarks and propose some directions for further research.

2 Literature Review

The concept of the "efficient market hypothesis" has been formalized by Fama (1970) and is related to the question of whether the pricing information is fully incorporated in stock prices. The question of market efficiency has been examined by an extant body of empirical studies. While providing a comprehensive review of the market efficiency literature is clearly out of the scope of this paper, we shall briefly review a number of widely cited studies in this field, related to the issue examined in this paper.1 Amihud and Mendelson (1987) report significant violations from the null of a random walk in favor of ARMA(1,1) model for the sample of Dow Jones components, suggesting that stock prices do not fully adjust to a new information. Lo and Mackinlay (1988) reject the null that stock prices follow a random walk based on the variance-ratio test. Damodaran (1993) estimates the speed of adjustment coefficients for the NYSE/AMEX sample of stocks and finds that in general the stocks are characterized only by the partial adjustment to the partial adjustment or "underreact". In a more recent work Lo and MacKinlay (1999) find that short-run serial correlations are not zero and the existence of "too many" successive moves in the same direction suggests the existence of "momentum" in short-run stock prices, a finding supported by Theobald and Yallup (2004) who report an overall tendency of NYSE/AMEX listed stocks to underreact in the short run. Lo, Mamaysky, and Wang (2000) by using nonparametric techniques report that some technical analysis based rules do have some predictive power. Overall, these papers suggest that in a short-run the adjustment of stock prices to a new information is less than full, possibly due to investors’ underreaction.

A number of studies examine the relationship between the market efficiency and information disclosure. Collins et al. (1987) and Freeman (1987) show that differences in information environment affect the extent to which price changes anticipate earnings changes.

1For a survey of the market efficiency literature a reader is referred to Schwert (2003)
a finding supported by Collins and Kothari (1989). Imhoff and Lobo (1992) find that earnings news have a greater impact on unexpected stock price change as the amount of pre-earnings-announcement uncertainty decreases. In a more recent study Douthett et al. (2003) find that earnings surprises tend to have a larger impact on the stocks of the firms which are required to report financial information under more stringent disclosure rules. Ng et al. (2006) provide evidence that the stocks of firms with higher quality disclosure are associated with a smaller underreaction. The overall conclusion of these studies is that the speed of adjustment of security prices to a new information is related to the quality of the information disclosure. In context of this study, an important question is whether the extent of underreaction has changed after the introduction of the SOX.

The second issue of interest is whether the enactment of the SOX act has influenced the performance of analysts, and, in particular, whether the analysts’ forecasts of the companies’ earnings became less biased. Numerous studies indicate the presence of a bias in the analysts’ earnings forecasts (see Brown,1993 for a review of the related literature). Fried and Givoly (1982) and O’Brien (1988) show that the analysts’ earnings forecasts generally are overoptimistic. Similar results are reported by Stickel (1990) and Abarbanel (1991), who find that the mean estimate of the analysts’ forecast errors is significantly negative. Lim (2001) proposes a model where analysts trade-off bias to improve management access leading to optimal forecasts being overoptimistic. More recent studies, however, find that the sign of the bias is unstable over time. Brown (1997) documents significant rightward temporal shifts in mean earnings surprises between 1984 and 1996. Similar findings are reported by Brown (2001), which are consistent with firms’ management desire to meet or beat analysts forecasts as reported by Degeorge et al. (1999).

Several studies examine the relationship between the forecast accuracy and the level of information disclosure. Waymire (1986) finds that the accuracy of the analysts earnings forecasts improves after the management earnings forecast is released. For the US firms Lang and Lundholm (1996) find that firms with more informative disclosure policies are followed by a larger number of the analysts and have more accurate analyst earnings forecasts. Basu, Hwang, and Jang (1998) find that country-average levels of disclosure are positively associated with analysts’ levels of accuracy for the sample of seven countries. Similar results are reported by Khanna, Palepu, and Chang (2000) for a sample of 37 countries and by Hope (2002) for a sample of 22 countries. The overall conclusion from these and other studies is that higher disclosure enhances the analysts’ forecast accuracy. Therefore, studying the question of whether the SOX act led to improvement in the analysts’ performance is of major importance both for policy makers and for the practitioners who incorporate the analysts’ earnings forecasts in their firm valuations.

2A term "analysts’ forecasts bias", as it is somewhat loosely used in academic literature, refers to a situation when the expected value of the analysts’ forecast error (actual earnings minus consensus forecast) is different from zero.
3 Data Description

3.1 Sample Selection

Our database consists of daily observations on all NYSE/AMEX stocks with continuous data from January 1998 to December 2005. This data includes closing prices adjusted for splits, daily trading volume and the number of shares outstanding for each security included in our sample. The data has been obtained from the CRSP tapes. Similar data, though over a different time span, has been used in related studies by Damodaran (1993) and Theobald and Yallup (2004), though for different purpose. Following the standard convention we exclude from our analysis NASDAQ stocks. This screening rule leaves us with a total number of 1513 firms.

To study potential implications of the SOX act for the analysts’s performance for all the firms included in our sample we collect data on their actual and predicted earnings over the period January 1998-June 2006 on a quarterly basis. More specifically, each firm-quarter observation includes actual earning, mean analysts’ forecast, highest/lowest forecast estimates, and the announcement date. All data has been obtained from the Institutional Brokerage Estimate System (I/B/E/S). Following related studies (Datta and Dhillon, 1993; Shane and Brous, 2001) we define the consensus forecast as the mean forecast from the last month before the announcement date for each firm-quarter observation. To be included in our sample a firm must have at least two years (eight quarterly observations) both before and after July 2002 when the SOX act has been signed into law. Also, to be included in our sample we require each observation to have at least two different analysts’ estimates. To take care of possible outliers caused either by special items or by data input errors we filter our data with the Grubbs algorithm (Barnett and Lewis, 1994). After applying these filtering rules our final dataset consists of 24380 firm-quarter observations.

3.2 Exploratory Analysis

In this subsection we conduct an exploratory analysis of our data. First, we present the descriptive statistics and discuss the properties of stock returns of the firms included in our sample. Next, we conduct a preliminary analysis of the earnings and the analysts’ forecasts data.

3.2.1 Stock returns data

A number of studies report the speed of adjustment of the large capitalization stocks being different from those of the small size firms (Damodaran, 1993; Theobald and Yallup, 2004). Therefore, for the purpose of further analysis we sort all the firms in our sample into ten deciles based on their average CRSP capitalization decile assignment during the time span of our study.

In Table 1 we present some descriptive statistics of the daily close-to-close log returns ranked by their market capitalization as described above. For each statistic (mean, standard deviation etc) we report its cross-section average across all the securities according to their capitalization decile assignment. Most of the stocks exhibit positive drift which, starting from the fifth decile, becomes statistically significant. Also, for the stocks included
in our sample the null hypothesis of normally distributed returns on average is strongly rejected based on a highly significant excess kurtosis. This excess kurtosis, however, can be partially attributed to GARCH-type effects, based on highly significant estimates of the autocorrelation coefficient between the squared returns. Raw returns on average also appear to be significantly autocorrelated, though both the sign and the magnitude of the estimates seem to change from significantly negative for the low decile portfolios to positive (though statistically insignificant) for the high-cap stocks. Interestingly, average estimates of both standard deviation and mean turnover almost monotonously increase from low to high-cap portfolios. Since both trading volume/turnover and volatility are usually considered to be a proxy for the information flow\(^3\), these findings suggest that trading in high-cap stocks is likely to be more informationally intensive. Finally, for all the stocks daily returns appear to be negatively skewed and the magnitude of skewness coefficient estimate tends to decline (though not monotonously) from the low to the high capitalization deciles. In conjunction with monotonously increasing turnover this finding can be attributed to Hong and Stein (2003) "dispersion of beliefs" model.

<table>
<thead>
<tr>
<th>Table 1: Descriptive statistics—daily returns</th>
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<tbody>
<tr>
<td>Decile</td>
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<tr>
<td>Mean</td>
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<tr>
<td>St. Deviation</td>
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<tr>
<td>Skewness</td>
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<td>Kurtosis</td>
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<tr>
<td>Corr(r_t, r_{t-1})</td>
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<td>Corr(r_t^2, r_{t-1}^2)</td>
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<tr>
<td>Daily Turnover</td>
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<tr>
<td>Decile</td>
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<tr>
<td>Mean</td>
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<tr>
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<td>Daily Turnover</td>
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</table>

The estimates of the mean, standard deviation and mean turnover are presented in percentage points. *(**) denoted significance at 10 (5)%. Corr\(r_t, r_{t-1}\) and Corr\(r_t^2, r_{t-1}^2\) denote serial correlation of raw and squared returns respectively. Turnover is measured as the number of share traded scaled by the number of shares outstanding.

\(^3\)See, for instance, Andersen (1996).
3.2.2 Earnings data

Next, we conduct a preliminary analysis of the "earning surprise" or the forecast error series which we define as actual earning minus consensus forecast. For the purpose of further analysis we group all firms in our sample into five portfolios based on the extent to which actual earnings can be predicted by the analysts. Define by \( \hat{\sigma}_{e,i}^2 \) the estimated variance of the actual earnings for a company \( i \). Also, denote by \( \hat{\sigma}_{e,i}^2 \) the estimated variance of the forecast errors for that company. We group all the firms included in our sample into five portfolios based on the variance-ratio criterion, that is, based on the value of their \( \frac{\hat{\sigma}_{e,i}^2}{\hat{\sigma}_{e,i}^2} \) ratio. This measure, calculated for each firm, provides us the information on how well the analysts' forecasts explain the variation of the actual earnings. The first portfolio includes the stocks with the lowest (first quintile) variance ratio while the fifth portfolio includes the firms with the highest (fifth quintile) variance ratio estimates.

Descriptive statistics are reported in Table 2. A number of interesting observations can be mentioned. First, the earning surprise series appears to be highly leptocurtic with estimates of kurtosis significantly exceeding the one corresponding to the normal distribution. Also, the earning surprises on average exhibit higher dispersion as we move from the lower to the upper quintiles, based both on the estimates of the standard deviation and average spread (highest forecast minus lowest forecast). The estimates of mean and skewness are particularly interesting. First, we find the mean of the earning surprise to be significantly negative for all the portfolios included in our sample, a finding which suggests that on average analysts tend to submit overpessimistic forecasts. This finding is consistent with the results of DeGeorge et al. (1999) and Burgstahler and Eams (1998) who report that companies’ management tends to report earnings that meet or beat analysts forecasts. Interestingly, both mean and skewness estimates tend to decline from lower to upper quintile portfolios, with skewness switching from positive and significant to slightly negative, suggesting that the degree of the analysts overpessimism tends to decline with an increase in ex-post uncertainty. These findings can be attributed to the bias-variance trade-off as suggested by Lim (2001). To test this conjecture in a last row of Table 2 we report sample correlations between the earning surprise and spread, which can be interpreted as an ex ante measure of the uncertainty, that is, a measure of the analysts’ dispersion of beliefs. While being statistically insignificant for the portfolios 3-5, for the lower quintile portfolios the estimates are positive and statistically significant, suggesting that during periods of high uncertainty analysts are likely to become more dependent on the information they receive from the company management.
Table 2: Descriptive statistics-earning surprises

<table>
<thead>
<tr>
<th>Quintile</th>
<th>1</th>
<th>2</th>
<th>3</th>
<th>4</th>
<th>5</th>
</tr>
</thead>
<tbody>
<tr>
<td>Mean</td>
<td>0.015**</td>
<td>0.016**</td>
<td>0.012**</td>
<td>0.011**</td>
<td>0.007**</td>
</tr>
<tr>
<td>St. Deviation</td>
<td>0.034</td>
<td>0.056</td>
<td>0.065</td>
<td>0.077</td>
<td>0.101</td>
</tr>
<tr>
<td>Skewness</td>
<td>0.841**</td>
<td>0.459**</td>
<td>0.093</td>
<td>-0.146</td>
<td>-0.17</td>
</tr>
<tr>
<td>Kurtosis</td>
<td>7.82**</td>
<td>7.71**</td>
<td>8.16**</td>
<td>7.78**</td>
<td>8.1**</td>
</tr>
<tr>
<td>Spread</td>
<td>0.057</td>
<td>0.086</td>
<td>0.088</td>
<td>0.097</td>
<td>0.084</td>
</tr>
<tr>
<td>Corr(spr,surp)</td>
<td>0.17**</td>
<td>0.11**</td>
<td>-0.001</td>
<td>0.004</td>
<td>0.018</td>
</tr>
</tbody>
</table>

** denotes significance at 5%

4 Research Questions and Methodology

4.1 Market Efficiency

Following a definition proposed by Schreiber and Schwartz (1986), a price discovery process is a process during which the stock price converges to its equilibrium or "intrinsic" value. When the markets are fully efficient any change in intrinsic value of the firm should be immediately reflected in its stock price. Following this concept, the question of whether the market has become more (less) informationally efficient due to some market-wide event can be analyzed by comparing the speed of adjustment coefficients before and after the event has occurred.

As noted by Fama (1992), testing the hypothesis of stock markets being informationally efficient is in fact a test of a joint hypothesis of market efficiency and correct specification of the model. Following Damodaran (1993), Theobald and Yallup (2004) and other related papers, our study of the market efficiency is based on a partial adjustment with noise model of Amihud and Mendelson (1987) (denoted as A&M model). This model explicitly specifies stochastic processes for the observed log-price series and the underlying latent intrinsic value series. The log of the intrinsic value is assumed to follow a random walk with drift, thus, assuming that the equilibrium unobserved price is fully efficient in a sense that it immediately incorporates any information shock. The observed price and the intrinsic value of firm $i$ at period $t$ are specified as

$$
\Delta P_{i,t} = \pi_{i,t}(V_{i,t} - P_{i,t-1}) + u_{i,t} \quad i = 1, \ldots, n, \ t = 1, \ldots, T
$$

$$
\Delta V_{i,t} = \mu_i + e_{i,t}
$$

where $\Delta P_{i,t} = P_{i,t} - P_{i,t-1}$ denotes the change in the observed price, $V_{i,t}$ is the unobserved intrinsic value (both expressed in natural logarithms), $\pi_{i,t}$ is the speed of adjustment coefficient which lies in interval $(0, 2)$ to keep the observed return process stationary, $\mu_i$ and $e_{i,t}$ are the drift term and the information shock to the intrinsic value, respectively, and $u_{i,t}$ is a bid-ask spread related noise. Both $u_{i,t}$ and $e_{i,t}$ are assumed to have zero mean and to be serially and cross sectionally-uncorrelated at all leads and lags. In this setting, a fully efficient market corresponds to the case of $\pi_{i,t}$ being equal to 1 for all the firms. When $\pi_{i,t}$ lies between 0 and 1 investors systematically underreact to the news, while for $\pi_{i,t}$ lying in
interval (1, 2) overreaction occurs.\(^4\)

Theobald and Yallup (2004) show that the observed price process can be easily reformulated as

\[
\Delta P_{i,t} = \pi_{i,t} \mu_i + (1 - \pi_{i,t}) \Delta P_{i,t-1} + \pi_{i,t} e_{i,t} + u_{i,t} - u_{i,t-1}
\]

A General Method of Moments (GMM) estimator of \(\pi_{i,t}\) based on instrument variables can be easily constructed and estimated (the choice of the instruments will be discussed in the following section).\(^5\)

The A&M model provides a simple and intuitive way of analyzing the process of markets incorporating new information with the speed of adjustment \(\pi\) being the key parameter of interest. Within this framework we are interested in studying the following research questions

\(Q1\): Was the speed of adjustment stable over the last decade? If not, when did this change occur?

In the original A&M model the speed of adjustment \(\pi\) is assumed to be time-invariant, that is, \(\pi_{i,t} = \pi_i\). We examine this issue by testing the following null hypothesis for each firm included in our sample

\[
H_0 : \pi_{i,t} = \pi_i \forall t = \{1, 2, \ldots, T\}
\]

against the alternative that there has been a structural shift at some \(t^*\) lying between 1 and \(T\), the time span of our study. That is, under the alternative for each firm the speed of adjustment is piecewise constant and equals to \(\pi_{i,S1}\) before the structural break has occurred and \(\pi_{i,S2}\) in the post-break period, with the indices \(S1\) and \(S2\) denoting before and after the structural shift, respectively. We apply two different tests: the supremum Wald test of Andrews (1993) and the exponential Wald test developed by Andrews and Ploberger (1994). While, as shown by Andrews and Ploberger (1994), the latter test enjoys certain optimality properties (which is not the case for the former), the supremum Wald test also allows us to determine when a structural break has occurred, given that the null of parameter stability has been rejected.

Consider the sample of length \(T\), which is partitioned into two subsamples with sample lengths \(\alpha T \in \mathbb{N}\) and \((1 - \alpha)T \in \mathbb{N}\) respectively for some \(\alpha\) lying in a given interval \([\alpha_l, \alpha_u]\) with \(\alpha_l\) and \(\alpha_u\) denoting the lower and the upper bound of the interval respectively. Also, define by \(\hat{\pi}_{\alpha T}\) and \(\hat{\pi}_{(1-\alpha)T}\) the GMM consistent estimators of \(\pi_{S1}\) and \(\pi_{S2}\), respectively, based on the first and second subsample and by \(\hat{\nu}_{\alpha T}\) and \(\hat{\nu}_{(1-\alpha)T}\) the estimates of the variance of \(\hat{\pi}_{\alpha T}\) and \(\hat{\pi}_{(1-\alpha)T}\), respectively. Then the standard Wald statistic which tests the null \(\pi_{S1} = \pi_{S2}\) takes the following form (see Andrews, 1993)

\[
W_T(\alpha) = T\left(\frac{\hat{\pi}_{(1-\alpha)T}}{\hat{\pi}_{\alpha T}} - 1\right) \left(\alpha^{-1/2} \hat{\nu}_{\alpha T} + (1 - \alpha)^{-1/2} \hat{\nu}_{(1-\alpha)T}\right)^{-1} (\hat{\pi}_{\alpha T} - \hat{\pi}_{(1-\alpha)T})
\]

\(^4\)In the original model of Amihud and Mendelson (1987) \(\pi_{i,t}\) is assumed to be constant for each firms. In this paper we allow it to be time varying which is important for our hypothesis development.

\(^5\)For the alternative estimation methods see Damodaran (1993) and Theobald and Yallup (2004).
and has asymptotic $\chi^2$ distribution under the null. The supremum Wald statistics is calculated as follows

$$SW_T = \sup_{\alpha} W_T(\alpha)$$

$$s.t. \alpha \in \{\alpha_l, \alpha_u\}$$

and the exponential Wald statistic equals to

$$\exp(W_T) = \ln \left\{ \frac{1}{\alpha_u T - \alpha_l T + 1} \sum_{i=0}^{K} \exp\left( \frac{1}{2} W_T(\alpha_l + (\alpha_u - \alpha_l) \frac{i}{K}) \right) \right\}$$

The limiting distributions of these two statistics under the null were derived by Andrews (1993) and Andrews and Ploberger (1994), respectively, who also provide the tables of the corresponding critical values for different levels of significance. More specifically, this procedure involves estimation of a series of usual Wald statistics over a finite grid of $K$ partitioning points of the whole time span of the study. Following Andrews we choose the interval of $\alpha$ equal to $[0.15, 0.85]$ with a grid of 22 trading days (approximately one trading month). This procedure leaves us with 65 potential change points per stock during the time span of our study.

We are interested not only in whether but also how the speed of adjustment changed after the introduction of the SOX and whether this change (if it occurred) has led to an increase in the informational efficiency of stock market. This leads to the following two research questions.

**Q2**: Has the average speed of adjustment for the firms included in our sample increased or decreased following the enactment of SOX reform?

To answer this question the following null hypothesis is tested

$$H_0 : N^{-1} \sum_{i=1}^{N} \pi_{i,S1} = N^{-1} \sum_{i=1}^{N} \pi_{i,S2}$$

with $\pi_{i,S1}$ ($\pi_{i,S2}$) denoting the speed of adjustment of firm $i$ before (after) the SOX act has been signed into law and $N$ is the number of firms in the population. For each cap decile we conduct a pairwise $t$-test where the population averages are replaced by their sample analogues $n^{-1} \sum_{i=1}^{n} \hat{\pi}_{i,S1}$ and $n^{-1} \sum_{i=1}^{n} \hat{\pi}_{i,S2}$ with $n$ denoting the number of firms included in each decile. A significant increase of the cross-sectional average of the speed of adjustment estimates in the post SOX period will indicate that following the reform US markets incorporate more rapidly new information.$^6$

**Q3**: Have the US markets become more efficient following the enactment of the SOX reform?

Since in the A&M model full market efficiency is defined as a case when the speed of adjustment $\pi$ is equal to unity for each stock traded on the market, a natural way of

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$^6$In this study we focus on the mean speed of adjustment as the measure of location of the distribution of the speed of adjustment coefficients. While it is out of scope of this paper, one could also test for the breakpoints in the higher moments of the speed of adjustment distribution, e.g., variance, skewness etc, though the results of these tests would be harder to interpret.
testing for change in market efficiency is by measuring the average distance between the vector of the speed of adjustment parameters and the unity vector. In this study we base our analysis on two different distance measures, the average $L_1$ distance (here $Abs.Dist$) and average squared $L_2$ distance (here $Sq.Dist$) which we define as $N^{-1} \sum_{i=1}^{N} |\pi_i - 1|$ and $N^{-1} \sum_{i=1}^{N} (\pi_i - 1)^2$, respectively, for each cap-portfolio. We test for a change in market efficiency by testing the following null hypotheses via a pairwise $t$-test

$$Abs.Dist_{S1} = Abs.Dist_{S2}$$

$$Sq.Dist_{S1} = Sq.Dist_{S2}$$

where again absolute and squared distance measures are replaced by their sample analogues $n^{-1} \sum_{i=1}^{n} (\hat{\pi}_{i,b} - 1)$ and $n^{-1} \sum_{i=1}^{n} (\hat{\pi}_{i,b} - 1)^2$ for pre-SOX period and $n^{-1} \sum_{i=1}^{n} |\hat{\pi}_{i,a} - 1|$ and $n^{-1} \sum_{i=1}^{n} (\hat{\pi}_{i,a} - 1)^2$, respectively. An increase in market efficiency should result in a significant decline in the distance measures following the reform.

4.2 Analysts’ Performance

A second issue we consider in this paper is the impact of corporate scandals and the SOX reform on the accuracy of the analysts forecasts, in particular, on their bias. The analysis is performed separately for each one of the five variance ratio ranked portfolios, formed as described in subsection 3.2. To set forth notations, we denote by $n^*$ the sample size for each portfolio which is equal to the total number of firms included in each portfolio, that is, $n$ from the previous subsection multiplied by the average number of observation over the time dimension per company. We examine the following research questions

$Q4$: Were the analysts’ consensus forecasts the "best" earnings forecasts both before and after the enactment of the reform?

Denote the actual quarterly earning by $e_{i,t}$ and the consensus forecast by $c_{i,t}$, where $i = 1, \ldots, n$ denotes the index of the company and $t$ denotes the forecast period. We test whether the analysts’ consensus forecasts were the "best" earnings forecasts both in the pre-and post SOX periods, which corresponds to testing the following null hypothesis

$$H_0 : E(e_{i,t} | c_{i,t} = c) = c \forall c \in C$$

where $E(e_{i,t} | c_{i,t} = c)$ denotes expected actual earning conditional on a latest consensus forecast $c$ and $C$ denotes the support of $c_{i,t}$. In particular, under the null, analysts’ consensus forecasts are neither "over pessimistic" nor "over optimistic". We test this hypothesis by using two alternative tests. The first one is the test proposed by Gozalo (1993). In our context, he suggests to use the following statistic

$$T = (n^* h)^{1/2} \bar{v}_c^{-1/2} \left( \frac{\hat{\pi}}{E(e_{i,t} | c_{i,t} = c) - c} \right)$$
Here \( c \) is the value of a randomly selected consensus forecast and \( \hat{E}(e_{i,t} | c_i = c) \) is a non-parametric Nadaraya-Watson estimate of the expected earning conditional on that value of the consensus forecast. Also, \( v_c = \hat{f}(c) \hat{\sigma}_c \int K(\psi)^2 d\psi \) with \( \hat{f}(\cdot) \) being a nonparametric estimate of a density function of \( c_i \) evaluated at the point \( c \), \( K(\cdot) \) is a kernel and \( \hat{\sigma}_c \) is the estimate of the variance of the forecast error \( e_{i,t} = e_{i,t} - \hat{E}(e_{i,t} | c_i = c) \). The choice of the bandwidth \( h \) is based on a cross-validation criterion. Under the null of a correct parametric specification the limit distribution of the test statistic is standard normal. Gozalo (1993) proposes to look at the supremum of \( T \) evaluated at \( d \) randomly chosen points. Since this test is potentially oversized, instead, we calculate the statistic \( T(G) = \sum_{j=1}^{d} T_j^2 \) which is asymptotically \( \chi^2(d) \) distributed under the null (see Pagan and Ullah, 1994 for a comprehensive discussion of estimation and testing in a nonparametric framework).

A second test has been proposed by Stute (1997). He proposes considering the empirical process \( R_n(c_0) = n^{-1/2} \sum_{i=1}^{n} \sum_{t=1}^{T} \epsilon_{i,t} I(c_{i,t} \leq c_0) \) where \( \epsilon_{i,t} = e_{i,t} - c_{i,t} \), that is, the earning surprise under the null of the consensus forecast being the "best" forecast, and using a functional of this process as test statistic. Following Miles and Mora (2003) we consider the Cramer-von Mises statistic
\[
T(S) = n^{* - 2} \sum_{t=1}^{T} \sum_{i=1}^{n} \left[ \sum_{t=1}^{T} \sum_{i=1}^{n} \epsilon_{i,t} I(c_{i,t} \leq c_{j,t}) \right]^{2}
\]
whose asymptotic distribution can be approximated by the "wild bootstrap" procedure (Stute, 1997).

Our interest is not only in whether the analysts’ forecasts were the optimal forecasts before and (or) after the reform, but also whether the former became more or less accurate, following the introduction of the SOX act. Therefore, we examine the following question.

**Q5:** Has there been a structural shift in the conditional dynamics of the analysts’ forecasts errors following the introduction of SOX act, and if so, in which direction?

Testing the intertemporal stability of the conditional dynamics of the analysts’ forecasts errors corresponds to testing the following null hypothesis
\[
E_{S1}(\epsilon_{i,t} | x_{i,t} = x) = E_{S2}(\epsilon_{i,t} | x_{i,t} = x) \quad \forall x \in X
\]
where \( \epsilon_{i,t} \) is the forecast error defined as \( e_{i,t} - c_{i,t}, x_{i,t} \) is a conditioning variable with support \( X \) and \( E_{S1}(\epsilon_{i,t} | x_{i,t} = x) \) and \( E_{S2}(\epsilon_{i,t} | x_{i,t} = x) \) are the conditional expectations of the analysts’ forecast error given \( x_{i,t} \) in pre-and post-SOX periods respectively. The choice of the conditioning variables will be discussed in the following section. To test this hypothesis we use the following Gozalo-type statistic
\[
T_{\text{break}} = (n^* h)^{1/2} \left( \alpha_{S1}^{-1} \hat{v}_{S1,x} + (1 - \alpha_{S1})^{-1} \hat{v}_{S2,x} \right)^{-1/2} \left( \hat{E}_{S1}(\epsilon_{i,t} | x_{i,t} = x) - \hat{E}_{S2}(\epsilon_{i,t} | x_{i,t} = x) \right)
\]
Here \( \hat{E}_{S1}(\epsilon_{i,t} | x_{i,t} = x) \) and \( \hat{E}_{S2}(\epsilon_{i,t} | x_{i,t} = x) \) are the nonparametric estimates of the expected forecast error for a given value of the conditioning variable \( x_{i,t} \) before and after the

---

\( ^7 \)Miles and Mora (2003) also provide a brief description of this procedure.
SOX has been signed into law, $\alpha_{S1}$ partitions our sample into pre-and-post SOX subsamples similar to Andrews (1993) test and $\hat{\sigma}_{S1,i,t}^2(\hat{\sigma}_{S2,i,t}^2)$ are the estimates of its variance $v$ before (after) the SOX act with the same formula as in standard Gozalo test. As in case of the standard Gozalo test we base our inference on the statistic evaluated over $d$ randomly selected points, $T_{(G),break} = \sum_{j=1}^{d} T_{break,j}^2$, which is asymptotically $\chi^2(d)$ distributed under the null. The underlying intuition behind this statistic is that $\hat{E}_{S1}(\epsilon_{i,t} | x_{i,t})$ and $\hat{E}_{S2}(\epsilon_{i,t} | x_{i,t})$ should asymptotically converge to the true conditional expectation which should be the same under the null that no structural shift has occurred. Finally, to test the direction of the structural shift, that is, whether the conditional bias has increased or declined, we test the equality of Spearman rho’s in both periods

$$H_0 : \rho_{S1}(\epsilon_{i,t}, x_{i,t}) = \rho_{S2}(\epsilon_{i,t}, x_{i,t})$$

via standard pairwise $t-$tests.

5 Empirical Results

In this section we present and discuss our empirical results. First, we study the implications of the SOX act for the informational efficiency of the US stock markets within the A&M model framework, as described in a previous section. The results are presented in the subsection 5.1. Next, we turn to the evaluation of the reform’s impact on the analysts’ performance. We present and discuss our findings in subsection 5.2.

5.1 SOX and Market Efficiency

Q 1. Was the speed of adjustment stable over the last decade? If not, when has a structural break occurred?

We start with the estimation results of the A&M model. More specifically, we test for a structural shift of the speed of adjustment coefficient $\pi$ after the reform has been signed on 30 July 2002. While testing for a structural break when the date of a break is assumed to be known is a straightforward procedure, we find it more sensible to start our analysis with testing for parameter stability without determining a change point a priori. The reason for applying this approach is that if the structural shift is detected, it still can be potentially attributed to some other event, such as the collapse of the internet "bubble" or the September 11 events, both of which happened during the time span used in our study. Thus, by allowing the structural breakpoint to be endogenously determined we do not only test for a structural shift but also determine the event which has potentially triggered the latter. To test the stability of the speed of adjustment we apply Andrews supremum Wald (1993) and Andrews and Ploberger (1994) exponential Wald tests as described in a previous section.

Testing results are presented in Table 2. Since it is unrealistic to present the results for each security separately, instead, we present the rejection rates which we define here as a number of securities for which the null of the speed of adjustment stability has been rejected divided by the total number of stocks. Since we partition our sample on monthly basis by using 22 days grid and not on a daily basis (which would be a highly computationally
intensive procedure), the test is likely to be conservative. Therefore, for both supremum Wald and the exponential Wald tests we take the significance level of 10 percent. The rejection rates are calculated for each cap-based portfolio.

<table>
<thead>
<tr>
<th>Decile</th>
<th>1</th>
<th>2</th>
<th>3</th>
<th>4</th>
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</tr>
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<tr>
<td>SW rejection rate</td>
<td>0.19**</td>
<td>0.14</td>
<td>0.29**</td>
<td>0.39**</td>
<td>0.42**</td>
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<tr>
<td>ExpW rejection rate</td>
<td>0.16*</td>
<td>0.17**</td>
<td>0.32**</td>
<td>0.45**</td>
<td>0.47**</td>
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<th>8</th>
<th>9</th>
<th>10</th>
</tr>
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<tbody>
<tr>
<td>SW rejection rate</td>
<td>0.52**</td>
<td>0.46**</td>
<td>0.49**</td>
<td>0.35**</td>
<td>0.37**</td>
</tr>
<tr>
<td>ExpW rejection rate</td>
<td>0.52**</td>
<td>0.48**</td>
<td>0.49**</td>
<td>0.36**</td>
<td>0.35**</td>
</tr>
</tbody>
</table>

Notes: Rejection rate is defined as the number of stocks for which a sup Wald statistic is significant at 10% level divided by a total number of shares. Critical values for sup(W) and exp(W) tests can be found in Andrews (2003) and Andrews and Ploberger (1994). *(**) denotes that the rejection rate is significantly different than 0.1.

The results of the supremum Wald test indicate that for the low-cap stocks, in particular those assigned to the first and second deciles, the null of a stable speed of adjustment is rejected only for a small number of firms. For the second decile the rejection rate statistically does not exceed ten percent, that is, the significance level of the test. On the other hand, the results are strikingly different as we move to higher deciles. Starting from decile 4 we find both statistically and economically significant evidence of a structural shift in the information adjustment mechanism. Rejection rates for deciles six, seven, and eight are especially striking, where a structural break in the speed of adjustment is detected for approximately every second security. The overall conclusion might be that, if the structural break is due to the reform, its impact is substantially more pronounced for the high-cap stocks. This finding can be related to the fact that high-cap stocks, being well known by (or "visible" to) the investors, are more intensively traded than the stocks of the firms with low market value, and, therefore, we would expect the former to respond more rapidly to the reform than the latter. These findings are corroborated by the results of the exponential Wald test where the evidence of a structural shift in a speed of adjustment mechanism is even more pronounced, possibly due to the optimality properties of this test. Overall, based on both tests, we find substantial evidence of the intertemporal instability of the speed of adjustment coefficients.

An interesting question is not only whether, but also when a structural shift has occurred, i.e., whether the structural shift in the speed of adjustment is indeed related to the Sarbanes-Oxley reform and not to some other event. To shed some light on this issue, we plot in Figure 3 the relative frequency of rejections over the time span used in our study. More specifically, for each trading month between March 1999 (which corresponds approximately to \( \alpha = 0.15 \)) and October 2004 (which corresponds to \( \alpha = 0.85 \)), we define the relative share of rejections which occurred during that particular month out of the total number of rejections. The results clearly show a substantial cluster of the structural breaks occurring in 2002, the year of the reform, with two spikes around June-July 2002 and August-September 2002.
Figure 3: Relative frequency of the structural shifts in the speed of adjustment coefficients for the NYSE/AMEX stocks

Interestingly, we also find a cluster (though of less substantial magnitude) of the structural breaks around March 1999-April 2000, a finding that can be related to the collapse of the "dot.com" bubble. Also, we find a single spike of structural breaks the week following September 11, 2001. Overall, the findings of the Andrews (1993) stability test clearly indicate both instability of the speed of adjustment mechanism and the reform of 2002 being a potential source of the former.

Q2 and Q3. Has the average speed of adjustment for the firms included in our sample increased or decreased following the enactment of SOX reform? Have the US markets become more efficient following the enactment of the SOX reform?

We now turn to the evaluation of the impact, the reform of 2002 had on the speed of adjustment of the US stock market to a new information. Based on the findings of Andrew's (1993) stability test, the most substantial cluster of structural shifts has occurred in 2002. Therefore, in the following analysis we split our sample into two subsamples: January 1998-December 2001 and January 2002-December 2005 which we denote as S1 and S2 respectively. For each security included in our sample, we estimate the speed of adjustment coefficient \( \pi \) during the first and second sub-period and report its cross-sectional average for each of ten cap-sorted portfolios. We apply the GMM theory of Hansen (1982) where as instruments we choose Fama-French factors. The following set of moments is used to estimate the speed of adjustment for each security

\[
E \left\{ \begin{array}{l}
MKT_{t-1} \\
SMB_{t-1} \\
HML_{t-1}
\end{array} \begin{array}{l}
(\Delta P_{t,t} - \pi_i \mu_i + (1 - \pi_i) \Delta P_{t,t-1})
\end{array} \right\} = 0
\]
Here $MKT_t$ denotes the return on market portfolio, $SMB_t$ denotes the return on a portfolio of small size stocks minus high capitalization stocks portfolio, and $HML_t$ denotes the return on "value" stocks portfolio minus "growth" stocks portfolio. Daily data on these factors has been obtained from CRSP. Our preliminary analysis suggests that for our sample these factors on average capture above 20 percent of the total variation of stock returns while on the other hand we would expect these factors to be uncorrelated with bid-ask spread related noise of the individual stock, properties that make these factors suitable instruments.

The results are presented in Table 3. A number of interesting findings can be noted. First, consistent with findings of Theobald and Yallup (2004) we find that on a daily basis investors tend to underreact to the news with most of the average estimates of $\pi$ being significantly lower than unity. Also, we find that on average the extent to which investors underreact tends to decline from low to high-cap portfolios, suggesting that high-cap stocks tend to reflect more rapidly new information, compared to the ones with low market capitalization. This finding is consistent with a "lead-lag" effect (e.g. Chordia and Swaminathan, 2001). Also, for the low-cap portfolios the estimates of the speed of adjustment tend to exhibit a higher cross-sectional dispersion.

Next, we turn to the main issue, namely, the impact of the Sarbanes-Oxley reform on the speed of adjustment and market efficiency. Starting with low cap portfolios, in particular, the first decile, we find that, though on average the speed of adjustment to the information shocks has increased, it lacks statistical significance. The picture is similar for the distance measures, both of which have declined after the reform, but the difference between the distance measures in the pre- and post-SOX periods lacks statistical significance. However, starting from second decile the results change dramatically. Starting with the analysis of the adjustment coefficients, we find a dramatic increase in the former, an increase which is both economically and statistically significant. An increase in the speed of adjustment coefficients is becoming more pronounced as we move from the low to high cap stocks. Our results indicate that, following the reform on July 2002, the speed of adjustment to the information shocks increased on average by more than 15%, suggesting that after the reform, the US stock market responds more rapidly to the new information. The analysis of the two distance measures supports these findings. Similarly to the speed of adjustment estimates, starting from second cap-decile, we find both an economically and statistically significant decrease in both the absolute and squared distance measures. Our findings suggest that on average the Sarbanes-Oxley reform resulted in a decrease of more than 36% of the mean absolute distance, while a mean squared distance declined almost twice! Overall, these findings suggest that based on the A&M model of partial adjustment, the reform of 2002 indeed led to an increase in market efficiency.

\footnote{Since in our study the time dimension $T$ is substantially larger than the dimension of the cross-section $N$ we assume that $T$ is of the higher order than $N$. Under this assumption the estimation inaccuracy of the $\hat{\pi}$s is asymptotically irrelevant.}
Table 3: Speed of adjustment estimates—A&M model

<table>
<thead>
<tr>
<th>Decile</th>
<th>1</th>
<th>2</th>
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<th>6</th>
<th>7</th>
<th>8</th>
<th>9</th>
<th>10</th>
</tr>
</thead>
<tbody>
<tr>
<td>Sample</td>
<td>S1</td>
<td>S2</td>
<td>S1</td>
<td>S2</td>
<td>S1</td>
<td>S2</td>
<td>S1</td>
<td>S2</td>
<td>S1</td>
<td>S2</td>
</tr>
<tr>
<td>Speed of Adj.</td>
<td>0.518**</td>
<td>0.589**</td>
<td>0.449**</td>
<td>0.474**</td>
<td>0.648**</td>
<td>0.755**</td>
<td>0.706**</td>
<td>0.853**</td>
<td>0.776**</td>
<td>0.939**</td>
</tr>
<tr>
<td>p-value</td>
<td>0.4</td>
<td>0.09</td>
<td>0.01</td>
<td>0.00</td>
<td>0.00</td>
<td>0.245</td>
<td>0.168</td>
<td>0.092</td>
<td>0.055</td>
<td></td>
</tr>
<tr>
<td>Abs.Dist</td>
<td>0.519</td>
<td>0.476</td>
<td>0.595</td>
<td>0.527</td>
<td>0.426</td>
<td>0.286</td>
<td>0.331</td>
<td>0.23</td>
<td>0.245</td>
<td>0.168</td>
</tr>
<tr>
<td>p-value</td>
<td>0.57</td>
<td>0.00</td>
<td>0.00</td>
<td>0.00</td>
<td>0.00</td>
<td>0.00</td>
<td>0.00</td>
<td>0.00</td>
<td>0.00</td>
<td>0.00</td>
</tr>
<tr>
<td>Sq.Dist</td>
<td>0.619</td>
<td>0.328</td>
<td>0.429</td>
<td>0.297</td>
<td>0.285</td>
<td>0.179</td>
<td>0.151</td>
<td>0.107</td>
<td>0.092</td>
<td>0.055</td>
</tr>
<tr>
<td>p-value</td>
<td>0.36</td>
<td>0.00</td>
<td>0.09</td>
<td>0.00</td>
<td>0.02</td>
<td>0.03</td>
<td>0.00</td>
<td>0.00</td>
<td>0.00</td>
<td>0.00</td>
</tr>
</tbody>
</table>

Notes: Speed of Adj. denotes sample mean estimate of $\pi$, Sq.Dist and Abs.Dist denote average squared and average absolute distance of $\pi$ from the unity vector. We test the null: Speed of Adj.$(S_1)$ = Speed of Adj.$(S_2)$ Abs.Dist.$(S_1)$ = Abs.Dist.$(S_2)$ and Sq.Dist.$(S_1)$ = Sq.Dist.$(S_2)$. Corresponding $p$-values are reported below.

We conduct a number of robustness checks. The results of Andrews (1993) stability test suggest that a substantial number of structural breaks can be potentially attributed to the collapse of the "dot.com" bubble and not to the SOX reform. Therefore, as a first robustness test we reestimate the speed of adjustment coefficients for the first subsample from which we exclude the period between March 1999 and March 2000. Overall, the results remain unaltered with cap-decile 7 being the only exception where an increase in the speed of adjustment turned out to be statistically insignificant. Next, we test whether our results could be due to the fact that we used only the stocks with continuous data over the whole sample period. If investors learn over time, one would expect to observe a gradual increase in the speed of adjustment, which could be mistakenly attributed to the impact of the reform. While indeed for a number of portfolios the speed of adjustment coefficients exhibit a significant time trend, the latter is significantly negative, a finding that suggests that our results are not driven by a selection bias problem. The investors’ hypothetical learning curve also fails to explain a large cluster of structural breaks around the reform date. Finally, we reestimate the speed of adjustment coefficients for the second subsample from which we excluded the period between January and December 2002, the "breakpoint year" when the reform has been signed. An increase in the speed of adjustment coefficients becomes even more pronounced and also becomes statistically significant for the second decile. An overall impression is that our findings are fairly robust.
5.2 SOX and the Analysts’ Performance

Next, we turn to the analysis of the analysts' performance in the pre- and post Sarbanes-Oxley periods. All firms are grouped into five portfolios based on a variance ratio criterion, as described in subsection 3.2. The bias-variance trade-off models of Das et al. (1998) and Lim (2001) suggest that the analysts issue intentionally biased forecasts in order to improve access to managers' private information. Consequently, an improved access to managers' private information would result in a lower variance ratio. Also, based on the underlying logic of the bias-variance trade-off models, the analysts' forecasts of the companies with low variance ratio are more likely to be biased, i.e. overoptimistic if the managers prefer overoptimistic forecasts, or overpessimistic if the firms' management follows "meat or beat" analysts' forecast tendency. Therefore, if the analysts covering the low variance ratio companies are more dependent on access to managers' private information, we would expect the accuracy of their forecasts to be more affected by the new disclosure rules on the one hand, and the corporate scandals on the other hand. In other words, if the structural shift in the analysts' forecasts accuracy is detected in the post-SOX act period, we would expect it to be more pronounced for the low variance-ratio firms.

Q.4. Were the analysts’ consensus forecasts the "best" earnings forecasts both before and after the enactment of the reform?

We begin with the analysis of the expected forecast error, conditional on the consensus forecast, which we define as the difference between the conditional expected earning given the consensus forecast for a particular quarter and the consensus forecast. Consider the following decomposition of the "best" earnings forecast

\[
E(e_{i,t} | c_{i,t} = c) = c + E(e_{i,t} | c_{i,t} = c) - c
\]

The last term at the right hand side of the equation can be considered as the correction factor, one can use to improve the performance of the consensus forecast. For each sub-period we test the null hypothesis that the correction factor is equal to zero or, in other words, that the consensus forecast is the "best" forecast. As discussed in the previous section, we base our analysis on both nonparametric (Gozalo, 1993) and conditional moments type (Stute, 1997) tests. Since nonparametric tests typically are substantially sensitive to the choice of bandwidth, using the "nonsmoothing" type of tests such as the one proposed by Stute (1997) is a useful check of the robustness of our results.
Table 4: Expected forecast error tests

<table>
<thead>
<tr>
<th>Quintile</th>
<th>Pre-SOX</th>
<th>Post-SOX</th>
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<tr>
<td></td>
<td>$T(G)$</td>
<td>$T(S)$</td>
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<tr>
<td>1</td>
<td>0.000</td>
<td>0.000</td>
</tr>
<tr>
<td>2</td>
<td>0.001</td>
<td>0.003</td>
</tr>
<tr>
<td>3</td>
<td>0.001</td>
<td>0.002</td>
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<tr>
<td>4</td>
<td>0.002</td>
<td>0.000</td>
</tr>
<tr>
<td>5</td>
<td>0.67</td>
<td>0.078</td>
</tr>
</tbody>
</table>

The null: $E(e_{i,t} | c_{i,t} = c ) = c \forall c \in C$

The numbers are the $p$-values of Gozalo ($T(G)$) and Stute ($T(S)$) statistics. For $T(G)$ the choice of bandwidth is based on cross-validation. The significance of $T(S)$ is tested by using "wild bootstrap" with 500 replications. Each test is conducted for both pre- and post SOX period for each portfolio.

Q.5. Has there been a structural shift in the conditional dynamics of the analysts’ forecasts errors following the introduction of SOX, and if so, in which direction?

Analysts’ performance- a visual inspection

To gain some preliminary insight into the dynamics of the analysts’ forecast errors before and after the enactment of the SOX reform we plot the estimated correction factor, $\hat{E}(e_{i,t} | c_{i,t} = c ) - c$, versus the consensus forecast $c$ for each value of $c$ lying in the range of $(q_{c,0.05}, q_{c,0.95})$ in Figures 4 to 6, with $q_{c,0.05}$ ($q_{c,0.95}$) denoting the 5% (95%) estimated quantile of the consensus forecasts distribution. The estimated correction factor is presented for the groups of variance ratio-ranked portfolios, where for each group we present the estimated correction factor before and after the SOX has been signed into law separately along with the corresponding uniform confidence bands. Starting with the analysis of the correction factor for the low variance-ratio firms, which is presented in the upper plot of Figure 4, it appears that there has been a structural shift in the forecast error-consensus forecast relationship. Surprisingly, however, it appears that the magnitude of the correction factor has increased, suggesting that in the post-SOX act period the analysts’ forecasts on average became less accurate. In addition, two interesting findings should be mentioned. First, the correction factor became more positive, suggesting that the analysts became more overpessimistic, or cautious. Secondly, it appears that in the post-SOX period the correction
factor is substantially more pronounced for the positive forecasts, i.e., we would expect the analysts' forecasts errors to be more positive when the firm's management is reporting "good" news. An inspection of the behavior of the expected forecast error for the high variance ratio firms leads to similar conclusions, with the quintile 3 portfolio being the only exception, where it seems to be hard to reach any specific conclusions regarding the change in direction of the bias. However, the shift in the correction factor is substantially more pronounced for the low variance-ratio stocks, that is, for the firms with more pronounced bias-variance trade-off by the analysts, suggesting that the analysts following these firms became substantially more cautious when the firms' management discloses "optimistic" information.

To study further the nature of the bias we are also interested in studying the relationship between the expected forecast error and the forecast uncertainty, and, in particular, whether this relationship has undergone any kind of structural shift, following the series of corporate scandals and the SOX legislation. Based on the underlying logic of bias-precision trade-off models we would expect the magnitude of the forecast errors to increase during the periods of higher uncertainty regarding the future earnings since the analysts are more likely to request additional information from the firm's management when the earnings exhibit a high degree of variation. Moreover, if the source of the structural shift is the SOX act, which made the information regarding the financial and economic conditions of the firm more accessible, we would expect to see a link between the forecast errors and the forecast uncertainty becoming weaker, since due to an increase in the information transparency we would expect the analysts forecasts' to become less dependent on the information provided by the management. On the other hand, it is also possible that a structural shift (if detected) is due to the corporate scandals which were related to the misreporting and the earning manipulations of the management, such as the Enron and Worldcome inquiries. In this case we would expect the analysts to become more cautious in their forecasts, which could lead to an increase in the forecast errors. Also, the link between the forecast errors and the forecast uncertainty is expected to become stronger, since, following the scandals, we would expect the analysts to become more cautious during the turbulent periods compared to the pre-SOX period.

Figures 7 to 9 provide some visual impression on the relationship between the analysts' forecast errors and the earnings forecast uncertainty, where, as a proxy of uncertainty, we take the analysts' forecast spread $s_{i,t}$, which we define as the highest forecast minus the lowest forecast for the last month before the actual earning is announced for each firm-quarter observation. As before, we plot a nonparametric estimate of the expected forecast error, $\hat{E}(\epsilon_{i,t} | s_{i,t} = s)$, for each value of $s$ lying in the range of $(q_{s,0.05}, q_{s,0.95})$ for the variance-ratio ranked portfolios. A number of interesting findings should be mentioned. First, by observing the overall level of the analysts' forecasts errors it seems that the latter was positive, that is, in both periods, on average, analysts tended to submit overpessimistic forecasts. This finding is consistent with the results obtained from the inspection of Figures 4-6. Interestingly, the magnitude of the forecast errors appears to be related to the level of the uncertainty regarding the future earnings, measured by the forecasts spread, a finding which is consistent with the results reported by Imhoff and Lobo (1990). Moreover, we find that the forecast error is likely to increase during the turbulent periods, a finding, which cannot be attributed to the bias-variance trade-off model of Lim (2001), who assumes that
firms’ managers prefer overoptimistic forecasts which would lead to the forecast errors being negative during the periods of high earnings uncertainty. On the other hand this finding is consistent with the bias-precision trade-off model if the firm’s management desire is to meet or beat analysts forecasts, as reported by Degeorge et al. (1999). This positive link appears to be more pronounced for the lower quintile portfolios, which is also consistent with the predictions of the bias-variance trade-off models with ‘meat or beat analysts forecasts’ management strategy.

However, the most intriguing results come from the comparison of the forecast error-forecast spread link in both periods. First, there is a clear visual evidence of a structural shift in the overall level of the forecast error. More specifically, we find that the overall level of the analysts’ forecast errors has experienced an upward shift, a finding, which holds both for the low and high variance-ratio firms. The magnitude of the observed shift is indeed striking, ranging from 50 percent for the lower quintile to more than 100 percent for the upper quintile portfolios. Moreover, it seems that there has also been a structural change in the bias-spread relationship. Comparing the forecast error-forecast spread plots for the pre-and post-SOX periods we find that the analysts’ forecast errors became substantially more positively related to the earnings uncertainty. This shift is also substantially more pronounced for the low-quintile portfolios, for which the estimated forecast error-forecast spread curve turns from being moderately increasing or almost flat to substantially increasing. The shift in forecast error-forecast uncertainty link, however, is also detected for the high variance ratio portfolios, where the estimated forecast error-forecast spread curve, which in the pre-SOX period exhibited an inverted "U" shape, turned to be substantially increasing in spread.

**Analysts’ performance-formal tests**

We start with a Gozalo-type nonparametric test described in Section 4. We test for the presence of structural shifts by using two different conditioning variables: the consensus forecast $c$ and the forecast spread $s$. As for the standard Gozalo test, we calculate the value of the statistic for ten randomly chosen values of $c$ over the range of $(q_{c,0.05}, q_{c,0.95})$ and for ten randomly chosen values of $s$ over the range of $(q_{s,0.05}, q_{s,0.95})$. 

22
Table 5: Structural shift tests

\[ H_0 : E_{S1}(\epsilon_{i,t} | c_{i,t}) = E_{S2}(\epsilon_{i,t} | c_{i,t}) \]

\[ H_0 : E_{S1}(\epsilon_{i,t} | s_{i,t}) = E_{S2}(\epsilon_{i,t} | s_{i,t}) \]

<table>
<thead>
<tr>
<th>Quintile</th>
<th>( T_{(G)} )</th>
<th>p-value</th>
<th>min / max</th>
<th>( \text{rho}_{S1} )</th>
<th>( \text{rho}_{S2} )</th>
<th>t.stat</th>
<th>p-value</th>
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<tr>
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<td>0.0002</td>
<td>[0.02, 13.47]</td>
<td>0.07**</td>
<td>0.194**</td>
<td>3.98</td>
<td>0.000</td>
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<tr>
<td>2</td>
<td>19.76</td>
<td>0.031</td>
<td>[0.08, 5.63]</td>
<td>0.134**</td>
<td>0.184**</td>
<td>1.59</td>
<td>0.11</td>
</tr>
<tr>
<td>3</td>
<td>17.39</td>
<td>0.06</td>
<td>[0.009, 5.82]</td>
<td>0.087**</td>
<td>0.191**</td>
<td>3.26</td>
<td>0.001</td>
</tr>
<tr>
<td>4</td>
<td>19.35</td>
<td>0.036</td>
<td>[0.002, 10.62]</td>
<td>0.056**</td>
<td>0.179**</td>
<td>4.14</td>
<td>0.000</td>
</tr>
<tr>
<td>5</td>
<td>42.51</td>
<td>0.0000</td>
<td>[0.095, 13.95]</td>
<td>0.074**</td>
<td>0.189**</td>
<td>3.72</td>
<td>0.000</td>
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</tbody>
</table>

<table>
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<tr>
<th>Quintile</th>
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<th>p-value</th>
<th>min / max</th>
<th>( \text{rho}_{S1} )</th>
<th>( \text{rho}_{S2} )</th>
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<td>0.000</td>
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<td>0.000</td>
<td>[0.77, 23.64]</td>
<td>0.046**</td>
<td>0.108**</td>
<td>1.94</td>
<td>0.05</td>
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<tr>
<td>3</td>
<td>95.57</td>
<td>0.000</td>
<td>[4.92, 13.11]</td>
<td>0.007</td>
<td>0.062**</td>
<td>1.74</td>
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<td>67.54</td>
<td>0.000</td>
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<td>0.03</td>
<td>0.06**</td>
<td>0.98</td>
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<td>5</td>
<td>103.86</td>
<td>0.000</td>
<td>[0.02, 22.21]</td>
<td>0.005</td>
<td>0.07**</td>
<td>1.95</td>
<td>0.05</td>
</tr>
</tbody>
</table>

min/max denote minimum/maximum value of ten randomly selected Gozalo-type statistics; \( \text{rho}_{S1} \) and \( \text{rho}_{S2} \) denote Spearman’s rho estimates for pre and post-SOX periods

Spearman’s rho standard errors’ estimates are based on 1000 Monte-Carlo replications

** denotes significance at 5% level

The results are presented in Table 5. For each variance ratio ranked portfolio we test two separate null hypotheses: \( H_0 : E_{S1}(\epsilon_{i,t} | c_{i,t} = c) = E_{S2}(\epsilon_{i,t} | c_{i,t} = c) \) and \( H_0 : E_{S1}(\epsilon_{i,t} | s_{i,t} = s) = E_{S1}(\epsilon_{i,t} | s_{i,t} = s) \). By testing the first hypothesis, we test for the presence of a structural break in the forecast error-consensus forecast relationship, while by testing the second one we test for the stability of the forecast error-forecast uncertainty link. We start with the analysis of the forecast error-consensus forecast relationship, with the results presented in the upper panel of Table 5. Our results strongly suggest the presence of a structural shift in the forecast error-consensus forecast linking function after the SOX has been signed into law. The null of intertemporal stability is strongly rejected for the first two quintile portfolios, with somewhat weaker, though still significant, evidence of a structural break for the upper quintiles. Overall, our findings suggest that a structural shift has occurred and that the latter appears to be more pronounced for the low variance-ratio firms, that is, the firms with more pronounced bias-variance trade-off by the analysts.

To provide an additional insight into the nature of this shift, for each quintile we calculate the estimates of Spearman’s rho between the forecast error and the corresponding consensus forecast, for both the pre- and post-SOX act sub samples. The estimates are presented in the sixth and seventh columns of the upper panel of Table 5. Starting with the pre-SOX period the results suggest that there exists a statistically significant and positive dependence between the forecast error and the consensus forecasts, suggesting that, on average, it is more likely to observe a positive earning surprise when the analysts submit high earnings estimates. In other words, the forecasts are relatively more "overpessimistic" when the firm management issues positive reports regarding the economic and financial fundamentals of the firm. This can be attributed either to the earnings management or to the analysts being more cautious when they receive too optimistic reports. However, the low magnitude of the estimates suggests that the economic significance of this dependence is somewhat limited, with the second quintile being the only exception. The results dramatically change in the
post-SOX period, where for all portfolios Spearman’s rho experienced a sharp increase in its value. The magnitude of increase is indeed dramatic, ranging between 40 percent for the second quintile and more than 200 percent for the 4-th quintile.

Next, we study the results of the structural shift test for the forecast error-forecast uncertainty link. The results of the test using the critical values of the $\chi^2_{10}$ distribution suggest that the null of the intertemporal stability is strongly rejected for any reasonable significance level. This confirms the visually based findings, which indicated an upward shift both in the overall level of the forecast errors and in the forecast errors-forecast spread link. These findings gain an additional support from the analysis of Spearman’s rho estimates between the earnings surprise and the forecast spread. For the pre-SOX period only for the first two quintiles we find a significant and positive dependence between the bias and the uncertainty, while for the rest of the firms the estimates are neither statistically nor economically significant. However, our findings dramatically change as we move to the post-SOX period, where for all the quintiles we find a statistically significant and positive relationship between the earnings surprise and the spread. As with the bias-consensus link, the changes in Spearman’s rho are indeed striking and are especially pronounced for the lower quintile portfolios.

We formally test the null $\rho_{S1} = \rho_{S2}$ via standard pairwise $t$-test. This test requires an estimate of the asymptotic standard errors of the estimators of both $\rho_b$ and $\rho_a$. Though a closed and compact formula for the asymptotic variance exists, it is hardly suitable for practical applications, since it requires estimation of high-dimensional integrals (see, for instance, Schmid and Schmidt, 2006). Instead, these authors propose to use bootstrap based estimates, an approach we shall adopt in this study as well. In the last two columns of Table 5 we report the test statistics and the corresponding $p$-values for each quintile. Overall, our findings indicate that the forecast error-consensus forecast and forecast error-forecast uncertainty links became both statistically and economically significant in the post-SOX period. These findings suggest that, following a series of corporate scandals, analysts became substantially more cautious in forming their forecasts.

5.2.1 SOX or Regulation FD - a Robustness Check

While we find a substantial evidence of structural shifts in the analysts’ forecasts bias in the post-SOX period, the remaining question is whether this shift is due to the corporate scandals or, perhaps, it can be attributed to some other event which occurred during our sample period. A natural candidacy for such an event is the Regulation Fair Disclosure (FD) act enacted on October 23, 2000. Regulation FD prohibits corporations from privately disclosing material information to a subset of investors or securities markets professionals, e.g. analysts, without simultaneously disclosing the same information to the public. Since the implementation of FD is likely to be associated with changes in the earnings-related information environment, this legislation could also lead to structural shifts in forecasting performance of the analysts. Thus, to examine the robustness of our results we conduct the same tests with the alternative partitioning of our sample period. More specifically, we partition the whole sample into three sub-samples: January 1998-October 2000 (encoded pre-FD), November 2000-July 2002 (encoded post-FD/pre-SOX), August 2002-June 2006 (encoded post-SOX). By partitioning the pre-SOX period into pre-FD and post-FD/pre-SOX periods...
we seek to disentangle the impact of Regulation Fair Disclosure act from the potential effect of the SOX act and the preceding corporate scandals. For each sample period we conduct the same structural shift tests we conducted for the two sub-sample partitioning. In particular, while testing for the structural shift we test separately for the presence of the structural shift between the pre-FD and post-FD/pre-SOX periods, and between post-FD/pre-SOX and the post-SOX periods. A structural break in the analysts' forecast bias between pre-FD and post-FD/pre-SOX periods can be attributed to the impact of FD regulation, while a shift in the forecast bias between the post-FD/pre-SOX and post-SOX periods is likely to be attributed to the impact of the SOX legislation.

We start with a visual inspection of the forecast error-consensus forecast and the forecast error-forecast spread plots depicted in Figures 10-15. Each figure depicts a nonparametric estimate of the forecast error-consensus forecast or the forecast error-forecast spread relation for the pre-FD, post-FD/pre-SOX and post-SOX periods. A number of interesting findings should be mentioned. First, there is an upward shift in the overall level of the analysts' forecast bias, a finding which suggests that over time analysts became more "over-pessimistic" regarding future firms' earnings. The shift in the overall level of bias, that is, the shift of the "intercepts" of the estimated curves, is pronounced both for the forecast error-consensus forecast and the forecast error-forecast spread links. This is consistent with findings of Brown (2001), who reports a similar trend in the earnings surprise, though for the earlier period. Second, no visual shift in the forecast error-consensus forecast link can be detected between pre-FD and post-FD/pre-SOX periods. The findings are similar for the forecast error-forecast spread link. On the other hand, we do find some evidence of a structural shift in both forecast error-consensus forecast and forecast error-forecast spread links between the post-FD/pre-SOX and post-SOX periods. For the forecast error-forecast spread link a shift is more pronounced for the firms with low variance ratio, which emphasizes, though informally, the need to control for the bias-variance trade-off. Overall, visual inspection informally suggests that the structural shifts detected in the analysts' bias seem to occur in the post-SOX period and are unlikely to be attributed to the Regulation FD.
Table 6: Structural shift tests - 3-period partition

Table 6.1: Forecast error-consensus forecast tests

<table>
<thead>
<tr>
<th>Quintile</th>
<th>$T(G)$</th>
<th>p-value</th>
<th>min / max</th>
<th>$T(G)$</th>
<th>p-value</th>
<th>min / max</th>
</tr>
</thead>
<tbody>
<tr>
<td>1</td>
<td>16.31</td>
<td>0.09</td>
<td>[0.44,8.59]</td>
<td>40.39</td>
<td>0.0001</td>
<td>[0.073,15.84]</td>
</tr>
<tr>
<td>2</td>
<td>4.65</td>
<td>0.91</td>
<td>[0.02,1.48]</td>
<td>15.65</td>
<td>0.11</td>
<td>[0.14,4.66]</td>
</tr>
<tr>
<td>3</td>
<td>20.46</td>
<td>0.025</td>
<td>[0.006,8.45]</td>
<td>32.67</td>
<td>0.0003</td>
<td>[0.09,6.83]</td>
</tr>
<tr>
<td>4</td>
<td>21.34</td>
<td>0.02</td>
<td>[0.009,3.82]</td>
<td>43.51</td>
<td>0.0000</td>
<td>[0.002,16.39]</td>
</tr>
<tr>
<td>5</td>
<td>6.31</td>
<td>0.79</td>
<td>[0.02,2.57]</td>
<td>53.15</td>
<td>0.0000</td>
<td>[0.18,11.79]</td>
</tr>
</tbody>
</table>

Table 6.2: Forecast error-forecast spread tests

<table>
<thead>
<tr>
<th>Quintile</th>
<th>$T(G)$</th>
<th>p-value</th>
<th>min / max</th>
<th>$T(G)$</th>
<th>p-value</th>
<th>min / max</th>
</tr>
</thead>
<tbody>
<tr>
<td>1</td>
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<td>0.0000</td>
<td>[1.38,18.31]</td>
<td>116.49</td>
<td>0.0000</td>
<td>[10.39,13.56]</td>
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<tr>
<td>2</td>
<td>2.62</td>
<td>0.98</td>
<td>[0.09,1.23]</td>
<td>139.48</td>
<td>0.0000</td>
<td>[3.84,15.62]</td>
</tr>
<tr>
<td>3</td>
<td>17.52</td>
<td>0.063</td>
<td>[0.68,3.11]</td>
<td>113.67</td>
<td>0.0000</td>
<td>[3.48,19.59]</td>
</tr>
<tr>
<td>4</td>
<td>37.66</td>
<td>0.0001</td>
<td>[1.78,5.74]</td>
<td>119.41</td>
<td>0.0000</td>
<td>[8.16,14.94]</td>
</tr>
<tr>
<td>5</td>
<td>16.87</td>
<td>0.08</td>
<td>[0.06,3.56]</td>
<td>178.01</td>
<td>0.0000</td>
<td>[11.62,21.8]</td>
</tr>
</tbody>
</table>

In this table we present the results of the Gozalo-type structural shift tests for the 3-period sample partition. We test for the structural shifts in the post-FD/pre-SOX and post-SOX periods. min/max denote minimum/maximum value of ten randomly selected Gozalo-type statistics. $T(G)$ denotes the value of statistic.

Next, we turn to the formal structural shift tests. The results of the nonparametric Gozalo-type test are presented in Table 6. The upper panel of Table 6, Table 6.1, presents the results of the structural shift test applied to the forecast error-consensus forecast relation. The results of the structural shift tests for the pre-FD versus post FD/pre-SOX periods suggest that there was a structural shift in the forecast error-consensus forecast relation for the third and forth quintiles. There is also strong evidence of a structural shift following the enactment of the SOX for all but second quintile, where the results are marginally significant. The results of the structural shift tests applied to the forecast error-forecast spread relation, reported in Table 6.2, depict a similar picture. There is a significant evidence of a structural shift in both post-FD/pre-SOX and post-SOX periods. However, these shifts can also be attributed to the overall shift in the analysts’ forecast bias. Thus, to study further the nature of these structural shifts we study the estimates of Spearman’s rhos.
In this table we present the results of the structural shift test for the forecast error-consensus forecast relation and forecast error-forecast spread relation for the 3-period sample partition. \( \rho_1 \), \( \rho_2 \), and \( \rho_3 \) are the Spearman rho estimates for the pre-FD, post-FD/pre-SOX and post-SOX periods. The following null hypotheses are tested: 

\[
H_0: \rho_1 = \rho_2, \\
H_0: \rho_2 = \rho_3
\]

where \( t_{\rho_1=\rho_2} \) and \( t_{\rho_2=\rho_3} \) are the corresponding t-statistics respectively. The estimates of Spearman's rhos as well as the t-statistics of the structural shift tests and the corresponding p-values are reported in Table 7. Here \( \rho_1 \), \( \rho_2 \), and \( \rho_3 \) denote Spearman's rho estimates for the pre-FD, post-FD/pre-SOX and post-SOX periods respectively. First, consistent with the results of the nonparametric tests reported in Table 6 we find some evidence of structural shifts in the forecast error-consensus forecast link in the post FD/pre-SOX period. Secondly, and more important, we find a strong evidence of a structural shift in the forecast error-consensus forecast relation in the post-SOX period. More specifically, we find the upward shift in the estimates of Spearman's rho, a shift which is both statistically and economically significant for all the quintiles. Turning to the forecast error-forecast spread link, we find almost no evidence of a structural shift with a first quintile being the only exception. On the contrary, we find an increase in Spearman's rho estimates in the post-SOX period for all the quintiles. For the first two quintiles the shift is also statistically significant. Overall, these results support the robustness of our findings, that there has been a structural shift in the analysts’ forecast bias following the enactment of the SOX.

### Conclusions and Topics for Further Research

In this paper we examine the implications of the Sarbanes-Oxley reform of 2002 for the informational efficiency of the US stock market and the performance of the stock market analysts. To the best of our knowledge, this is the first paper that addresses these issues and the questions raised and studied in this paper are of major importance, both for the policy makers and practitioners.

To study the impact of the SOX reform on the informational market efficiency we estimate the partial adjustment model with noise of Amihud and Mendelson (1987) for all
the firms listed on the NYSE/AMEX during the last decade. By applying an endogenous structural break tests of Andrews (1993) and Andrews and Ploberger (1994), we find 2002 to be the year when a lions share of the structural breaks in the speed of adjustment occurred. Further tests indicate that, following the enactment of the SOX reform, the average speed of adjustment to the new information has substantially increased, suggesting that, following the legislation, investors incorporate more rapidly the information released by the firms in the stock prices, thus, making the US stock market more informationally efficient.

We also study the implications of the Act of 2002 for the accuracy of the analysts’ forecasts. By applying nonparametric tests to the large span of I/B/E/S forecasts and actual earnings data, we find that both before and after the reform the analysts’ forecasts were significantly "overpessimistic". Moreover, we find that in the post-SOX period the degree of the "overpessimism" has not declined, but rather increased. An increase is especially pronounced during the turbulent periods and/or when "good" news is released. These findings suggest that an increase in the magnitude of the forecast errors can be attributed to the analysts becoming more cautious following the series of corporate scandals when severe earnings overestimations were uncovered.

Our findings also propose a number of promising directions for further research. First, it may be interesting to dichotomize the analysts’ forecasts into those submitted by the analysts who work for the underwriting firms and those who do not. Several studies find that the former are, in general, "overoptimistic", and, therefore, it is quite possible that their overoptimism on the one hand, will be balanced by the impact of the corporate scandals on the other hand, leading to an increase in the analysts’ forecasts accuracy. Secondly, our findings suggest that the stock market investors became more efficient in incorporating new information while the picture is reverse for the analysts, suggesting that the importance of the analysts’ forecasts for the investors’ valuation of the firm, in particular, using the consensus forecasts as a proxy for the investors (market) expectations, seems to be overstated. An alternative explanation is that the investors are sophisticated enough to correct for the analysts’ bias, which seems to be a promising research direction. Finally, our results suggest that while in a short run investors are underreacting, the degree of the underreaction following the reform of 2002 seems to decline. Therefore, it will be interesting to compare the profitability of "momentum" based strategies both before and after the reform has been signed into law.
Figure 4: Forecast error vs consensus forecast before and after the enactment of the SOX. Dashed lines denote 95% uniform confidence bands.
Figure 5: Forecast error vs consensus forecast before and after the enactment of the SOX (continued)
Figure 6: Forecast error vs consensus forecast before and after the enactment of the SOX (continued)
Figure 7: Forecast error vs forecast spread before and after the enactment of the SOX. Dashed lines denote 95% uniform confidence bands.
Figure 8: Forecast error vs forecast spread before and after the enactment of the SOX (continued)
Figure 9: Forecast error vs forecast spread before and after the enactment of the SOX (continued)
Figure 10: Forecast error vs consensus forecast - 3-period partition. Dashed lines denote 95% uniform confidence bands.
Figure 11: Forecast error vs consensus forecast - 3-period partition (continued)
Figure 12: Forecast error vs consensus forecast - 3-period partition (continued)
Figure 13: Forecast error vs forecast spread - 3-period partition. Dashed lines denote 95% uniform confidence bands.
Figure 14: Forecast error vs forecast spread - 3-period partition (continued)
Figure 15: Forecast error vs forecast spread - 3-period partition (continued)
References


