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STOCK PRICE DISCOVERY IN EARNINGS SEASON

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ABSTRACT

This study investigates whether the timing of earnings announcement in earnings season affects stock price discovery process. This paper documents that market reaction is more favorable for earnings announcements made at the beginning of earnings season ("timing effect"). Price reaction on earnings announcement dates and post-announcement price drift are significantly stronger for positive earnings surprises released at the beginning of earnings season. Negative earnings surprises announced at the end of earnings season have the most pronounced post-announcement price decline. The timing effect associated with positive earnings surprises is consistent with industry information transfer theory. The timing effect associated with negative earnings surprise is mainly driven by market penalty on companies' strategic delay of bad news announcements.

JEL G12, G14, G30

KEYWORDS Market Reaction to Earnings News; Timing of Earnings Announcement

INTRODUCTION

To assist market efficiency, the Securities and Exchange Commission (SEC) requires publically traded companies to file their earnings reports in a timely manner. Form 10-K annual reports (Form 10-Q quarterly reports) are required to be filed within 90 (45) days after fiscal year (quarter) end. As a result, earnings announcements are clustered in time. The time window in which the majority of corporate earnings are released to the public is generally referred as "earnings season". Prior research on earnings announcements has shown that the timing of a company's earnings announcement is highly predictable (e.g. Begley and Fischer (1998)). Therefore, if a company reports earnings on time, its stock price discovery process should be independent of when the earnings announcement is made in earnings season. However, a company's stock price discovery process could be complicated by earnings season. On the one hand, clustered earnings announcements and massive media coverage of news events enrich investors' information set, which promotes stock price discovery. On the other hand, competing news events complicate investors' searching, processing, and interpretation of earnings news.

If investors are overwhelmed or distracted, stock price may not fully adjust to reflect the value implication of a company's released earnings news. Using a sample of quarterly earnings announcements released between years 1985 and 2003, this study documents that the timing of earnings announcement affects stock price discovery process (referred as "timing effect" thereafter). Specifically, market reaction is more favorable to earnings news announced at the beginning of earnings reason compared with those announced at the end of earnings season. Stocks with extreme positive earnings surprise released at the beginning of earnings season on average gain 3.1% over a three-day window around the earnings announcement date, and 2.9% in the following sixty trading days. The total price increase of 6% exceeds the total price increase of 2.8% for companies in the same earnings surprise quintile but report at the end of earnings season. Both immediate price reaction and post-announcement price drift (1.9% and 0.9% respectively) are significantly smaller for late-in-season earnings announcements. Stocks with extreme negative earnings surprise reported in early earnings season experience smaller price decline. The average total price drop is -3.5% (-2.4% on earnings announcement date and -1.1% in the following sixty trading days), significantly smaller than the

-6.5% (-2.4% on earnings announcement date and -4.1% in the following sixty trading days) for companies in the same earnings surprise quintile but report in late earnings season.

Two possible explanations are explored for the documented timing effect: industry earnings information transfer and a firm's strategic timing of news announcements. The timing effect associated with extreme positive earnings surprise can be attributed to industry earnings news transfer. The price reaction to late-in-earnings-season announcements is significantly positively correlated with the industry earnings news announced at the beginning of earnings season. The weaker price reaction associated with late-in-earnings-season announcements is mainly driven by the significant pre-announcement price increase.

The timing effect associated with extreme negative earnings surprise is mainly driven by market penalty on a firm's strategic delay of bad news announcement. An examination of the reporting pattern of individual firms shows the proportion of delayed bad news announcements increases from 7.6% in early earnings season to 37.2% toward the end of earnings season. The timing effect is the most pronounced among delayed bad news announcements: both price reaction (-3.0%) and price drift (-5.2%) are significantly more negative than those associated with on-time announcements. In addition, market appears to be forgiving if companies announce bad news earlier than expected, regardless whether the bad news is reported in early or late earnings season. Stocks of these earlier-than-expected bad news announcers do not significantly underperform their size-B/M matching portfolios in the following sixty trading days.

This study contributes to the literature with evidence of how news is incorporated into stock price in a unique setting, i.e. during a predictable time period with intensive mandatory earnings news announcements. It is yet unclear in the literature how a cluster of mandatory news announcements of the same nature affects stock price discovery process, and whether the value implication of the timing of such mandatory news announcement reflects management's opportunism or desire for better information dissemination. Evidence in this study suggests that stock price discovery process is distorted in earnings season. The reward to good news announced at the beginning of earnings season and penalty on bad news released toward the end of earnings season suggest that the timing of earnings announcement has a strong effect on investors' perceptions of the information content of reported earnings. As a result, price discovery is complicated by investors' rational information extrapolation, which promotes market efficiency, and investors' representativeness bias, which distorts price discovery. The remainder of the paper proceeds as follows. Section II reviews related literature on market response to corporate news announcements. Section III describes data measurements. Section IV discusses empirical results. The last section concludes.

RELEVANT LITERATURE REVIEW

If the timing of a company's mandatory earnings announcement is highly predictable, stock price reaction to such a mandatory news announcement should be solely determined by the value implication of earnings news. However, existing studies suggest that the timing of earnings announcement in earnings season may have value implication as well. The value implication of the timing of earnings announcement in earnings season is associated with the facts that corporate earnings announcements are clustered in time, and a firm's earnings news conveys information not only relevant to the announcing firm but also to firms that are linked to the announcing firm in product markets. Studies (e.g. Foster (1981), Han and Wild (1990), Hou (2007)) have shown that a firm's stock price reacts to earnings announcements made by competing firms in the same industry. The intra-industry diffusion of earnings news is due to the fact that firms in the same industry have similar cash flow characteristics. Firms in the same industry react similarly to changes in regulations and shift in supply and demand conditions. Firms' growth opportunities and the corresponding investment and financing decisions are closely related to the industry's expansion and contraction. These commonalities suggest that a firm's earnings news not only discloses the impact of firm specific policies, but also reflects the impact of common industry externalities.

The clustered earnings announcements in earnings season create a unique environment for information transfer. In earnings season, intensive corporate earnings news release not only broadens information transfer channels, but also stimulates market participants' information searching activities. At the beginning of earnings season, investors hold their trading interests to wait for massive earnings news to set a distinctive tone to trade in equity market. As more earnings news hits the market, investors' trading interests grow, which promotes information dissemination. As investors and financial analysts extrapolate from the released earnings news to update their understanding of macroeconomic condition, and industry challenges and opportunities, there is smaller likelihood that a yet-to-be-released earnings announcement surprises the market. Alternatively, the value implication of the timing of earnings announcement in earning season could be attributed to companies' strategic timing of corporate news announcements to optimize post-announcement stock price. Prior research suggests that companies strategically choose the time of the day (e.g. Patell and Wolfson (1982), Francis, Pagach and Stephan (1992), Gennotte and Trueman (1996)), the day of the week (e.g. Damodaran (1989), DellaVigna and Pollet (2009)) to make news announcements, or intentionally accelerate or delay news announcements (e.g. Givoly and Palmon (1982), Chambers and Penman (1984), Kross and Schroeder (1984), Begley and Fischer (1998), Bagnoli et al. (2002), Graham et al. (2005), Cohen et al. (2007a), Kothari et al. (2009))). Evidence of strategic timing of news disclosure includes negative earnings news is announced more often outside trading hours than during trading hours, companies tend to release bad news on Friday, and earnings report delay is negatively related to the magnitude of earnings news ("good news early, bad news late").

The strategic timing of corporate news announcements has been argued as driven by managers' opportunism. For example, managers opportunistically hold bad news until the market closes or when investors are more distracted from job-related tasks, e.g. on Friday, to minimize negative market reaction. Managers may accelerate the release of good news before competing firms weigh in with strong earnings performance that will downplay the value implication of the announcing firm's good news. Managers may also accelerate bad news announcement to minimize litigation and reputation risks, or when they sense a "best possible light" to position the bad news. The strategic timing of news announcements could also be driven by managers' attempt to be informative. For example, managers may accelerate the release of good or bad news to provide investors with more timely information, or build the company's reputation for transparent reporting. Similarly, the withholding of bad news may reflect managers' desire for better information dissemination, giving less informed investors more time to digest complex information so that they can fairly assess the value implication of bad news.

In summary, information transfer literature implies that stock price discovery process will be distorted in earnings season, with weaker market reaction to earnings news as earnings season extends. Strategic timing literature suggests that market reaction to late-in-earnings-season announcements is less favorable. Because investors' ex ante expectation is that it takes managers more time to prepare for the release of manipulated or bad earnings news. Managers may use the extra time to resolve disagreement with auditors, prepare for analysts' questioning, come out a recovery plan to offset the negative impact of bad earnings news, or explore complicated accounting practices that may reverse the poor earnings result.

DATA AND METHODOLOGY

The value implication of the timing of earnings announcement in earnings season is studied with a sample of quarterly earnings announcements made between calendar years 1985 and 2003 by companies listed on the New York Stock Exchange (NYSE), American Stock Exchange (AMEX), and National Association of Securities Dealers Automated Quotation system (NASDAQ). Real estate investment trusts (REITs), American Depository Receipts (ADRs), and closed-end mutual funds are excluded. To mitigate microstructure effect, stocks priced below $5 per share are also excluded. To ensure that fiscal quarters are aligned, the research sample is restricted to firms with March, June, September and December fiscal quarter end. Earnings season is identified with Compustat quarterly earnings announcement date. Following Della

Vigna and Pollet (2009), when the earnings announcement date is also available at the Institutional Brokerage Estimate System (I/B/E/S) but different from the Compustat date, the earlier date is used. While the accuracy of earnings announcement date is likely to be higher when it is available at both I/B/E/S and Compustat, Compustat earnings announcement dates are used for the identification of earnings season because I/B/E/S shows a coverage bias toward large firms. The magnitude of earnings news is estimated by earnings surprise, defined as the difference between the I/B/E/S reported Earnings-Per-Share (EPS) and I/B/E/S consensus EPS forecast, which is the median of the most recent forecasts from individual analysts. The difference is normalized by the stock price at the end of the reported fiscal quarter. To exclude stale forecasts, only the one- or two-quarter ahead forecasts issued or reviewed in the last sixty calendar days before the earnings announcement are included. If an analyst issues multiple forecasts during that period, the most recent forecast by this individual will be used.

Both the reported EPS and the consensus EPS forecast are from the I/B/E/S raw unadjusted earnings dataset. The commonly used standard I/B/E/S dataset reports the actual and forecasted EPS that are adjusted for stock splits (i.e. EPS is based on the number of shares outstanding as of today along with a split factor). The estimated and actual EPS are rounded to the nearest cent after making retroactive and cumulative stock split adjustments, which makes the comparison of actual and estimated EPS problematic for firms that have executed stock splits. A disproportionate number of firms will be categorized as "exactly met " analysts' expectations if standard I/B/E/S data file were used, while in fact they have missed or exceeded analysts' expectations. Stock price discovery process is evaluated by market reaction to earnings surprise, which includes immediate price reaction (CAR) to earnings surprise and post-announcement price drift (PEAD) over a window of sixty trading days. To calculate CAR and PEAD, each stock is matched with 1 of 25 size and book-to-market (B/M) portfolios at the end of June. Size is measured by the market capitalization at the end of June. B/M is measured as the ratio of the book value of equity at the end of last fiscal year-end in the prior calendar year to the market value of the equity at the end of December of the prior calendar year. Daily individual stock returns and company financials are from CRSP/Compustat merged database. Daily returns of the 25 size-B/M portfolios are from professor Kenneth French's web site. CAR is calculated as the sum of daily abnormal returns over a window of three trading days [-1, 1], where *t=0* is the earnings announcement date. PEAD is calculated as the difference between the buy-and-hold return of the announcing firm and that of a size-B/M matching portfolio over a window of [2, 61] in trading days,

$$CAR_{i,q} = \sum_{t=-1}^{t=1} \left(r_{i,t} - r_{p,t}\right) \quad (1)$$

$$PEAD_{i,q} = \prod_{t=2}^{t=61}(1 + r_{i,t}) - \prod_{t=2}^{t=61}(1 + r_{p,t}) \quad (2)$$

where $r_{i,t}$ is the daily return of stock *i* and $r_{p,t}$ is the daily return of its matching size-B/M portfolio on trading day *t*, where *t=0* is the announcement date of quarter *q*'s earnings. PEAD is calculated over sixty trading days because Bernard and Thomas (1989) report that most of the drift occurs during the first sixty trading days after the announcement. For a small number of earnings announcements, next quarter's earnings announcement date falls into the sixty-trading-days window. In this case, a stock's raw return is replaced by its matching size-B/M portfolio's raw return for the remaining sixty-days window for PEAD calculation.

RESULTS AND DISCUSSIONS

The Timing of Quarterly Earnings Announcements

The SEC explicitly requires that quarterly earnings reports to be filed within 45 days after the fiscal quarter ends. As a result, quarterly earnings announcements are clustered in time, forming the so called "earnings season"- the time period in which a majority of corporate earnings are released to public. Table 1 reports

the descriptive statistics of the number of calendar days between earnings announcement date and fiscal quarter end based on a sample of 420,900 observations of Compustat quarterly earnings announcements made between calendar years of 1985 and 2003. As shown in Table 1, in accordance with SEC's requirements, 90% of companies release their quarterly earnings results within 45 calendar days after fiscal quarter end. Earnings season kicks off approximately two weeks after fiscal quarter end, as only 5% of earnings announcements are released in the first 15 calendar days after fiscal quarter end. Earnings season lasts about one month with 90% of earnings announcements are made during earnings season, which starts from the 16th calendar day after fiscal quarter end.

Table 1: Descriptive Statistics of The Number of Calendar Days Between Earnings Announcement Date and Fiscal Quarter End

Fiscal Quarter	Mean	Median	P5	P10	P20	P25	P30	P40	P60	P70	P75	P80	P90	P95
1	32	29	15	18	21	23	24	27	33	37	38	40	45	47
2	32	30	16	18	21	23	24	27	33	37	38	40	45	47
3	31	29	15	17	21	22	24	26	33	36	38	40	45	46
Average	32	29	15	18	21	23	24	27	33	36	38	40	45	47

This table shows the percentile descriptive statistics of the number of calendar days between earnings announcement date and fiscal quarter end.

Based on the percentile descriptive statistics in Table 1 and frequency distribution of reporting lag (untabulated for brevity), the timing of a quarterly earnings announcement is defined as "at the beginning of earnings season (BGN)" if it is made between the 16th and 25th calendar days after fiscal quarter end, " in the middle of earnings season (MID)" if it is made between the 26th and 35th calendar days after fiscal quarter end, or "at the end of earnings season (END)" if it is made between the 36th and 47th calendar days after fiscal quarter end. If an earning announcement is made in the first 15 calendar days or beyond calendar day 47 after fiscal quarter end, its timing is labeled as "Before" and "After" respectively. Table 2 presents the percentage of total quarterly earnings announcements, and percentage of announcements of extreme earnings news released before, during, and after earnings season, based on a sample of Compustat quarterly earnings released between calendar years 1995 and 2003. Extreme earnings news is identified based on quarterly sorting of reported earnings surprises into quintiles. Extreme good and bad earnings news refers to earnings surprises ranked in the top and bottom quintiles.

Table 2: Percentage of Earnings Announcements Made Before, During, and after Earnings Season

Timing of Announcement	Time Interval	% of Total Announcements (N=420,900)	% of Extreme Bad Earnings News (N=35,846)	% of Extreme Good Earnings News (N=36,137)
Before	D1-D15	5%	5%	7%
BGN	D16-D25	30%	28%	40%
MID	D26-D35	30%	27%	28%
END	D36-D47	30%	21%	15%
After	D48 ~	5%	19%	10%

This table reports the percentage of quarterly earnings announcements made before, during, and after earnings. An announcement is labeled as "Before" if it is released between calendar days 1 and 15 after fiscal quarter end (D1-D15), "BGN" if between calendar days 16 and 25 (D16-D25), "MID" if between calendar days 26 and 35 (D26-D35), "END" if between calendar days 36 and 47 (D36-D47), and "after" if beyond calendar day 47 after fiscal quarter end. The last two columns report the distribution of extreme news announcements. Extreme news is defined based on quarterly quintile ranking of earnings surprises for companies with fiscal quarter ends in March, June, September, or December.

As shown in the column "% of Total Announcements", with the above partition of earnings season, 90% of quarterly earnings announcements are announced during earnings season, 5% are announced before

earning season starts, and 5% are announced after earnings season ends. Quarterly earnings announcements spread evenly in earnings season with approximately 30% announcements made at the beginning, in the middle, and at the end of earnings season. Such a partition of earning season is not affected by whether company has a calendar fiscal quarter end, or a December fiscal quarter end (results untabulated for brevity). The timing of extreme earnings news announcement is determined by whether the news is good or bad. Extreme good earnings news tends to be released in early earnings season: 40% of extreme good news is released over the first 10 calendar days of earnings season while only 15% is announced at the end of earnings season. While extreme bad news announcements spread evenly in earnings season, 19% of extreme bad news is reported after earnings season ends.

The Timing Effect

To examine the effect of the timing of earnings announcement on stock price discovery process, a univariate analysis is performed. At the end of each calendar quarter, earnings surprises announced in the quarter are sorted into quintiles. For each earnings surprise quintile, the mean price reaction and post-announcement price drift, as defined in equation (1) and (2), are calculated based on the timing of news announcement. For extreme good and bad earnings news, the mean price reaction and price drift, and the difference in price reaction and price drift between announcements made at the beginning versus at end of earnings season are reported in Table 3.

Table 3: Price Reaction and Post-Announcement Price Drift For Extreme Earnings Surprise Quintiles by Timing of Earnings Announcement

Panel A: Extreme Negative Earnings Surprise Quintile						
	Earnings Season					
	Before	BGN	MID	END	After	BGN-END
Price Reaction	-1.76*	-2.44**	-2.51**	-2.38**	-1.87**	-0.06
Price Drift	0.39	-1.11	-2.58*	-4.12**	-2.18*	3.01**
Panel B: Extreme Positive Earnings Surprise Quintile						
	Earnings Season					
	Before	BGN	MID	END	After	BGN-END
Price Reaction	3.99*	3.10**	2.66**	1.93**	1.91*	1.17*
Price Drift	4.13*	2.90**	1.65*	0.94	1.34	1.96*

*This table reports the mean price reaction and price drift of extreme earnings surprise quintiles grouped by the timing of earnings announcements. The timing of earnings announcement is labeled "BGN" ("END") if the announcement is made in the first (last) 10 calendar days of earnings season, and "MID" otherwise. An earnings announcement made in the first 15 calendar days (beyond calendar day 47) after fiscal quarter end is labeled "Before" ("After"). Extreme earnings surprises are those ranked top and bottom 20 percent in quarterly quintile sorting. The p-values are calculated using standard errors adjusted for heteroskedasticity and clustering by date. Notation *, **, and *** indicate significance at the 10%, 5%, and 1% levels respectively.*

Evidence in Table 3 shows market reaction is more favorable for earnings announcements made at the beginning of earnings season ("timing effect"). For extreme positive earnings surprise quintile, price reaction and post-announcement drift are stronger if the announcement is made in early earnings season. Stock price of announcements made at the beginning of earnings season on average gain 3.1% over three trading days around earnings announcement date and 2.9% over the post-announcement sixty trading days. In comparison, stocks with announcements at the end of earnings season appreciate significantly less with price reaction of 1.93% and price drift of 0.94%. As earnings season proceeds, the magnitude of price reaction and post-announcement price drift monotonically decreases.

For extreme negative earnings surprise quintile, market reaction is more negative if the bad news is released in late earnings season. Stocks with bad news released at the end of earing season on average are

accompanied with a total price decline of 6.5% with 2.38% on earnings announcement date and 4.12% over the following sixty trading days. The total price decline is only 3.5% for bad news announced at the beginning of earnings season and primary on earnings announcement date. Early announcers of bad news do not significantly underperform their size-B/M matching portfolios. Their post-announcement price drift of -1.11% is not significantly difference from zero. The last column of Table 3 reports the difference in price reaction and post-announcement price drift between at-the-beginning and at-the-end announcements. To verify that these differences are statistically meaningful, the following regression is conducted using announcements in the top and bottom of the earnings surprise quintiles and released at the beginning and end of earnings season,

$$CAR\ (or\ PEAD) = \propto_0 + \propto_1 FE_5 + \propto_2 BGN + \propto_3 (FE_5)(BGN) + \varepsilon \quad (3)$$

where CAR is the immediate price reaction and PEAD is the post-announcement price drift as defined in equations (1) and (2), FE_5 is an indicator variable that is equal to one for the top quintile of earnings surprises and zero for the bottom quintile of earnings surprises, BGN is a dummy variable that is equal to one if an announcement is made at the beginning of earnings season, or zero if an announcement is made at the end of earnings season. The regression coefficient α_2 tests whether the timing effect associated with extreme negative earnings surprise announcement is statistically significant, while the sum of regression coefficients of α_2 and α_3 tests whether the timing effect associated with extreme positive earnings surprise announcement is statistically significant. The timing effect associated with extreme positive earnings surprise is statistically significant. The price reaction difference between the announcements made at the beginning and at the end of earnings season is 1.17%. The post-announcement price drift difference is 1.96%. Both differences are statistically different from zero at the 10% level using standard errors adjusted for heteroskedasticity and clustering by date.

The timing effect associated with extreme negative earnings surprise is mainly due to the significant difference in the post-announcement price decline. The 3% difference in price drift is statistically significant at the 5% level, while the price reaction difference is not statistically significant. Prior research shows that market reaction to earnings news varies with firm size, B/M ratio, number of analyst following, institutional ownership, earnings persistence, and day of the week (e.g. Chamber and Penman (1984), Bernard and Thomas (1989), DellaVigna and Pollet (2009)). Thus, to control for other determinants of market reaction to earnings news, a multivariate test is performed. Price reaction (CAR) or post-announcement price drift (PEAD) is regressed on the earnings surprise quintile rank (FE), the timing of earnings announcement (Timing), the interaction term (FE x Timing), control variables (X_i), and the interaction of control variables with FE:

$$CAR\ or\ PEAD = \propto_0 + \propto_1 FE + \propto_2 Timing + \propto_3 (FE \times Timing) + \sum_{i=1}^{n} c_i X_i + \sum_{i=1}^{n} \beta_i (FE \times X_i) + \varepsilon \quad (4)$$

where control variables include size and B/M decile ranks, the log of the number of analysts following during the most recent fiscal year, institutional ownership, earnings persistence, earnings volatility, day of week/month/year dummies, and industry dummies using Fama-French twelve industry classification. The size and B/M deciles are formed at the end of June of each year. Size is based on the market value of equity at the end of June. B/M ratio is calculated as the book value of equity for the last fiscal year-end in the previous calendar year dividend by the market value of equity for December of the previous calendar year. Earnings persistence is the first-order autocorrelation coefficient of quarterly EPS during the past 4 years (minimum four observations required). Earnings volatility is the standard deviation during the preceding 4 years of the deviations of quarterly earnings from a-year-ago earnings (minimum four observations required). Institutional ownership is measured as the percentage of shares owned by institutions at the end of the most recent calendar quarter. The indictor of timing takes value of 1 if earnings news is reported at

the beginning of earnings season, 2 if earnings news is reported in the middle of earnings season, or 3 if earnings news is reported at the end of earnings season.

Following existing literature, the quintile rank of earnings surprises is used in regression, as apposed to the magnitude of earnings surprises, to reduce the impact of outliers. This is because the relationship between announcement-day abnormal returns and earnings surprise is highly nonlinear with small negative surprises having big effects (Kothari (2001)). Regression results with all observations (Regressions (1) and (2)) and with observations in the top and bottom earnings surprise quintiles (Regressions (3) and (4)) are reported in Table 4.

Table 4: Regression Analyses of Price Reaction to Earnings Surprise and Post-Announcement Price Drift

Independent Variables	**(1) Price Reaction (Car)**	**(2) Price Drift (Pead)**	**(3) Price Reaction (Car)**	**(4) Price Drift (Pead)**
FE	0.0193***	0.0098***		
FE interacted with Timing	-0.0012**	-0.0013*		
$FE_{extreme}$			0.0196***	0.0121***
$FE_{extreme}$ interacted with Timing			-0.0010**	-0.0024*
Control variables interacted with FE	Yes	Yes	Yes	Yes
Constant	-0.032***	-0.016***	-0.031***	-0.012***
Observations	105,833	105,833	38,790	38,790
R^2	7.6%	0.2%	11.4%	4%

*This table reports the multivariate test of the effect of announcement timing on price reaction to earnings surprise and post-announcement price drift. FE is the earnings surprise quintile (FE=1: lowest, 5: highest) based on quarterly ranking of earnings surprises. $FE_{extreme}$ is an indicator variable for the top and bottom earnings quintiles. Timing is an indicator variable that equals to 1/2/3 if an announcement is made at the beginning/middle/end of the earnings season. Control variables include size and B/M deciles, log of the number of analyst following, institutional ownership, earnings persistence, earnings volatility, indicator variables for year, month, day of week, and Fama-French 12 industry classification. Notation *, **, and *** indicate significance at the 10%, 5%, and 1% levels respectively, and based on standard errors adjusted for heteroskedasticity and clustering by the day of announcement.*

The regression coefficient on the timing interaction term is significantly negative at 5% level for price reaction and at 10% level for price drift (Regressions (1) and (2)). The coefficient estimates on FE and the interaction term (FE x Timing) imply that market reaction to earnings surprise is significantly stronger if an announcement is made at the beginning of earnings season. For example, the price reaction to earnings announcements released at the beginning of earnings season is more sensitive: the sensitivity is 0.0181 (= 0.0193-0.0012 x 1) for announcements at the beginning of earning season, 15.3% larger than the sensitivity of 0.0157 (=0.0193-0.0012 x 3) associated with announcements made at the end of earnings season. Similarly, price drift is 44% more sensitive if news is released at the beginning of earnings season. Results are similar when extreme earnings surprise quintiles are used (Regressions (3) and (4)).

Industry Information Transfer and the Timing Effect

To examine whether the documented timing effect is due to industry earnings information transfer, I examine whether the stock price of companies that report earnings at the end of earnings season is positively correlated with earnings surprises announced at the beginning of earnings season by companies in the same industry. At the end of June of each calendar year, companies in research sample are assigned to one of twelve Fama-French industries based on the firm's four-digit SIC code. For each Fama-French industry, the average earnings surprise of the companies in the same earnings surprise quintile that have released earnings results at the beginning of earnings season is calculated.

A Fama-McBeth regression is conducted as follows. At the end of each calendar quarter, for each company that reports at the end of earnings season, its immediate price reaction (CAR) is regressed on its own earnings surprise (FE), the reported industry average earnings surprise ($FE_{industry}$), size, book-to-market ratio (B/M), and the its abnormal buy-and-hold returns over the first and second ten-day windows of earnings season (R_{BGN}; R_{MID}). The time series averages of the cross-sectional regression coefficients for companies in extreme earnings surprise quintiles are presented in Table 5.

Table 5: Industry Information Transfer and Timing Effect

	Log (Size)	**Log (B/M)**	**FE**	**$Fe_{industry}$**	**R_{BGN}**	**R_{MID}**
Regression (1)	−0.0006***	0.0004***	0.0075***			
Regression (2)	−0.0006***	0.0004***		0.1252***		
Regression (3)	−0.0006***	0.0004***	0.0064***	0.1246**		
Regression (4)	−0.0006***	0.0004***	0.0059**	0.0014*	0.1481**	0.0621**

*This table reports the time-series average of the regression coefficients of regressing price reaction (CAR) to extreme earnings surprises announced at the end of earnings season on industry earnings news and the announcing firm's earnings surprise and its pre-announcement price change. Size and B/M ratio are defined in data section. FE is the earnings surprise of announcing firm. $FE_{industry}$ is the average earnings surprise of the companies in the same earnings surprise quintile that have reported at the beginning of earnings season. R_{BGN} (R_{MID}) is the announcing firm's abnormal buy-and-hold return (BHR) over the first (second) ten days of the earnings season, measured as the difference between the BHR of the announcing firm and its size-/B/M matching portfolio. Notation *, **, and *** indicate significance at the 10%, 5%, and 1% levels respectively.*

Regression results in Table 5 confirm that earnings news disseminates within industry. The price reaction to late-in-earnings-season announcements is positively correlated with the industry earnings news announced at the beginning of earnings season (Regression (2)). The regression coefficient of $FE_{industry}$ remains positive at 5% significance level after the earnings surprise of announcing firm is added (Regression (3)), suggesting that earnings news announced at the beginning of earnings season contains information about the future prospects of companies in the same industry, and thus affects the stock price discovery of companies announcing at the end of earnings season. In the last regression in Table 5, the abnormal cumulative returns of the announcing firm over the first and second ten days of earning season are added for additional evidence on industry earnings news transfer. The regression coefficients on both cumulative returns are significant at 5% level, while the regression coefficient of announced industry earnings news is much smaller and significant at 10% level. This result suggests that market participants extrapolate from the announced earnings news for common industry opportunities and constraints that apply to all firms within. The pre-announcement price adjustment to industry earnings news helps explains the weaker price reaction and post-announcement price drift associated with extreme positive earnings surprise.

Information transfer theory, however, does not well explain the timing effect associated with extreme negative earnings surprise. An examination of the pre-announcement price change associated with extreme negative earnings surprise announced in the last ten days of earnings season shows a significant price decline of 1.3% over the first twenty days of earnings season (results untabulated for brevity). Despite the significant pre-announcement price drop, market reaction (measured by both price reaction and post-announcement price decline) is still significantly more negative for bad news announced at the end of earnings season. Therefore, an alternative explanation is explored in the following section.

Strategic Timing of News Announcements and the Timing Effect

Corporate managers have become increasingly aware of the potential impact that corporate disclosure strategies can have on a firm's value. An important component of a firm's disclosure strategy is the timing of its public news announcements. Existing studies show that companies strategically choose the time of the day and the day of the week to make news announcements for optimization of post-announcement stock price. Companies also intentionally accelerate or delay news announcements, depending on the nature of

the news and market condition. If investors interpret the timing of earnings announcement in earnings season as indication of the likelihood of manipulated earnings news or bad earnings news, and delayed announcements dominate the announcements released at the end of the earnings season, the timing effect associated with extreme negative earnings surprise can be attributed to market penalty on delayed bad news announcements. Table 6 provides descriptive statistics on the stability of companies' reporting pattern. The mean of reporting lag (RepLag), which is the number of calendar days between fiscal quarter end and quarterly earnings announcement date, indicates that firms spend approximately one month on earnings announcement preparation. The standard deviation of RepLag (std Replag) is relatively small because weekends and holidays naturally introduce some variations.

Table 6: Reporting Pattern of Individual Firms

Fiscal Quarter	Mean Replag	Std Replag	Mean Del	Median Del	Mean |Del|	Median |Del|
1	32	4.9	-0.23	0.00	4.41***	3.56***
2	32	4.7	-0.24	-0.06	4.14***	3.44***
3	31	4.9	-0.19	0.00	4.33***	3.46***

This table reports evidence on the reporting pattern of individual firms. Reporting lag ("RepLag") is defined as the number of calendar days between fiscal quarter end and earnings announcement date. Announcement Delay ("DEL") is the difference between the RepLag of the reported quarter and the same fiscal quarter of previous year. The mean and median of the absolute value of DEL are reported in the last two columns.

The stability of a firm's reporting pattern is indicated by announcement delay (DEL). Following earlier studies (e.g. Begley and Fischer (1998)), *DEL* for firm *i* in quarter *q* and year *t* is defined as $DEL_{i,q,t} = RepLag_{i,q,t} - RepLag_{i,q,t-1}$. A negative (positive) *DEL* indicates that an earnings report is accelerated (delayed) compared with the same quarter last year. As shown in Table 6, the mean and median DEL are negative but not significantly different from zero, representing little time trend of reporting lags for individual firms. The mean and median of the absolute value of announcement delay (|DEL|) are of similar magnitude as the standard deviations of reporting lag. This shows that the within-firm time variation of reporting lag is due to the swing associated with reports of the same fiscal quarter in consecutive years. Table 7 presents the average announcement delay (DEL) of extreme earnings surprise announcements. Based on the reporting pattern of individual firms, an earnings announcement is classified as "on time" if its *DEL* satisfies $-5 \leq DEL \leq 5$, or as "advanced (delayed)" if its *DEL* is less (more) than 5. Robustness check shows similar results when classification is based on variation of 3 or 4 days.

There is little evidence that firms intentionally accelerate release of good news. Among 26, 125 announcements of extreme positive earnings surprise, the proportion of advanced announcements is similar to that of delayed announcements (16.7% vs. 13.2%). In addition, advanced good news announcements spread evenly over earnings season. For example, 15.6% of announcements released at the beginning of earnings season are accelerated, compared with the 14.4% for announcements released at the end of earnings season. In comparison, there is clear evidence that firms intend to delay bad news announcements. Out of 26,244 announcements of extreme negative earnings surprise, 25.3% are delayed announcements while 12.3% are accelerated. There is clear evidence that the proportion of delayed announcements increases as earnings season proceeds. At the beginning of earnings season, only 5.6% of extreme positive earning surprise announcements are delayed. The proportion increases to 23.8% at the end of earnings season. Such increase is more dramatic for extreme negative earnings surprise announcements. There are 37.2% announcements made at the end of earnings season that are delayed, compared with only 7.6% at the beginning of earnings season.

Table 7: Earnings Announcement Delay (DEL) and the Timing of Earnings Announcement

Panel A	**Extreme Positive Earnings Surprise Quintile**					
Timing	**Before**	**BGN**	**MID**	**END**	**After**	**Nobs**
Advanced	-12.8*	-12.2*	-13.7*	-15.1*	-16.8*	4,350
	[25.1]	[15.6]	[17.0]	[14.4]	[16.6]	[16.7]
On Time	-1.1	-0.6	0.0	-0.10	0.1	18,323
	[73.3]	[78.8]	[67.0]	[61.8]	[49.5]	[70.1]
Delayed	7.1*	7.4*	8.5*	12.1*	23.6*	3,452
	[1.6]	[5.6]	[16.0]	[23.8]	[33.9]	[13.2]
Number of observations	1,928	10,996	7,250	3,528	2,367	26,125
	[100]	[100]	[100]	[100]	[100]	[100]
Panel B	**Extreme Negative Earnings Surprise Quintile**					
Timing	**Before**	**BGN**	**MID**	**END**	**After**	**Nobs**
Advanced	-14.1*	-13.2*	-13.8*	-16.7*	-17.5*	3,235
	[20.5]	[12.9]	[12.8]	[9.5]	[11.1]	[12.3]
On Time	-0.8	-0.4	0.4	0.3	0.7	16,383
	[77.3]	[79.5]	[63.7]	[53.3]	[36.8]	[62.4]
Delayed	7.4*	7.2*	8.9*	12.7*	25.1*	6,626
	[2.2]	[7.6]	[23.5]	[37.2]	[52.1]	[25.3]
Number of observations	1,429	7,879	7,183	5,189	4,564	26,244
	[100]	[100]	[100]	[100]	[100]	[100]

This table reports the announcement delay (DEL) for announcements of extreme earnings surprises. DEL is defined as $RepLag_{i,q,t} - RepLag_{i,q,t-1}$, *where RepLag is the number of calendar days between the fiscal quarter end and the earnings announcement date. An earnings announcement is "On Time" if its DEL satisfies -5≤DEL≤5. An announcement is "Delayed" ("Advanced") if its DEL>5 (DEL<-5). Timing indicators (e.g. "BGN") are defined in Table 3. Numbers in bold brackets are the percentage proportions of observations relative to the column sum of number of observations (Nobs). Notation * indicate significance at 5% level.*

Table 8 summarizes the comparison of price reaction and price drift associated with extreme negative earnings surprises announced at the beginning versus at the end of earnings season.

Table 8: Timing Effect and Strategic Delay of Extreme Negative Earnings Surprise Announcements

	Price Reaction (CAR)				**Price Drift (PEAD)**			
	BGN	**MID**	**END**	**BGN-END**	**BGN**	**MID**	**END**	**BGN-END**
Advanced	-2.05*	-1.83*	-1.60*	-0.45	-1.4	-0.96	-1.27	-0.13
Delayed	-2.41***	-2.67***	-2.96***	0.55*	0.73	-1.88**	-5.16***	5.89***
On Time	-2.46***	-2.47***	-2.18***	-0.28	-0.51	-1.73**	-2.43***	1.92**

This table shows the price reaction and price drift associated with extreme negative earnings surprises announcements categorized by the timing of announcement in earnings season ("BGN", "MID", "END") and whether the announcement is accelerated, on-time, or delayed ("Advanced", "On Time", "Delayed").

Consistent with the argument of strategic timing theories, the timing effect is the most pronounced for delayed bad news announcements. In particular, the more negative price reaction associated with announcements released at the end of earnings season only applies to delayed bad news announcements. Delayed bad news announcements are also accompanied by the largest negative post-announcement price drift (-5.16%), which quadruples the price decline associated with advanced announcements (-1.27%), and doubles that for on-time announcements (-2.43%). It is worth of notice that timing effect does not apply to bad news announced earlier than expected. The immediate price reaction is independent of the timing of advanced announcements. Moreover, there is no significant underperformance relative to the size-B/M matching portfolio in the following sixty trading days. This result suggests that investors perceive the acceleration of bad news announcement as driven by manager's attempt to be informative. In Table 9, a regression framework is used to separate the impact of strategic delay of bad news announcement from the

impact of the timing of news announcement. Results are based regressions of price reaction and price drift associated with extreme negative earning surprise announcements made at the beginning and at the end of earnings season.

Table 9: Regression Analyses of the Impact of Strategic Timing on Price Discovery Associated With Extreme Negative Earnings Surprises

Independent Variables	**(1) Price Reaction (Car)**	**(2) Price Reaction (Car)**	**(3) Price Drift (Pead)**	**(4) Price Drift (Pead)**
FE	0.0068***	0.0068***	0.0051***	0.0051***
FE interacted with END	0.0031*	0.0031	0.0047**	0.0040**
FE interacted with Strategic		0.0002*		0.0013**
Control variables interacted with FE	Yes	Yes	Yes	Yes
Constant	-0.051***	-0.051***	-0.067***	-0.067***
Observations	26, 244	26, 244	26,244	26,244
R^2	7.5%	7.5%	6.2%	6.2%

*This table reports the regression analyses of the impacts of announcement timing and strategic delay of bad news release on stock price discovery process. FE, as defined in data section, is worsorized at the 1% and 99% levels. "END" is a dummy variable that equals to 1/0 if an announcement is made at the end/ beginning of earnings season. "Strategic" is a dummy variable that equals to 1 if an announcement is delayed and 0 otherwise. Control variables include size and B/M deciles, log of the number of analyst following, institutional ownership, earnings persistence, earnings volatility, indicator variables for year, month, day of week, and Fama-French 12 industry classification. Notation *, **, and *** indicate significance at the 10%, 5%, and 1% levels respectively, and based on standard errors adjusted for heteroskedasticity and clustering by the day of announcement.*

Table 9 results reinforce that strategic delay of bad news announcements significantly contributes to the documented timing effect associating with extreme negative earnings surprise announcements, especially to the significant post-announcement price decline. The post-announcement price decline is stronger if an announcement is made at the end of earnings season (Regression (3)). The regression coefficient of 0.0047 is statistically significant at 5% level. The post-announcement price decline is the most pronounced if an end-of-earnings-season announcement is a delayed announcement (Regression (4)). Similar results apply to immediate price reaction (Regressions (1) and (2)). The stronger negative price reaction is mainly due to penalty on delayed bad news announcement. After the indicator of strategic timing is added to regression, the explanatory power of timing indicator becomes insignificant.

CONCLUSION

This study investigates stock price discovery in a unique setting: over a time window called "earnings season" when a cluster of corporate earnings announcements arrive the market and complicate the information environment. Results in this study show that the immediate price reaction to earnings news and post-announcement price drift are more favorable if a news announcement is released at the beginning of earnings season rather than at the end of earnings season ("timing effect"). Specifically, price increase on and after earnings announcement date is significantly more positive for good news announced at the beginning of earnings season, while bad news announced at the end of earnings season is associated with the most pronounced price decline on and after earnings announcement date. Two possible explanations for the documented timing effect are explored: industry earnings news transfer and market penalty on delayed news announcement. The timing effect associated with positive earnings news announcement is consistent with information transfer theory. Investors extrapolate from the earnings announcements by competing firms in the same industry to update their expectations of to-be-released earnings news. The stock price of companies report at the end of earnings season shows a pre-announcement price increase, which explains the weaker price reaction to earning announcement and post-announcement price drift. The timing effect associated with negative earnings news announcements is mainly due to market penalty on delayed bad

news announcement. Announcements made at the end of earnings season that are delayed have the most significant price decline on and after earnings announcement date. In comparison, earlier-than-expected bad news announcements have the smallest price decline on earnings announcement date. In addition, these accelerated announcers do not significantly underperform their size-B/M matching portfolios.

After controlling for the impact of strategic delay of bad news announcement, the timing effect still exists. End-of-earnings-season on-time announcements are accompanied by significant post-announcement price decline, while stocks of on-time announcements made at the beginning of earnings season do not significantly underperform. This suggests that investors' cognitive bias plays an important role in stock price discovery process. It appears that investors perceive the timing of announcement as an indicator of manipulated or poor quality earnings news, although the timing of announcement may be merely due to a firm's accounting complexity that calls for more time to prepare and release earnings results.

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BANK STOCK AND OPTION TRANSMISSIONS IN FINANCIAL CRISIS

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ABTRACT

We investigate bank stock and option transmissions during the financial crisis in 2008. Contemporaneous and lagged-one stock order imbalances have a significant impact on option returns. A time-varying GARCH model is employed to confirm the results. We develop an imbalance-based call (put) trading strategy that buys the call (put) if the previous day's stock imbalance is positive, and sells the stock if the previous day's stock imbalance is negative. The empirical results do not show a positive premium, which implies market efficiency between option and stock markets in financial crisis.

JEL: G01, G14, G21

KEYWORDS: Order Imbalance, Market Efficiency, Investment Bank, Commercial Bank, Financial Crisis, Option

INTRODUCTION

In 2008, the subprime crisis in U.S. spilled over and became the catalyst for a much broader global financial crisis. Intervention by the Treasury and the Federal Reserve in the financial markets was intend to avoid broader spillovers to other markets and sectors of the economy. An extensive literature about financial contagion in financial crisis examines the consequences for cross-nation contagion (e.g. Aloui et al., 2011; Baur, 2012). Less is known about the spillovers from stock markets to option markets in U.S. Because the financial sector is most vulnerable and subject to inside information during financial crisis, we examine the transmission from stock markets to option markets in U.S. banking sector around the financial crisis. That is, we use daily data of the financial sector in the U.S. during financial crisis in 2008 to examine the market efficiency between option and stock markets.

Choy and Wei (2012) argue that abnormal option turnovers and abnormal stock returns are significantly related (positively for calls and negatively for puts) around earnings announcements. Once they control for pre-announcement returns, the pre-announcement turnovers no longer predicts post-announcement returns. Hence, option trading doesn't appear to be driven by information around earnings announcements. Nevertheless, Hu (2014) decompose stock order imbalance into the option-induced imbalance and the imbalance independent of option trading activities. He finds that stock exposure imbalance induced by option transactions has strong predictive power of stock returns, while the independent stock order imbalance has a transitory price impact but shows no significant predictive ability for stock returns on the next day. We further explore whether the stock order imbalance has influence on option returns.

Lamoureux and Lastrapes (1990) employ a time-varying Autoregressive Conditional Heteroskedasticity (ARCH) model to test the relation between daily stock returns and trading volume and show that trading volume is a significant explanatory variable on the variance of daily returns. Gallant et al. (1992) shows that correlation between conditional volatility and volume is positive. Moreover, larger price movements are followed by higher trading volumes. Thus, we use a GARCH model to examine the relation between volatility, order imbalance, and return around financial crisis.

We find that contemporaneous stock imbalances have a significant impact on option returns, and lagged-one imbalances also have a significantly positive impact on call but not on put. Conditional on contemporaneous imbalance, the impact of lagged-one imbalance on call return is still positive. Employing a time-varying GARCH model based on the argument of a volatile market in financial crisis, we find that volatility plays an important role in the return-order imbalance relation. Moreover, we develop an imbalance-based call (put) trading strategy that buys the call (put) if the previous day's stock imbalance is positive, and sells the stock if the previous day's stock imbalance is negative. This trading strategy does not outperform the original buy and hold return.

The remainder of this paper is organized as follows. In the literature review section, we present the related literature. The data and methodology section describes the data and defines the variables. In results and discussion section, we show the regression results and discussions. Finally, the conclusion comments section provides our conclusion and suggestion.

LITERATURE REVIEW

Chordia et al. (2002) document that market order imbalances, defined as aggregated daily market purchase orders less sell orders, are positively autocorrelated. Therefore, investors continue buying or selling for a period, suggesting that traders are herding, or splitting large orders over time. Chordia and Subrahmanyam (2004) study the relation between stock order imbalances and daily returns of individual stocks. Price pressures caused by autocorrelated imbalances cause a positive relation between lagged imbalances and returns. Informed traders can use the imbalance-based trading strategies to yield statistically significant returns.

Instead of trading stocks, informed traders may trade options for the following reasons. First, greater leverage, lower transaction costs and built-in downside protection may attract informed traders to participate in the option market (e.g. Chakravarty et al., 2004). Second, investors who only have private information about the volatility of underlying security price could bet on such volatility in the option market (e.g. Ni et al., 2008). Third, the option market provides strategic flexibility to informed traders because they can trade contracts on the same underlying security but with different exercise prices and maturities (Kaul et al., 2004). On the contrary, lower liquidity may discourage informed traders from trading options (Vijh, 1990). Overall, Easley et al. (1998) indicated that informed traders may simultaneously trade in options and underlying stock markets. Further, Chakravarty et al. (2004) and Kaul et al. (2004) suggest that informed traders would trade-off between options leverage and the transaction costs associated with options liquidity.

Although there is informed trading in the option market, it does not necessarily imply that there is no market efficiency, because option volume is not publicly observable. Information-based models proposed by Easley et al. (1998) imply that prices react immediately to public information contained in the trading process but adjust slowly to the private information possessed by informed traders. Vega (2006) shows that not all information acquisition variables have the same effect on the market's efficiency. Cao and Wei (2010) find that information asymmetry is greater for options than for the underlying stocks, implying that the options market is a more efficient venue for informed traders.

DATA AND METHODOLOGY

We choose the largest nine investment and commercial banks to represent the U.S. banking sector, with a sample from June 1, through December 31, 2008. We collect intraday data from NYSE TAQ (Trade and Automated Quotations) and daily data from Option Metrics. Sample stocks and options are included or excluded in our samples according to the following criteria. First, all objective included in our sample must be large investment banks and commercial banks in the U.S., including Goldman Sachs, Merrill Lynch, Citi Group, Morgan Stanley, JP Morgan, Bank of America, Wells Fargo, American Express and Lehman Brothers. Second, all objectives whose transaction data are not available in both NYSE TAQ and Option Metrics are excluded from our samples. Third, to avoid noise trading, we delete those transaction recorded within the first 90 seconds after the market opens.

We employ Lee and Ready (1991) trade assignment algorithm to determine whether it is buyer-initiated or seller-initiated transaction. The average return of call options is -452.48%, with a median of -288.80%. The standard deviation of call option returns is 529.40%, with a maximum value is 68.35% and the minimum is -2393.52%. On the other hand, the average return of puts is -10.52%, with a median of 11.56%. The standard deviation of put option returns is 250.38%, with a maximum of 351.98% and minimum of 1068.90%.

We employ a time-varying GARCH model to re-examine the above relations based on the argument that volatility might play an important role in financial crisis.

$$R_t = \alpha + \beta * OI_t + \varepsilon_t \qquad \varepsilon_t \mid \Omega_t \sim N(0, h_t) \qquad h_t = A + B\ h_{t-1} + C\ \varepsilon_{t-1}^2 \tag{1}$$

Where R_t is the return at time t, and is defined as ln (Pt/Pt-1). OI_t denotes the explanatory variable of order imbalance. β is the coefficient describing the impact of the order imbalance on option return. ε_t is the residual value of the option return at time t. h_t is the conditional variance at time t. Ω_{t-1} is the information set in at time t.

To investigate volatility-order imbalance relation, we use another GARCH model.

$$R_t = \alpha + \varepsilon_t\ , \qquad \varepsilon_t \mid \Omega_t \sim N(0, h_t)\ ,\ h_t = A + B\ h_{t-1} + C\ \varepsilon_{t-1}^2 + \gamma\ OI_t \tag{2}$$

Where R_t is the return at time t, and is defined as ln (Pt/Pt-1). OI_t denotes the explanatory variable of order imbalance. ε_t is the residual value of the stock return at time t. h_t is the conditional variance at time t. Ω_{t-1} is the information set in at time t. γ is the coefficient describing the impact of the order imbalance on volatility of the return.

RESULTS AND DISCUSSION

Chordia and Subrahmanyam (2004) propose that lagged stock order imbalances are positively related to stock returns. We investigate unconditional return-order imbalance relation by regressing option return on five lagged stock imbalances. We expect the relation between call (put) option returns and stock order imbalance is po6sitive (negative). The results are presented in Table 1.

At the 5% significant level, the positive and significant percent of lagged-one imbalance is 21.0%, while the negative and significant one is null. Surprisingly, there is no significant relation between lagged-one imbalance and put return. The empirical result of call options is consistent with Chordia and Subrahmanyam (2004), while put option results show a totally different picture. A possible explanation is as follows. First of all, we find an asymmetric return pattern in option, namely, there is no upper limit for a long option.

Table 1: Unconditional Lagged Return-Order Imbalance Relation

	Positive	Positive and Significant	Negative and Significant
Panel A: Call			
OI_{t-1}	82.72%	21.0%	0.0%
OI_{t-2}	33.33%	2.5%	4.9%
OI_{t-3}	75.31%	2.5%	0.0%
OI_{t-4}	16.05%	3.7%	16.0%
OI_{t-5}	33.33%	0.0%	18.5%
Panel B: Put			
OI_{t-1}	44.44%	0.0%	0.0%
OI_{t-2}	71.60%	3.7%	0.0%
OI_{t-3}	33.33%	0.0%	1.2%
OI_{t-4}	88.89%	8.6%	3.7%
OI_{t-5}	43.21%	7.4%	0.0%

This table shows the regression estimates of lagged order imbalances on the current stock return of the individual stock. Panels A and B present the results in call and put options respectively. "Significant" denotes significance at the 5% level.

During financial crises, markets tend to be more volatile. Investors are prone to long a call instead of writing a call. In addition, market makers react in a different ways to a positive or a negative imbalance. Market makers with call options tend to have a stronger reaction on a positive than a negative imbalance. On the other hand, market makers with put options are prone to overreact more to a negative than a positive imbalance. Secondly, market participants show a pessimistic attitude toward the financial sector in financial crises. Market makers tend to keep a high level inventory of puts. When a large negative order imbalance shows up, option market makers interpret it as a noise trade. Therefore, from a daily perspective, market makers are reluctant to lower bid-ask spreads on put options in the following period to accommodate a negative imbalance. On the other hand, market makers with put option are short call option. When a large positive order imbalance appears, market makers tend to interpret it as a discretionary trade. Therefore, market makers are eager to raise bid-ask spreads on calls.

We also examine conditional return-order imbalance through a contemporaneous imbalance and four lags of order imbalances. The empirical result is exhibited in Table 2. At the 5% significant level, the positive and significant percent of contemporaneous and lagged-one imbalance are 76.5% and 24.7% respectively. We conclude that contemporaneous relations between imbalances and returns confirms both inventory and asymmetric information effects of price formation. However, our empirical results are inconsistent with Chordia and Subrahmanyam (2004), which predict a reversal return-lagged imbalance relation. There is a possible explanation: When a large positive imbalance appears, option market makers tend to interpret it as a discretionary trade and rush to raise bid-ask immediately because markets are in a panic. Nonetheless, they quickly find themselves overreacted. Therefore, they slow the price adjustments.

Table 2: Conditional Lagged Return-Order Imbalance Relation

	Positive	Positive and Significant	Negative and Significant
Panel A: Call			
OI_t	100.00%	76.5%	0.0%
OI_{t-1}	90.12%	24.7%	0.0%
OI_{t-2}	28.40%	3.7%	8.6%
OI_{t-3}	59.26%	1.2%	0.0%
OI_{t-4}	33.33%	3.7%	13.6%
Panel B: Put			
OI_t	0.0%	0.0%	82.7%
OI_{t-2}	25.93%	0.0%	0.0%
OI_{t-3}	74.07%	2.5%	0.0%
OI_{t-4}	43.21%	0.0%	1.2%
OI_{t-5}	70.37%	7.4%	3.7%

This table shows the regression estimates of current and lagged order imbalances on the current stock return of the individual stock. Panels A and B present the results in call and put options respectively. "Significant" denotes significance at the 5% level.

As for put options, the negative and significant percent of contemporaneous and lagged-one imbalances are 82.7% and 0% at the 5% significant level. One possible explanation is that, during a severe financial crisis with a pessimistic environment, market makers regard large negative imbalance as noise trading. Therefore, market makers in the option market lower quotes of put options to manage inventory levels in the same trading day, and do not further adjust quotes significantly to react to imbalances of the last trading day.

Does volatility play a role in return-order imbalance relation, especially in a financial turbulence? We employ a time-varying GARCH model to capture time-variant properties in the return-order imbalance relation. Table 3 presents the results that the positive and significant percent of calls is 79.0% and the negative and significant percent of puts is 81.5% at 5% significant level. These results reconfirm the return-order imbalance relation. If option markets are efficient, there should be no significant GARCH relation between stock order imbalances and option returns in financial crisis.

Table 3: Dynamic Return-Order Imbalance GARCH (1,1) Relation

	Percent	Percent Positive and Significant	Percent Negative and Significant
Call	95.1%	79.0%	0.00%
Put	4.9%	1.2%	81.5%

This table shows the regression estimates of the equation. $R_t = \alpha + \beta * OI_t + \varepsilon_t \quad \varepsilon_t \mid \Omega_t \sim N(0, h_t) \quad h_t = A + B\ h_{t-1} + C\ \varepsilon^2_{t-1}$ *where R_t is the return in period t, and is defined as $ln(P_t/P_{t-1})$, OI_t is the order imbalance, ε_t is the residual value of the stock return in period t, Ω_{t-1} is the information set in period t-1. "Significant" denotes significant at the 5% level.*

We take a further step to examine volatility-order imbalance relation through a time-varying GARCH model. The empirical results are exhibited in Table 4. The prior belief is that the larger order imbalance is associated with the higher volatility. Nonetheless, the imbalance impact on volatility is not what we had thought. For call, we document that the positive and significant percent is 33.3% and negative and significant percent is 32.1% at 5% significant level respectively. For put, we find that the positive and significant percent is 33.3% and negative and significant percent is 33.3% at 5% significant level respectively. It implies that market makers have sufficient inventories to mitigate option market volatility.

Table 4: Dynamic Volatility-Order Imbalance GARCH (1,1) Relation

Panel A: By Option Type			
	Percent	**Percent Positive and Significant**	**Percent Negative and Significant**
Call	50.6%	33.3%	32.1%
Put	45.7%	33.3%	33.3%
Panel B: By Bank			
χ	**Percent**	**Percent Positive and Significant**	**Percent Negative and Significant**
AXP	38.9%	33.3%	11.1%
BAC	22.2%	11.1%	55.6%
C	88.9%	77.8%	0.0%
GS	27.8%	22.2%	55.6%
JPM	16.7%	11.1%	50.0%
LEH	50.0%	5.6%	27.8%
MER	72.2%	50.0%	22.2%
MS	16.7%	11.1%	72.2%
WFC	100.0%	77.8%	0.0%

This table shows the regression estimates of the equation. $R_t = \alpha + \varepsilon_t \quad \varepsilon_t / \Omega_{t-1} \sim N(0, h_t),\ h_t = A + Bh_{t-1} + C\varepsilon^2_{t-1} + \gamma * OI_t$ *where R_t is the return in period t, and OI_t is the order imbalance ε_t is the residual value of the stock return in period t, Ω_{t-1} is the information set in period t-1. Panel A presents the results by option type and Panel B shows the results by bank. "Significant" denotes significant at the 5% level.*

We further regroup our sample into nine small groups according to their institution to demonstrate market maker behaviors in different option markets. The results show variable abilities of market makers in different markets. We find the positive and significant percent is 77.8%, 50.0%, and 77.8% at 5% significant level on Citi (C), Merrill Lynch (MER), and Wells Fargo (WFC). This result implies that market makers are poor at mitigating volatility in these markets. The empirical results are consistent with the intuition that these three financial institutions are deeply involved in the financial crisis. On the other hand, we show the negative and significant percent is 55.6%, 50.0%, and 72.2% in options of Bank of America (BAC), JP Morgan (JPM), and Morgan Stanley (MS) respectively.

To test spillover efficiency, we develop an order imbalance-based trading strategy. In our sample period, the average return on calls is -452.5%, while that on puts is 82.23%. We implement our imbalance-based trading strategy as following. First, we trim 90% noisy trades on liquidity on a daily bases. Then we long a call when a positive stock order imbalance appears and long a put when a negative imbalance appears. We hold the position until reversal appears. The performance is exhibited in Panel A of Table 5. Through this trading strategy, the returns are -179.56% and -26.42% for calls and puts respectively. At the 1% significant level, the return of trading strategy for calls is significantly negative, whereas the return of trading strategy for puts is insignificant negative. This result is self-explained in the financial crisis. Panel B shows the return of imbalance-based trading strategy for calls is significantly lower than open-to-close returns at the

1% significant level. In Panel C, we find significant differences in returns between the two strategies at the 1% significant level.

Table 5: Trading Profit under the Basis of Quote price

Panel A: Returns Compared with Zero		
	P-value	
Call	0.0001***	
Put	0.1041***	
Panel B: Returns Compared with Returns of Buy-and-hold Strategy		
	Mean Original	P-value
Original open-to-close return	-0.1050	
Call Return of strategy	-1.7956	0.0001***
Put Return of strategy	-0.2642	0.2273
Panel C: Differences in Returns between Two Strategies		
P-value	Mean	Two-tail P-value
Call return of strategy	-1.7956	0.0001***
Put return of strategy	-0.2642	

This table shows the trading profit under the quote price. We long a call when a positive stock order imbalance appears and long a put on a negative one. We hold the position until reversal shows up. Panel A presents the p-values to be used to examine whether the return of imbalance-based trading strategy is positive. Panel B shows the p-values to be used to explore whether the return of imbalance-based trading strategy is higher than open-to-close return. Panel C exhibits the p-values to be used to examine whether there is no difference in return between two strategies.

CONCLUDING COMMENTS

An extensive literature about financial contagion in financial crises examines the consequences across nations. Less is known about the spillovers from stock markets to option markets in the U.S. Since financial sector is most vulnerable and subject to inside information during financial crises, we focus on bank stocks and option transmissions during the financial crisis of 2008. From daily trading of nine leading banks in the U.S. from June 1, 2008 to December 31, 2008, we first examine the relations between returns and lagged order imbalances by using an OLS model. We find that lagged imbalance is a good predictor of calls, which is consistent with Chordia and Subrahmanyam (2004). We examine the role of volatility on return-order imbalance relations in financial crisis and find that the relations are significant. From a volatility-order imbalance GARCH model, we argue that market makers are capable of mitigating volatility. After examining relations among returns, order imbalance, and volatility, we develop an imbalance-based trading strategy. Our strategy suggests longing a call (put) if the previous day's stock imbalance is positive, and selling a stock if the previous day's stock imbalance is negative. We document that imbalance-based trading strategies on call options cannot beat the market. It implies market efficiency on bank stocks and option transmissions.

This paper focuses on the impact of stock order imbalances on option returns in financial crisis. Future research could examine the influence of option order imbalances on stock returns to enrich the literature.

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EXECUTIVE COMPENSATION STICKINESS AND PEER GROUP BENCHMARKS: EVIDENCE FROM CHINESE FIRMS

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ABSTRACT

This paper examines the phenomenon and effect of peer group on executive compensation stickiness in China's listed firms. We find there has been substantial growth in executive compensation in the past 10 years. Consistent with agency theory, executive compensation is positively related to firm performance. However, pay-for-performance sensitivity is asymmetric, and it is lower when firm performance declines suggesting that there is a characteristic of executive compensation stickiness in Chinese firms. Further, we test the effect of peer group on compensation stickiness. We find that the characteristic of compensation stickiness only exists in the firms whose executive compensation is lower than the compensation of peer group. The evidence suggests that compensation stickiness is an important mechanism to provide retention incentives to firm managers, rather than an agency problem in Chinese firms.

JEL: G34

KEYWORDS: Corporate Governance, Pay-For-Performance Sensitivity, Peer Group, Compensation Stickiness

INTRODUCTION

There are two main strands of research on the relationship between executive compensation and firm performance. The first focuses on the effects of compensation as a tool to decrease the agency cost, and they suggest that the compensation of managers should be linked to firm performance to motivate CEOs to maximize shareholder value (Jensen and Meckling, 1976; Jensen and Murphy, 1990; Murphy, 1999). The second strand focuses on the effects of the managerial power on the compensation, and they suggest that managers use their power to influence the compensation-setting process and acquire more compensation than on the basis of firm performance (Bebchuk et al, 2002; Bebchuk and Fried, 2004). In recent years, some scholars have begun to research on the compensation stickiness, which means the relationship between executive compensation and firm performance is asymmetric where managers are rewarded more for good performance but are punished less for bad performance. They point out that the compensation stickiness is a representation of the problem of corporate governance, and it is harmful to the value of companies (Garvey and Milbourn, 2006; Fang, 2009).

The purpose of this study is to investigate the phenomenon of executive compensation stickiness, and understand the implication of compensation stickiness in Chinese firms. China began a process of market economic reforms in the late 1970s and these reforms are being continued today. The reform process has brought about substantial changes in institutional arrangements for top executive compensation. Prior to market economic reforms, there was a highly structured pay scale system for the managers (Firth et al., 2006), and it had no incentive mechanism to motivate managers. Therefore, early studies do not find any relationship between executive compensation and firm performance in China (Huang and Zhang, 1998;

Wei, 2000). With the enactment of the Company Law and The Code of Corporate Governance for Listed Firms in China in 2002, executive compensation had a major impact, and executives were allowed to share in the profits generated by the companies. Most executive pay is composed of cash salary and bonus at present, and performance based pay is common in China, although very few firms have executive stock option schemes (Firth et al., 2006; Fang, 2009). According to a survey, more than 80% of the firms use incentive pay systems based on performance (Liu and Otsuka, 2004). These reforms on executive compensation are conducive to aligning interests of managers and owners. A series of recent studies reported a positive relationship between management pay and firm performance after the compensation reforms (Firth et al., 2006; Conyon, 2011; Cao, 2011).

Despite China having introduced market reforms some 30 years ago and having become the largest emerging economy with the fastest growing stock market, corporate governance and compensation of top executives are still at the developmental stage. Most listed firms in China originated from partial privatizations, and the Chinese government usually tends to be the largest shareholder in these firms (Sun and Tong, 2003). The government ownership is represented by various entities such as government agencies, state asset management companies, and state-owned enterprises (SOEs) (Chen et al., 2009). These representatives have their own interests, rather than maximize the value of listed firms, and they cannot provide effective monitoring of management. In addition, the product market does not function freely, and the takeover market for firms is almost non-existent in China. Insider control worsens the agency problems, and managers tend to be opportunistic and seek personal benefit rather than company success (Lin et al., 1998). Although, appointment committees, compensation committees, and other committees are now becoming common in China's listed firms, the effectiveness of them has yet to be fully understood (Chen et al., 2006). In this situation, managers are likely to influence the boards of directors in the pay process, and gain relatively more for good performance but lose less for bad performance.

On the other hand, since the managerial labor market is not well established and there is a shortage of high quality managers in China, management turnover could be more costly for Chinese firms. There is evidence that boards of directors are concerned about managerial retention. Therefore, the boards are more likely to be reluctant to decrease the top executive pay to retain talented managers when firm performance decreases, especially when executive compensation is lower relative to their peer group. There is evidence that in lower pay companies, the boards will reduce the pay-for-performance sensitivity when performance is down, but increase pay-for-performance sensitivity when performance is up. In this aspect, compensation stickiness may be considered a mechanism to provide retention incentives to managers, rather than an agency problem in China's listed firms. The remainder of this paper is organized as follows: The next section is the literature review. Section 3 discusses the methodology and data. Section 4 examines the relationship between growth of top executive compensation and firm performance. It also examines the sticky characteristic in executive compensation. Based on these analyses, we examine the role of peer group on executive compensation stickiness. Section 5 offers further checks and Section 6 concludes the paper.

LITERATURE REVIEW

The most important function of compensation committees is that design suitable compensation arrangements for corporate managers. Agency theory suggests that executive compensation should be linked to firm performance to motivate CEOs to maximize shareholder value (Jensen and Meckling, 1976; Jensen and Murphy, 1990; Murphy, 1999). However, Bebchuk et al (2002) pointed out that executives have power to influence their own pay, and they use that power to extract rents. Furthermore, the desire to camouflage rent extraction might lead to the use of inefficient pay arrangements that provide suboptimal incentives and thereby hurt shareholder value. In addition, Gaver and Gaver (1998) found most top executives in American firms were rewarded with extra bonuses when firm performance increased, but were not penalized when firm performance decreased. Garvey and Milbourn (2006) split executives pay into two elements: pay for luck and pay for skills, and they argued there is an asymmetry benchmarking in

executive compensation: executives are rewarded for good luck while they protect themselves from being penalized for bad luck (i.e., executive compensation increases for good industry performance is larger than marginal decreases for bad industry performance). Fang (2009) also found there is a sticky characteristic in top executive compensation in Chinese firms, and he pointed that it is a signal of poor corporate governance. That means managers trend to influence the compensation-setting process, and get the best compensation arrangements for themselves in China.

In fact, however, firm performance is not the only consideration factor in the executive pay-setting process. An important method used to determine executive compensation is to compare the current level of compensation with the compensation at a peer group of similar companies. Most compensation committees consider the peer group factor in the pay-setting process, and affirm that pay above the median of the peer group is competitive and below the median is below market (Bizjak et al., 2008). In a contrasting viewpoint, it is an efficient way to determine the reservation wage of the CEO (Holmstrom and Kaplan, 2003). That means, when firm performance decreases, the board of directors cannot decrease their top executive compensation immediately. They need to compare it with the compensation in similar companies, and decrease it less if it is lower than other executive pay in similar companies.

Bizjak et al. (2008) provide empirical evidence on the relation between executive compensation and peer group compensation. They find that the use of peer group compensation is widespread and has a significant impact on CEO compensation. Faulkender and Yang (2010) also find that peer group compensation can affect the executive compensation. In addition, they analyze the features of the disclosed compensation peer groups, and they find that firms appear to select highly paid peers to justify their CEO compensation. Bizjak et al. (2011) further point out that although peer groups are largely selected based on characteristics that reflect the labor market for managerial talent, but they are constructed in a manner that biases compensation upward, particularly in firms outside the Standard & Poor's (S&P) 500. Albuquerque et al.(2013) argue that peer group compensation is an efficient mechanism used to gauge the market wage necessary to retain valuable human capital, and the choice of highly paid peers represents a reward for unobserved CEO talent. Although some studies have focused on the existence of executive compensation stickiness, few have further analyzed the implications of it. Why is compensation stickiness widespread? Is it a problem of corporate governance or an important mechanism of compensation design? In this paper, we explore the implications of compensation stickiness in the background of the Chinese institutional environment.

METHODOLOGY AND DATA

Relationship between Compensation Growth and Firm Performance

We employ panel data regression with fixed-effects to estimate our model. This approach controls for heterogeneity in firm quality, and the problem of omission of variables that could cause statistical bias in the estimation of parameters in our models. To test the relationship between pay and performance we use the following regression model:

$$\Delta PAY_{it} = \alpha_i + \beta_{it}\Delta PERF_{it} + ControlX_{it} + \varepsilon \quad (1)$$

where, ΔPAY_{it} is the growth of executive cash compensation in firm i at time t, and it equals $PAY_t - PAY_{t-1}$.

Executive compensation is the aggregated pay of the top three officers, not only the CEO, defined as the sum of basic salary, bonus, stipends, and other benefits to the top three officers. We only use the cash compensation of the top officers, because it is the most important part of incentive contract in Chinese companies, and it is consistent with previous research in China (Firth et al., 2007; Kato et al., 2006; Wang and Xiao, 2011; Conyon and He 2011). The main independent variable is $\Delta PERF_{it}$, which is the change of

firm performance in firm i at time t, and it equals $PERF_{it} - PERF_{it-1}$. Firm performance is defined in two ways. First, we use the change in stock value (stock $value_t$ - stock $value_{t-1}$) following Jensen and Murphy (1990). Second, we use an accounting-based measure of performance. Fang (2009) points out that most boards of directors in Chinese companies use net profit-based criteria to evaluate firm performance and management contribution. We select the change of net profit and net profit (excluding extraordinary gains or losses) to represent accounting-based performance. It is expected that these independent variables are positively related with compensation; i.e., $\beta_1 > 0$, regardless of the chosen performance variable. In addition, following previous studies (e.g. Firth et al., 2007; Faulkender and Yang, 2010; Conyon and He 2011), a set of control variables, X_{it}, are also included in model (1), and these variables are defined in Table 1.

Table 1: Variable Descriptions

Variable	Description
PAY_t	Total executive compensation, i.e., the aggregated pay of the top three officers, defined as the sum of basic salary, bonus, stipends, and other benefits.
ΔPAY_t	Periodic change in PAY or $PAY_t - PAY_{t-1}$
$\Delta WEALTH_t$	Change in stock value or stock $value_t$ - $stockvalue_{t-1}$
$WDOWN_t$	Dummy variable of unity if the ΔWEALTH < 0, and zero otherwise.
$PROFIT1_t$	Net profit in year t
$\Delta PROFIT1_t$	Change in net profit or $profit1_t$ - net $profit1_{t-1}$
$P1DOWN_t$	Dummy variable of unity if the ΔPROFIT1 < 0, and zero otherwise.
$PROFIT2_t$	net profit (excluding extraordinary gains or losses) in year $_t$
$\Delta PROFIT2_t$	Change in net profit (excluding extraordinary gains or losses) or $profit2_t$ - net $profit2_{t-1}$
$P2DOWN_t$	Dummy variable of unity if the ΔPROFIT2 < 0, and zero otherwise.
$\Delta ASSET_t$	Change in total assets or total $asset_t$ – total $asset_{t-1}$
$\Delta SALES_t$	Change in sales or $sales_t$ –$sales_{t-1}$
$LN(S)_{t-1}$	Natural log of sales, used to measure firm size
$GROWTH_{t-1}$	The market value of the firm divided by the book value of assets (market-to-book), used to measure firm growth opportunities
$LG1\_OWN_t$	Share ownership of the largest shareholder, used to measure ownership concentration
$STATA\_OWN_t$	Share ownership of the State, used to measure the degree of government control.
$BOARD\_SIZE_t$	Number of individuals on the main board
IND_DIR_t	Percentage of independent directors in the board
$COMBINE_t$	Equal to one if the posts of CEO and chairperson are combined, and zero otherwise
EPS_t	Earnings per share, used to measure the profitability of equity
LEV_t	The ratio of total liabilities to total assets, used to measure the degree of financial leverage

This table shows the definitions of variables.

Compensation Stickiness

To test compensation stickiness in Chinese firms we introduce a dummy variable $PERDOWN_{it}$ in regression model, which equals to one if performance decreases ($\Delta PERF_{it} < 0$) in firm i at time t, and otherwise equal to zero. The model is:

$$\Delta PAY_{it} = \alpha_i + \beta_1 \Delta PERF_{it} + \beta_2 PERDOWN_{it} + \beta_3 \Delta PERF_{it} \times PERDOWN_{it} + ControlX_{it} + \varepsilon \quad (2)$$

Given our use of three performance measures, there are three $PERDOWN_{it}$ variables (WDOWN, P1DOWN, and P2DOWN) in model (2). Following Garvey and Milbourn (2006), we include an interaction-term, $PERF_{it} \times PERDOWN_{it}$ to make specific inferences regarding the relationship between compensation and its drivers in model (2). If executives are paid for good performance but not punished equally for bad performance, the pay-performance sensitivities will be lower when performance is bad. If there is pay-performance stickiness in Chinese firms, then β_3 should be negative, regardless of the performance variable we choose. The control variables, X_{it}, remain the same as in model (1).

Effect of Peer Group

To analyze the influence of peer groups on compensation stickiness we construct a peer group for every company's executives, and rank them according to compensation levels of their peers. Firms seldom report the composition of their compensation peer groups in China, but most peer groups are selected by similar size in similar industries（Bizjak et. al, 2008）. Following Bizjak et. al (2008), we divide our sample into 21 industries (There are 13 "industries" in China: A: Agriculture and fishery, B: Mining, C: Manufacturing; D: Electricity, water and other energy manufacturing and supply; E: Construction; F: Transportation and logistics; G: Information technology; H: Wholesale and retail; I: Finance and Insurance; J: Real estate; K: Service; L: Communication; M: Others. Manufacturing industry (C) is the biggest in China, and scholars typically divide it into 10 sub-industries. In addition, Due to its highly regulated nature, Finance and Insurance industry (I) is excluded from our sample. As a result, there are 21 "industries" in our sample). For each year, we rank all firms in every industry according to the size (total assets). We classify firms as large (small) firm group if their size is larger (smaller) than the median size in the industry. Thus, we form 42 peer groups for this study. We rank all firms in every peer group based on the level of the prior-period compensation. We classify executives as "lower compensation" if their prior-period compensation is lower than median executive compensation, or else "higher compensation." Thus we form two groups: one is higher compensation, and the other is lower compensation.

Most compensation committees consider pay above the median in a peer group to be competitive and consider compensation levels below the median to be below market (Bizjak et. al, 2008). Because there are retention incentives in firms (to prevent falling below the reservation wage), the firms whose prior executive compensation is below the median tend to be under pressure to increase their executive compensation, regardless of their performance. As a result, there is a significant potentially sticky characteristic in the lower compensation group. This means that, in model (2), we expect the value of β_3 in the lower compensation group to be larger and more significant than that of the higher compensation group.

Sample

Our study uses data of firms listed on the Shanghai Stock Exchange or the Shenzhen Stock Exchange, and the sample period is from 1/1/2002 to 12/31/2012. We collect annual data from the China Center for Economic Research Sinofin Information Service (CCER/SinoFin). This database has been used in previous research of China's listed firms (Cao et al., 2011; Conyon and He, 2011). Our original data consists of almost the entire population of publicly traded firms from years 2002 to 2012. In order to estimate the regression models we required non-missing data on main variables, such as executive compensation, stock value, profit, size, sales, and so on. Also, in our empirical work we analyze the relationship between the change of compensation ($PAY_t - PAY_{t-1}$) and the change of performance ($PERF_t - PERF_{t-1}$), so this requires the firms to have at least two consecutive years of data. In addition, consistent with the previous literature, we exclude firms in the financial services industries, due to their unique accounting standards and incomplete information on the main variables used in our analysis. Our final sample consists of 2050 firms and 13,517 firm-year observations between 2002 and 2012.

Descriptive Statistics

Table 2 presents the descriptive statistics on executive compensation (Panel A) and firm performance (Panel B) for the sample firms. Panel A shows the mean (median) of executive compensation over the 2002-2012 sample period is about 2,652,420 (1,847,000) RMB. This compensation level is much lower than for top executives in the U.S., and other developed countries (Brick et al., 2006; Kato et al., 2007; Wang and Xiao, 2011; Conyon and He, 2011). However, we find a steady increase in executive compensation from 2002 to 2012. In 2002, the mean (median) is 966,140 (729,960) RMB, and it increases to 4,579,570 (3,446,250) RMB in 2012. The growth rate in executive compensation across our sample period is as high as 374% (372%).

Table 2: Descriptive Statistics

Panel A: Executive Compensation(Thousands Rmb)

Year	Mean	Median	Standard Deviation	Number of Observations
2002	966.14	729.96	827.45	726
2003	1,204.28	904.76	1,092.05	975
2004	1,439.15	1,082.60	1,282.77	1,099
2005	1,589.51	1,191.58	1,480.41	1,188
2006	1,753.96	1,328.95	1,511.00	1,186
2007	2,312.21	1,700.00	2,228.81	1,221
2008	2,742.20	1,998.28	2,919.88	1,318
2009	2,995.40	2,260.00	2,784.91	1,383
2010	3,620.29	2,638.00	3,678.94	1,481
2011	4,041.82	3,089.20	3,738.48	1,820
2012	4,579.57	3,446.25	4,461.44	1,120
Total Compensation	2,652.42	1,847.00	3,003.24	

Panel B: Firm Performance (Millions Rmb)

	ΔWEALTH		PROFIT1		PROFIT2	
Year	Mean	Median	Mean	Median	Mean	Median
2002	2,583.91	1,918.73	49.21	28.93	44.52	25.54
2003	2,475.98	1,539.76	60.48	31.33	54.18	29.55
2004	2,151.13	1,276.12	78.00	31.55	70.46	28.54
2005	1,754.72	969.56	70.75	26.26	63.35	23.96
2006	2,889.85	1,449.62	97.35	36.31	72.79	26.62
2007	9,037.65	4,154.80	199.71	72.72	132.65	39.51
2008	3,547.90	1,683.22	153.63	54.86	108.78	31.69
2009	7,746.56	4,104.02	200.22	70.29	141.43	44.92
2010	8,456.90	4,522.45	314.86	97.13	231.70	68.44
2011	5,812.43	3,025.58	315.86	94.92	230.37	67.34
2012	6,074.80	3,064.36	296.76	91.69	215.49	67.35
Total Sample	5,067.93	2,398.75	183.23	56.63	135.30	39.99

This table shows the descriptive statistics results. Panel A shows results for executive compensation. Panel B shows results for firm performance. Data are for the period 1/1/2001 until 12/31/2012.

Performance statistics of the firms (panel B) show that the mean (median) of the change in stock value is 5,067,930,000 (2,398,750,000) RMB. The mean (median) of net profit is 183,230,000 (56,630,000) RMB. The mean (median) of net profit (excluding the extraordinary gains or losses) is 135,300,000 (39,990,000) RMB. We also find a steady increase in executive compensation from 2002 to 2012. The growth rates across our sample period are 135% (59%), 503% (217%), 384% (164%), respectively. To estimate the annual growth rates of executive compensation and firm performance, we run regressions of the logarithm of executive compensation and the logarithm of performance on a linear time trend. The results indicate that executive compensation has grown by about 16% per year ($\beta = 0.157$, $t = 71.52$), stock value has grown by about 12% per year ($\beta = 0.123$, $t = 45.78$), net profit has grown by about 14% per year ($\beta = 0.135$, $t = 29.47$), and net profit (excluding the extraordinary gains or losses) has grown by about 13% per year ($\beta = 0.129$, $t = 27.08$). We observe that the growth rate of executive compensation has grown faster than firm performance, and the time trend of the growth of compensation is more significant than the performance variable.

RESULTS

Pay-For-Performance Sensitivities

Table 3 reports estimates of models (1) and (2). Column (1), (3), and (5) show the relationship between executive compensation and firm performance. Column (1) uses change in stock value as the measure of performance, column (3) uses change in net profit, and column (5) uses change in net profit (excluding the extraordinary gains or losses). As predicted, the regression coefficient on performance is positive ($\beta_1 > 0$) in all three performance measures.

Table 3: Regression Results of Pay-For-Performance Sensitivities

Independent Variables	**Dependent Variable: ΔPay**					
	(1)	(2)	(3)	(4)	(5)	(6)
Δwealth	0.01***	0.01***				
	(4.66)	(4.38)				
wdown		-190.94***				
		(-4.24)				
Δwealth×wdown		-0.01***				
		(-2.33)				
Δprofit1			0.18***	0.19***		
			(5.52)	(5.67)		
p1down				-71.28		
				(-1.22)		
Δprofit1×p1down				-0.23		
				(-1.44)		
Δprofit2					0.38***	0.40***
					(7.82)	(7.91)
p2down						-133.95***
						(-2.62)
Δprofit2×p1down						-0.35**
						(-2.08)
Δasset	0.00	0.00	0.00	0.00	0.00	0.00
	(0.11)	(0.17)	(0.26)	(0.28)	(0.27)	(0.30)
Δsales	0.02***	0.02***	0.00	0.00	0.00	0.00
	(2.79)	(2.74)	(0.06)	(-0.02)	(-0.34)	(-0.44)
ln(s)	58.87***	60.32***	48.24***	46.71***	42.38**	36.88**
	(3.28)	(3.35)	(2.68)	(2.59)	(2.36)	(2.04)
Growth	176.88	121.55	201.74*	205.99*	223.91*	228.29**
	(1.56)	(1.05)	(1.78)	(1.82)	(1.98)	(2.02)
lg1_own	531.58**	517.28**	503.02**	493.56**	481.53**	453.38**
	(2.47)	(2.40)	(2.33)	(2.29)	(2.24)	(2.10)
stata_own	1.33	1.21	1.47	1.49	1.41	1.40
	(1.41)	(1.27)	(1.55)	(1.58)	(1.50)	(1.48)
board_size	11.12*	10.46*	10.55*	10.29*	9.92*	9.35
	(1.88)	(1.76)	(1.78)	(1.73)	(1.68)	(1.58)
ind_dir	0.56	0.51	0.29	0.29	0.05	0.13
	(0.25)	(0.23)	(0.13)	(0.13)	(0.02)	(0.06)
Combine	43.61	37.29	37.37	37.58	32.75	32.22
	(0.84)	(0.72)	(0.72)	(0.73)	(0.63)	(0.62)
Eps	265.66***	245.02***	216.30***	220.30***	179.97***	177.53***
	(7.68)	(7.04)	(5.99)	(5.30)	(4.93)	(4.41)
lev	22.61	22.82	21.80	22.29	21.54	24.46
	(0.89)	(0.89)	(0.85)	(0.87)	(0.85)	(0.96)
#obs.	13517	13517	13517	13517	13517	13517
r^2	0.033	0.036	0.033	0.033	0.034	0.035

*This table shows regression results based on equation (1) and (2): (2) The first figure in each cell is the regression coefficient. The second figure in each cell is the t-statistic. *, **, *** indicate significance at 10%, 5% and 1% level, respectively. We also control industries fixed-effect, year fixed-effect, and firm fixed-effect in our model.*

Executive compensation in our data is in thousands of RMB, and performance is in millions of RMB. Therefore, the regression coefficient β_1 can be interpreted as the RMB increase in executive compensation for every 1,000 RMB increase in firm performance. The coefficient β_1 in Column (1) implies that the executive compensation will be 0.01 RMB more per 1,000 RMB increase in shareholder value, all else equal. This result is much lower than the reported in research for the U.S., U.K., and other countries (Jensen

and Murphy, 1990; Brick et al., 2006; Merhebi et al., 2006). The coefficient β_1 in column (3) (column (5)) imply that the executive compensation will be 0.18 (0.38) RMB more per 1,000 RMB increase in net profit (net profit (excluding the extraordinary gains or losses)), all else equal. We find the pay-for-performance sensitivity is much higher when we used accounting-based measures of performance. This result is consistent with the argument of Fang (2009), and indicates that most boards of directors in Chinese use net profit-based criteria to evaluate firm performance and management contribution, especially the net profit excluding the extraordinary gains or losses. These results show that the growth of executive compensation is associated with firm performance and it could be interpreted as an important motivation mechanism to make executives focus on annual firm performance and management in China.

Compensation Stickiness

In order to test the executive compensation stickiness, we observe the coefficients of PERFDOWN dummy variables and interaction-term of firm performance and PERFDOWN dummy variables in column (2), (4), and (6) of Table 3. The coefficient on PERFDOWN is negative (-190.94, -71.28, and -133.95, respectively), and most PERFDOWN dummy variables are significant (except P1DOWN), implying that executive compensation decreases when performance is bad. We also find that the coefficients on the interaction-term of firm performance and PERFDOWN dummy variables are negative ($\beta_3 < 0$), and significant in column (2) and (6). As anticipated, the pay-performance sensitivity is lower when performance is poor.

The regression coefficient β_3 in column (2) is -0.01, and it implies that when stock value decreases, the pay-for-performance sensitivity of executive compensation is equal to $\beta_1 + \beta_3 = 0$ (and the F value of ($\beta_1 + \beta_3$) is statistically insignificant in the regression model). Similarly, the regression coefficient β_3 in column (6) is -0.35, and it implies that when net profit (excluding the extraordinary gains or losses) decreases, the pay-for-performance sensitivity of executive compensation is equal to $\beta_1 + \beta_3 = 0.05$ (and the F value of ($\beta_1 + \beta_3$) is statistically insignificant in the regression model). From these interaction-term regression coefficients we find that the sticky characteristic is substantial, and almost equal to the relationship between executive compensation and firm performance. The result means that the compensation of managers is no longer linked strongly to firm performance when firm performance decreases. Overall, these results indicate that there is a sticky characteristic in executive compensation, and managers are rewarded more for good performance but are punished less for poor performance in Chinese companies. These findings are consistent with studies Gaver and Gaver (1998) and Garvey and Milbourn (2006) in U.S.

Effect of Peer Group

We analyze the effect of peer group on compensation stickiness, and interpret the implication of compensation stickiness in this subsection. Based on the discussion above, we expect that the compensation stickiness is larger when performance is down for executives who are paid below their peers; i.e., β_3 is more negative for executive compensation below the peer group median. To test how the compensation stickiness varies by peer group pay ranking, we divided the sample into two groups: higher compensation group, and lower compensation group, and then we estimated model (2) using these two groups, separately. The results are presented in Table 4. The first (last) three columns represent the alternative measures of firm performance in higher (lower) compensation group. Table 4 makes it clear that there is an asymmetry in compensation stickiness between the two groups. The coefficients on the interaction-term of firm performance and PERFDOWN dummy variables (β_3), as predicted, are larger and more significant in the lower compensation group. Even more important, coefficients of compensation stickiness are not statistically significant in higher compensation group.

Table 4: Regression Results of Pay-For-Performance Sensitivities（Higher Compensation Group Vs. Lower Compensation Group）

Independent Variables	**Dependent Variable: ΔPay**					
	Panel A			**Panel B**		
	Higher Compensation Group			**Lower Compensation Group**		
	(1)	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**
Δwealth	0.01**			0.01***		
	(2.34)			(5.18)		
wdown	-251.60***			-66.27**		
	(-2.98)			(-2.08)		
Δwealth×wdown	-0.01			-0.02***		
	(-1.26)			(-4.06)		
Δprofit1		0.17***			0.72***	
		(3.57)			(11.59)	
p1down		-98.79			-85.74**	
		(-0.77)			(-2.38)	
Δprofit1×p1down		-0.16			-0.80***	
		(-0.59)			(-6.29)	
Δprofit2			0.41***			0.76***
			(5.52)			(10.58)
p2down			-144.29			-113.35***
			(-1.33)			(-3.46)
Δprofit2×p1down			-0.31			-0.86***
			(-1.06)			(-6.17)
Δasset	0.00	0.00	0.00	0.00	-0.01	0.00
	(0.13)	(0.30)	(0.47)	(-0.03)	(-0.77)	(0.05)
Δsales	0.02*	0.00	-0.01	0.01**	0.00	0.01
	(1.89)	(-0.06)	(-0.69)	(2.04)	(0.17)	(0.73)
ln(s)	129.29*	70.34	25.28	66.78***	55.16***	54.41***
	(1.91)	(1.03)	(0.37)	(6.67)	(5.56)	(5.47)
growth	27.96	139.73	188.44	287.37***	294.02***	320.37***
	(0.12)	(0.59)	(0.80)	(3.73)	(3.94)	(4.29)
lg1_own	280.60	239.31	239.51	707.49***	601.72***	586.15***
	(0.58)	(0.49)	(0.49)	(4.97)	(4.25)	(4.12)
stata_own	2.09	2.40	2.19	0.05	0.33	0.41
	(1.08)	(1.24)	(1.13)	(0.07)	(0.52)	(0.63)
board_size	4.16	3.39	2.22	19.75***	18.71***	18.64***
	(0.35)	(0.29)	(0.19)	(4.84)	(4.64)	(4.62)
ind_dir	1.04	0.82	0.51	-0.86	-0.98	-0.88
	(0.24)	(0.19)	(0.12)	(-0.54)	(-0.62)	(-0.56)
combine	24.95	28.29	23.06	30.59	25.47	20.43
	(0.25)	(0.28)	(0.23)	(0.84)	(0.71)	(0.57)
eps	442.40***	407.26	345.79***	105.54***	-2.86	28.53
	(5.93)	(4.68)	(4.04)	(4.54)	(-0.10)	(1.08)
lev	133.19	134.19	130.45	9.91	6.89	9.38
	(0.74)	(0.74)	(0.72)	(0.78)	(0.55)	(0.74)
#obs.	6835	6835	6835	6682	6682	6682
r^2	0.033	0.032	0.033	0.041	0.041	0.042

*This table shows regression results based on equation (2). (2) Panel A shows the results for the sample of higher compensation group. Panel B shows the results for the sample of lower compensation group. The first figure in each cell is the regression coefficient. The second figure in each cell is the t-statistic. *, **, *** indicate significance at the 10%, 5% and 1% level, respectively. We also control industries fixed-effect, year fixed-effect, and firm fixed-effect in our model.*

The regression coefficient β_3 in column (1) is -0.01, and in column (4) is -0.02. It implies that when stock value decreases, the decline in pay-for-performance sensitivity will be 100% more in the lower compensation group, all else equal. The regression coefficient β_3 in column (2) is -0.16, and in column (5) is -0.80, implying that when net profit decreases, the decrease in pay-for-performance sensitivity is 400% more in the lower compensation group, all else equal. The regression coefficient β_3 in column (3) is -0.31, and in column (6) is -0.86. This implies that when stock value decreases, the decrease in the pay-for-performance sensitivity is 177% more in the lower compensation group, all else equal. Furthermore, all the regression coefficients on ΔPERF × PERDOWN (β_3) in the higher compensation group are

insignificant. In addition, we find that the pay-for-performance sensitivity is more pronounced when firm performance is up for executives who are paid below their peers. The regression coefficient β_1 in column (1) is 0.013, and in column (4) is 0.014 for the performance of value to shareholders; the regression coefficient β_1 in column (2) is 0.17, and in column (5) is 0.72 for the performance of net profit; the regression coefficient β_1 in column (3) is 0.41, and in column (6) is 0.76 for the performance of net profit excluding the extraordinary gains or losses. These results imply that when performance increases, the increase of pay-for-performance sensitivity is 8%, 324%, and 85% more relative to the performance criterion in the lower compensation group, all else equal.

The results are consistent with our expectation that compensation stickiness is larger and more significant and pay-for-performance sensitivity is higher in the lower compensation group. Even more important, firms who paid managers higher compensation in the prior year do not display the characteristic of compensation stickiness. The result suggests that compensation stickiness is an important mechanism to provide a retention incentive, rather than the problem of corporate governance in Chinese firms. Under a retention motive, boards of directors need to prevent executive compensation from falling below the reservation wage. So when executive compensation is already low in the peer group, boards of directors will insulate the CEO from getting penalized for bad performance.

CONCLUDING COMMENTS

In this paper we investigate the relationship between the level of executive compensation and change in firm performance in China's publicly traded firms. Previous research (Firth et al., 2006; Conyon and He, 2011) has examined how firm performance affects executive compensation. We extend this research and analyze a sticky characteristic in executive compensation, and how peer groups affect the sticky characteristic. We use sample of Chinese firms listed on the Shanghai Stock Exchange or the Shenzhen Stock Exchange to test the sticky characteristic, and document a number of important findings. First, we find that the growth of executive compensation is positively correlated to firm performance, regardless of how we measure performance (as market-based performance or accounting-based performance). The result is consistent with agency theory that executive compensation is an effective tool to decrease the agency cost. Second, we find that there is a sticky characteristic in executive compensation. That means top managers are rewarded more for good performance but are punished less for bad performance. Third, we show that peer group is an important determinant of the executive compensation stickiness. The characteristic of compensation stickiness only exists in the firms whose executive compensation is lower than the median compensation in their peer group. The result supports the retention incentive view that boards of directors compare the current level of the executive compensation with the compensation in a peer group of similar companies when they determine executive compensation. If the compensation is lower than the median compensation in the peer group, boards of directors are reluctant to decrease their executive compensation to prevent the executive compensation from falling below the reservation wage and exposing the firm to the risk of losing executives. The paper has a natural limitation in the selection of peer groups. While industry-size-based grouping is a simple way of identifying peer groups, it may not be the most suitable approach. Therefore, to better understand the effect of peer group on CEO compensation, future studies could focus on the development of theories that optimally define firm-specific peer groups. Such an approach would lead to stronger conclusions due to better modeling of this phenomenon.

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INTERNATIONAL EARNINGS TO PRICE RATIO CONVERGENCE: EVIDENCE FROM THE EUROPEAN UNION

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ABSTRACT

This paper investigates whether any pattern of convergence of international earnings-to-price ratios that exist for a sample of 19 European Union (EU) countries over the period 1994-2012 can be detected through the methodology of Phillips and Sul (2007). This methodology is based on a general form of a nonlinear time varying factor model and allows for cross sectional heterogeneity as well as for different transitional time paths towards equilibrium. The results show that such a convergence is not present. Next, the study aims at detecting any potential factors supporting the pattern of divergence. The empirical findings reveal that such divergence patterns mainly reflect divergence in economic factors.

JEL: G10, C23

KEYWORDS: Earnings-Price Ratios; Club Convergence; Clustering Procedure; European Markets

INTRODUCTION

Price-earnings ratios measure the willingness of investors to pay for current earnings. They are used either as a valuation tool for assessing stocks vis-à-vis other stocks in terms of growth and risk or determinants in cost-of-capital computations. Bekaert (1995) argues that higher price-earnings ratios imply a stronger degree of market convergence. These ratios have also been widely used as indicators for the quality of investments (Dontoh *et al.*, 1993) as well as indicators of convergent markets. In the case that expected returns from various markets depend on location, they are characterized as segmented (Karolyi and Stulz, 2002). In addition, the P/E ratio has been considered an indicator of transitory earnings, future earnings or risk. If investors' information sets are not homogeneous, low P/E ratios may signal undervalued stocks and portfolios of low P/E stocks should yield excess returns even after they are adjusted for risk. Even assuming non-homogeneous information sets, the hypothesis that all market agents evaluate the price stocks according to a discounted cash flow approach is still a useful benchmark, which can be tested and rejected in favor of alternative hypotheses. Its implications are that the expected growth of earnings and payout (risk and persistence) should be negatively (positively) related to the earning price ratio.

Benefits of market convergence include lower volatility and an increased ability to absorb shocks. In convergent markets the cost of equity capital diminishes, which in turn boosts investments, and accordingly, economic growth. Colacito and Croce (2010) argue that in convergent markets, shocks are at least partially diversifiable across countries. They conclude that the implied benefits of keeping international financial markets open can be as high as ten percent of lifetime consumption. Similarly, Umutlu et al. (2010) demonstrate that increased market convergence results in lower volatility and an increased ability of global markets to absorb risk. Possible costs, or downside effects, of market convergence include fewer

diversification benefits and increased chance of a joint crash across markets. Financial markets that are at least partially convergent place less weight on local risk factors (Carrieri *et al.*, 2007). This may lead to fewer diversification benefits, which might lead to slower stock price gains (Eun and Lee, 2010). Beine *et al.* (2010) find that financial liberalization increases co-movements between markets most strongly during crashes, implying that increased openness in markets increases the likelihood of a joint crash across all markets.

The objective of this paper is, first, to investigate whether earnings-to-price ratios from 19 European Union (EU) countries have been converging over the recent years, and second, if such convergence or divergence patterns can be documented, it is the role of macroeconomic factors that can explain them. The overall related literature acknowledges that EU positive earnings stocks have a P/E ratio around 25, while the expected 1-year ahead rate of growth is highly volatile. Finally, average betas –around 0.8 for European stocks- show that these stocks are not very sensitive to systematic non diversifiable risk. According to Zarowin (1990) and Bildersee *et al.* (1990), earnings-price ratios differ widely across firms, a fact that can be attributed to differences in certain factors such as economic growth and accounting practices. This paper uses the approach of Bekaert and Harvey (1995) who model time-varying market convergence and assess the relative importance of information arising from the above factors on capital market characteristics.

The contribution of this paper to the above-mentioned framework of convergence is that of directly testing the presence of convergence patterns across a number of EU countries. More specifically, this study contributes to the literature by examining a number of countries for which no relevant study, to the best of our knowledge, occurs relative to the convergence of such financial ratios. The second contribution of this study is the employment of a new methodological approach, that of clustering or clubbing panel convergence testing, recommended by Phillips and Sul (2007). This methodology has several advantages. First, no specific assumptions concerning the stationarity of the variables of interest and/or the existence of common factors are necessary. Second, it is based on a quite general form of nonlinear time varying factor models. More importantly, it takes into account that countries experience transitional dynamics, while it abstains from the hypothesis of homogeneous technological progress, an assumption extensively employed in the majority of growth studies (Phillips and Sul, 2006).

Evidence regarding convergence patterns of P/E ratios is relevant for policymaking. The adoption of common fundamentalist rules is of high concern for investors and traders. A deeper understanding of the role of potentially unexplored variables, which may proxy for the unknown part of the fundamental value of the stock in a framework of asymmetric information, may reveal the presence of significant differences in the way stocks are valued across European financial markets.

The remaining of the paper is organized as follows. Section 2 presents a brief literature review while Section 3 discusses the hypotheses tested. Section 4 reports the description of the data used along with the econometric methodology. Section 5 reports the empirical findings and discusses them, while Section 6 concludes the paper.

LITERATURE REVIEW

The concept of convergence is affected by various factors, such as investment factors and macroeconomic factors. More specifically, investment factors are legal barriers that differentiate between foreign and domestic investors (Bekaert, 1995) or other regulatory framework characteristics (Kim and Singal, 2000) or other type of barriers, such as differences in information, in accounting standards and in investors' protection (Bekaert, 1995). For the case of emerging markets there also present barriers related to specific risks, such as liquidity risks, political risks, economic policy risks and currency risks (Bekaert, 1995), while a group of researchers has examined business cycle or macroeconomic factors affecting market convergence. Such factors involve interest rates and income levels (Calvo *et al.*, 1993; Chuhan *et al.*, 1993).

Aydogan and Górsoy (2000) analyze the P/E effect and the explanatory power of the P/E ratio in 19 emerging market countries. To test the P/E effect, they rank the E/P ratios in descending order, then divide the sample into five quintiles. The authors find that, for all return horizons, a significant P/E effect: average returns decreased as the E/P ratio declined. Additionally, the effect was much stronger for longer (12-month) return horizons. Estrada (2003) proposes a new tool that adjusts the P/E ratio by growth and risk. He compares the performance of value strategies based on the P/E ratio, the P/E ratio adjusted by growth (PEG) and the P/E ratio adjusted by growth and risk (PERG). He finds that the portfolio formed based on PERG significantly outperforms the portfolios based on the P/E ratio and PEG. Chahine and Choudhry (2004) verify whether value outperforms growth strategies across all euro-zone countries. They argue that a high P/E ratio may result two things: first, from a high price relying on expected future cash flows, or second, from a sudden decrease in the earnings level not yet included in the stock price. They find that value stocks (low P/E level) with high earnings growth record the highest performance whereas growth stocks (high P/E ratio) with low earnings growth show the lowest performance. They conclude that the strategy of selling short shares with a price to earnings growth (PEG) ratio higher than one and buying shares with a PEG ratio less than one out performs other strategies.

In another strand of the relevant literature, Campbell and Shiller (1998) show that the ratio of long-run moving average of earnings to the current stock price is negatively associated with future stock returns, while Bekaert *et al*. (2007) show that country-specific growth opportunities are capable of predicting future output and investment growth. By contrast, others find that, at the firm level, the driving force for convergent trends can be primarily explained by the accounting practices employed in the participating capital markets (French and Poterba, 1991; Land and Lang, 2002). The desire for convergence of accounting standards is greater than ever. Convergence aspires to, ultimately, all standard setters agreeing on a single set of high quality accounting standards applied even-handedly. This is expected to greatly reduce uncertainty about comparability of published accounts, while enhancing the transparency of information to the market place. Anderson and Brooks (2006) show that although P/E ratios have been calculated on the basis of the previous year's earnings, the power of the effect has until now been seriously underestimated due to taking too short-term a view of earnings. They use data from all UK firms since 1975 and the traditional P/E ratio and their empirical findings display that the difference in average annual returns between the value and glamour deciles to be 6%. They are able to almost double the value premium by calculating the P/E ratio using earnings averaged over the previous eight years.

Giannetti (2007) studies the predictive ability of the E/P ratio for the S&P 500 Index. He does so by estimating a linear market-timing model, which relates the stock's return to the previous E/P ratio and the latest change in the ratio. The E/P level illustrates the long-term profitability, while the change in the ratio acted as proxy for shorter economic fluctuations. He finds that the earnings price ratio effectiveness as the index returns predictor had spectacularly declined from 1997 to 2002, which may be interpreted in either a rational risk premium or an investor sentiment type framework. Furthermore, he documents strong evidence for market timing ability for unconstrained (no restrictions on short selling and leveraged positions) strategies. Modares et al. (2008) analyze the relationship between different financial ratios and the excess rate of return. For the Teheran Stock Exchange from 2001 to 2004, they find significant results showing that the E/P ratio has a meaningful relationship with the excess rate of return and that it can explain part of the change in the excess rate of return. Azhar et al. (2009) move away from using the traditional P/E ratio to predict future stock returns. They find it problematic to use the traditional P/E ratio because as it is not a symmetric, proportional and scaled measure, and can lead to measurement errors in empirical analysis. Instead, they design a geometric tool and apply the traditional P/E ratio in this framework to arrive at a measure that is free of the problems mentioned above. They estimate a model with data between 1872 and 2008 and conclude that using the geometric version of the P/E ratio is a better measure than the traditional P/E ratio for predicting future returns.

Finally, the literature has offered a number of studies that assess the merit of the P/E ratio as a predictor of stock returns. Although the literature prior to 2000 is extensive, we offer the most recent studies on the issue. In particular, Saleh (2007) examines the performance of the value-growth strategies based on two measures: earnings-to-price and divided-to-price. He uses the CAPM and the Fama-French three-factor model to investigate the ability of these two measures to explain the cross-sectional stock returns. Inconsistent with previous research he provides evidence suggesting that the SMB and HML factors are not significant when he sorted stocks based on E/P and D/P ratios. He attempts to explain such findings by using a multi-factor model, the Fama-French three-factor model augmented with liquidity, leverage, volatility, and winner-loser effects. Jiang and Lee (2009) investigate the prediction of excess returns and fundamentals by financial ratios (e.g. earnings-to-price ratio, book-to-market ratio, and dividend-to-price ratio) by splitting financial ratios into two components: cyclical components and stochastic trend components. They find that the cyclical components predict increases in future stock returns, whilst the stochastic trend components predict declines in future stock returns, while Pietrovite (2009) confirms that managers use the information contained in the P/E ratios to make investment decisions.

Hypotheses Development

The following questions motivate our analysis. First, do earnings-to-price ratios exhibit a time-trend convergence pattern? Recent trends in world financial markets suggest this possibility. The removal of barriers between capital markets in the case of the EU area and movements towards international harmonization of accounting standards (through the adoption of International Accounting Standards-IAS regime) may contribute to market convergence. If the sources of systematic risk are common and are priced internationally, it would follow that earnings-to-price ratios for similar firms would be the same on a global basis because the growth opportunities would be priced in a more homogeneous manner across these markets.

Next, if there is a time trend convergence pattern for earnings to price ratios, are there factors that drive this convergence? One factor of interest is market efficiency, as proxied by variance ratios. This study assesses the relative differences in efficiency at the market level across equity markets. Measuring whether efficiency holds is a crucial matter because pricing drives capital allocation in equity markets. Markets where share prices slowly incorporate information are less efficient than markets where share prices rapidly incorporate all available information. In other words, efficiency testing implies the degree to which equity market returns follow a random walk process. The methodology of testing for efficiency, recommended by Lo and MacKinlay (1988) and Campbell *et al.* (1997), is based on variance ratios that study autocorrelations in such returns. To this end, the association between earnings-price ratios and efficiency is extremely crucial, since it indicates how investors react to new information, especially earnings' information (Kryzanowski and Zhang, 1992; Chan *et al.*, 1996; Bartholdy, 1998).

Ferson and Harvey (1991) demonstrate that predictability may be driven by time-varying risk premiums instead of market inefficiency. Thus, it is not possible to distinguish between high variance ratios due to market inefficiency and those due to time-varying risk premiums. However, very low variance ratios would be evidence that markets are efficient, since time-varying risk premiums would only increase the predictability of returns.

In addition to efficiency and variance ratios, another factor that has been described as driving convergence is the harmonization of accounting practices on a worldwide basis. This harmonization practices imply that accounting variables are being measured similarly and, thus, any such convergence is expected to lead to a similar convergence in the earnings-to-price ratios. Ball *et al*. (2000) argue that there are systematic differences in accounting incomes as well as in accounting systems on an international basis, while Choi and Meek (2005) report that the International Accounting Standard Board has issued IAS that are closely compatible with accounting standards applied to countries, such as the U.S. and the U.K.

DATA AND METHODOLOGY

The data used in this study are quarterly earnings-to-price ratios, consumer prices, proxied by the consumer price index, industrial production, short- and long-term interest rates, proxied by the 3-month T-Bill rate and the 10-year Treasury bond yield, respectively, market capitalization, and dividend yields for 19 EU markets spanning the period 1994:q1-2012:q2. We compute the real interest rate as the nominal rate minus the quarterly consumer price realized inflation. Market capitalization is expressed in terms of national currency. We did not convert market capitalization to the same currency in order to avoid introducing the stochastic properties of the exchange rates into the original series. To be able to compare different market capitalization series, we divided all the observations by the first observation and then multiplied all observations by a factor of 100. In other words, market capitalization is expressed as an indicator.

Although the process of international financial convergence was initiated in 1980 (Gultekin *et al*. 1989; Mittoo, 1992; Karolyi and Stulz, 2002), our empiricasl analysis begins at 1994 due to data availability for all countries under investigation. These 19 markets are selected due to data availability for earnings. Earnings-to-price ratios, market capitalization and dividend yields come from the Datastream's value weighted total market index and, naturally, obtained from Datastream. Daily series of the Datastream's market index are used to construct quarterly measures of market efficiency, described later. Consumer prices and industrial production are collected from the International Financial Statistics database, except for those on consumer price index for Germany obtained from the OECD National Accounts database. Finally, data on short-term interest rates are obtained from the OECD Economic Outlook, while those on long-term interest rates are obtained from Oxford Economics. For Cyprus no interest rate time series data are available, and, thus, the country was not included in the interest rate section of convergence.

Phillips and Sul (2007) propose a new econometric approach for testing the convergence hypothesis. Their method is based on a nonlinear time varying factor model and provides the framework for modeling transitional dynamics as well as long-run behavior. This section describes the econometric methodology for testing convergence for identifying convergence clubs. The new methodology is based on the following time varying common factor representation for y_{it} of economy (country) i:

$$y_{it} = \delta_{it}\mu_t, \tag{1}$$

where μ_t is a single common component and δ_{it} is a time varying idiosyncratic element, which captures the deviation of economy i from the common path which is defined as μ_t. Within this framework, all N economies will converge, at some point in the future, to the steady state if $\lim_{k\to\infty}\delta_{it+k} = \delta$ for all i=1,2,…,N, irrespective of whether they are near the steady state or in transition. This is of major importance since paths to the steady state (or states) across economies can differ significantly. Moreover, the data sets available to empirical researchers most likely contain transition periods as well as periods near the steady state. Since δ_{it} cannot be directly estimated from (1) due to over-parametrization, Phillips and Sul (2007, 2003) eliminate the common component μ_t through the rescaling by the panel average:

$$h_{it} = \frac{y_{it}}{\frac{1}{N}\sum_{j=1}^{N} y_{jt}} = \frac{\delta_{it}}{\frac{1}{N}\sum_{j=1}^{N}\delta_{jt}}, \tag{2}$$

The relative measure h_{it} captures the transition path, but with respect to the panel average. In order to define a formal econometric test of convergence as well as an empirical algorithm of defining club convergence, the authors assume the following semiparametric form for the time varying coefficients δ_{it} **:**

$\delta_{it} = \delta_i + \sigma_{it}\xi_{it}$, where $\sigma_{it} = \frac{\sigma_i}{L(t)t^{\alpha}}$, $\sigma_i > 0$, $t \geq 0$, and ξ_{it} is weakly dependent over t, but iid(0,1) over i**.** The function $L(t)$ is a slowly varying function, increasing and divergent at infinity.

In this paper we set $L(t) = \log(t+1)$ (3)

Under this specific form for δ_{it},

the null hypothesis of convergence for all i, takes the form: $H_0: \quad \delta_i = \delta, \quad \alpha \geq 0$, while the alternative hypothesis of no convergence for some i, is expressed as: $H_A: \quad \delta_i \neq \delta \quad or \quad \alpha < 0$. Phillips and Sul (2007) show that the null of convergence can be tested in the framework of the following regression (the analytic proof that under the convergence hypothesis this regression equation holds is reported in Appendix B of Phillips and Sul, 2007):

$$\log\left(\frac{H_1}{H_t}\right) - 2\log L(t) = \hat{c} + \hat{b}\log t + \hat{u}_t, \tag{4}$$

for $t = [rT], [rT]+1, ..., T$, and $r > 0$**.** Following Phillips and Sul (2007), the methodology selects r in [0, .. ,3]. Then, this r is the value multiplied by T to truncate the data sample.

In this regression, $H_t = \frac{1}{N}\sum_{i=1}^{N}(h_{it} - 1)^2$ (5)

and

$$\hat{b} = 2\hat{\alpha} \tag{6}$$

where h_{it} is defined in (2) and $\hat{\alpha}$ is the least squares estimate of α. Under the null hypothesis of convergence, the dependent variable diverges whether $\alpha > 0$ or $\alpha = 0$, thus the convergence hypothesis can be tested by a t test of the inequality null $\alpha \geq 0$. The t-test statistic $t_{\hat{b}} = \frac{\hat{b}}{s_{\hat{b}}}$ (7)

follows asymptotically the standard normal distribution and it can be constructed using a heteroskedasticity and autocorrelation consistent (HAC) standard error.

Following Phillips and Sul, we estimate standard errors by

$$s_{\hat{b}} = l\hat{\text{var}}_r(\hat{u}_t)\left[\sum_{t=[Tr]}^{T}\left(\log t - \frac{1}{T-[Tr]+1}\sum_{t=[Tr]}^{T}\log t\right)^2\right]^{-1} \tag{8}$$

where $\hat{l}\,\widehat{\text{var}}(\hat{u}_t)$ is the estimated HAC long-run variance of the regression residuals. At the 5% level we reject the null hypothesis of convergence is logt < - 1.65. Phillips and Sul (2007) call the one sided t test, which is based on $t_{\hat{b}}$, the $\log t$ test due to the presence of the logt regressor. The logt test has good asymptotic and finite sample properties and it is easy to be implemented.

One important issue in the empirical convergence literature is the possible existence of multiple equilibria. In such a case, rejection of the null hypothesis that all economies in the sample are under convergence does not imply the absence of different convergence clubs in the panel. In our study we implement the "Club Convergence and Clustering" procedure proposed by Phillips and Sul (2007). The procedure is summarized in the following steps: First, order the N economies according to the value of the final times series. Second, form all possible core groups C_k by selecting the first k highest economies, with $k = 2,3,...N$. Test for convergence using the $\log t_k$ test within each subgroup of size k. Define the core group C^* of size k^* as the group for which the maximum $\log t_{k^*}$ statistic is computed, given of course that all $\log t_k$ statistics over which the maximization is performed support the convergence hypothesis. Third, find all the economies that according to the logt test converge to the same steady state with the core group C^*. This identifies the first convergence club in the panel. For the remaining economies (if any) the procedure is repeated in order to determine the next convergence club, if one exists. We stop when the remaining economies fail to converge, i.e. according to the logt test, they form a club.

RESULTS

Tables 1 and 2 present summary statistics on the mean and standard deviation of inflation, dividend yields, industrial production, long-term interest rates and market capitalization, P/E ratios, stock prices and short-term interest rates, respectively. Means are calculated as the average of the quarterly values for each quarter.

Table 3 reports the results of the panel convergence methodology for the earnings-price ratios. The table involves three columns. The first column reports the results of the full convergence *logt* test, i.e., convergence among all sample countries. As the first column in this table indicates, the null hypothesis of full convergence is rejected at the 5% level. Specifically, the point estimate of the t-value is -1.660. These findings imply the absence of a converging behavior in the earnings-price ratios in all 19 EU countries.

Given the absence of full convergence, we proceed with implementing the club clustering algorithm to identify subgroups of countries that satisfy the convergence criterion (hereafter, the terms group and club are used interchangeably). The results are reported in columns two through three. As these findings indicate, one convergent club is formed by Austria, Cyprus, Czech Republic, Finland, France, Greece, Hungary, Italy, Luxembourg, Poland, Portugal, Spain, Sweden and the UK, while the second group involves Belgium, Germany, Ireland and the Netherlands. Finally, Denmark displays an independent (non-converging) behavior. The lack of full convergence of the earnings-price ratios, and more importantly, the formation of two convergent clubs, calls for the investigation of possible factors responsible for such lack of convergence.

Table 1: Summary Statistics Part 1

Countries	Mean	Min	Max	Standard Deviation	Mean	Min	Max	Standard Deviation
	Panel A: Inflation				**Panel B: Dividend Yields**			
Austria	0.019	0.000	0.037	0.009	2.055	0.930	5.510	0.817
Belgium	0.020	-0.012	0.054	0.011	3.168	1.440	11.420	1.811
Cyprus	0.027	-0.010	0.053	0.011	8.662	0.580	41.560	10.393
Czech Rep.	0.040	-0.004	0.126	0.032	3.921	0.630	9.110	2.155
Denmark	0.021	0.009	0.041	0.006	1.553	0.900	3.090	0.407
Finland	0.016	-0.010	0.045	0.012	2.946	0.690	7.440	1.378
`France	0.016	- 0.005	0.033	0.007	2.793	1.280	5.810	1.030
Germany	0.040	0.006	0.096	0.019	2.588	1.140	6.460	1.096
Greece	0.090	0.024	0.265	0.063	2.739	1.260	6.410	1.060
Hungary	0.025	-0.063	0.064	0.025	2.524	1.060	7.390	1.279
Ireland	0.024	0.001	0.055	0.010	3.445	1.480	10.090	1.375
Italy	0.021	-0.002	0.042	0.009	2.295	1.150	4.040	0.685
Luxembourg	0.021	-0.001	0.042	0.009	2.827	1.330	7.860	1.127
Netherlands	0.021	0.003	0.043	0.008	2.446	0.480	6.070	1.201
Poland	0.068	0.003	0.286	0.066	2.798	0.320	8.270	1.544
Portugal	0.027	-0.015	0.048	0.012	3.066	1.050	6.959	1.414
Spain	0.028	-0.011	0.051	0.012	2.884	1.240	6.090	0.950
Sweden	0.013	-0.014	0.042	0.012	2.901	1.400	5.230	0.811
UK	0.021	0.005	0.048	0.010	3.611	2.720	4.090	0.421
	Panel C: Industrial Production				**Panel D: Long Term Interest Rates**			
Austria	90.600	57.600	123.800	18.985	4.651	1.840	7.680	1.273
Belgium	92.001	61.300	125.200	17.222	4.838	2.900	8.340	1.248
Cyprus	92.919	77.500	111.100	8.501	------	-----	-----	------
Czech Rep.	93.219	64.900	129.000	20.149	6.733	2.560	19.670	3.953
Denmark	94.774	76.200	111.600	8.639	4.762	1.200	8.760	1.647
Finland	90.814	57.400	118.500	16.075	4.861	1.850	10.350	1.844
France	95.364	84.500	102.600	5.401	4.665	2.200	8.260	1.345
Germany	98.785	80.600	124.300	11.594	4.447	1.400	7.570	1.375
Greece	90.980	71.541	106.361	9.038	8.051	3.410	31.740	5.689
Hungary	83.288	39.200	128.900	26.284	9.324	6.180	14.450	2.777
Ireland	81.171	29.700	112.900	26.912	5.597	1.970	10.800	1.791
Italy	97.732	77.000	109.800	8.493	5.748	3.390	12.740	2.398
Luxembourg	85.225	61.900	105.700	12.476	4.486	1.770	7.850	1.478
Netherlands	96.195	73.100	120.200	10.915	4.548	1.750	7.560	1.274
Poland	91.862	46.000	151.900	29.463	10.689	4.740	28.450	6.620
Portugal	99.479	78.100	113.000	8.369	6.147	3.330	13.470	2.727
Spain	91.867	75.400	106.800	8.825	5.538	3.180	11.910	2.182
Sweden	91.829	65.400	116.200	11.949	5.045	1.380	11.110	2.260
UK	97.714	88.000	103.600	4.466	5.417	3.960	7.680	1.200

Notes: The table reports basic statistics, i.e. mean, min, max and standard deviation for a number of macroeconomic and financial variables, i.e. inflation, dividend yields, industrial production, and long-term interest rates.

This part of the paper attempts to examine the convergence or divergence pattern of certain variables influencing the behavior of the earnings-price ratios. According to Kumar and Hyodo (2001), certain factors can be used to explain price-earnings ratios patterns, which include economic factors related to the macroeconomic environment, i.e. inflation, industrial production and interest rates, differences in accounting rules and differences in market efficiency. One of those economic variables is the market capitalization (Table 4), which is used as a proxy for the size of a market. The reason is that smaller markets are considered to offer greater opportunities for growth. By contrast, the small size of such markets may be an indication of their inability to grow or their possession of high risk.

Table 4 reports the results of the panel convergence methodology for market capitalization. The table involves four columns. As the first column indicates, the null hypothesis of full convergence is rejected at any reasonable significance level. Specifically, the point estimate of the t-value is -19.773. These findings imply the absence of a converging behavior in the earnings-price ratios in all 19 EU countries. The results also report that three distinctive clubs are formed, with the results providing empirical support to those

displayed in Table 3 and indicating the absence of full convergence. In other words, the results imply that the market size is a factor giving rise to a divergent behavior.

Table 2: Summary Statistics Part 2

Countries	Mean	Min	Max	Standard Deviation	Mean	Min	Max	Standard Deviation
	Panel A: Market Capitalization				**Panel B: PE Ratio**			
Austria	5,5567	15,145	155,556	39,868	16.212	6.000	32.100	4.896
Belgium	149,455	45,641	268,368	76,600	14.293	5.200	25.100	3.985
Cyprus	5,567	826	19,482	4,680	13.941	0.600	39.500	10.368
Czech Rep.	601,138	127,724	1,397,984	316,440	15.430	7.600	32.000	5.368
Denmark	701,654	1,99,132	1,308,502	311,221	18.624	8.900	33.000	4.578
Finland	133,406	18,605	359,082	79,455	17.401	6.600	71.400	11.509
France	1,008,026	253,360	1,928,818	454,231	15.546	8.000	25.400	3.629
Germany	834,159	288,704	1,384,551	300,134	16.031	9.100	26.800	4.063
Greece	61,545	4,914	157,963	39,939	16.816	7.300	36.300	6.395
Hungary	3,851,204	74,020	8,932,417	2,422,041	15.038	7.100	29.100	4.350
Ireland	56,778.3	13,127	123,325	27,969	14.674	4.000	26.200	4.361
Italy	459,087	104,243	803,866	207,311	17.523	6.400	32.500	5.290
Luxembourg	19,810	4,638	37,183	9,336	17.091	4.300	32.900	6.875
Netherlands	425,507	166,575	754,968	491,951	16.605	5.400	31.900	5.620
Poland	176,044	5,594	491,950	151,451	15.503	6.600	27.900	5.464
Portugal	51,439	9,248	99,567	21,741	17.746	4.400	40.200	6.607
Spain	373,586	82,474	746,106	184,162	15.565	7.200	25.600	4.655
Sweden	2,163,947	552,425	3,786,593	922,590	17.093	7.500	31.200	5.467
UK	1,379,103	549,087	2,003,365	420,123	15.884	7.100	26.700	3.988
	Panel C: Stock Price Index				**Panel D: Short Term Interest Rates**			
Austria	1,339.158	600.940	3,292.150	776.838	3.066	0.660	5.830	1.347
Belgium	2,626.552	935.030	5,079.450	1,089.817	3.143	0.660	6.810	1.455
Cyprus	582.094	82.71	2,344.090	473.981	----	----	----	----
Czech Rep.	325.611	74.460	828.590	253.569	5.598	0.980	19.670	4.678
Denmark	5,173.997	1,444.970	10,614.110	2,606.701	3.551	0.380	6.860	1.601
Finland	737.916	122,520	1,666.090	406.262	3.266	0.660	7.580	1.609
France	5,612.773	1,847.870	10,315.700	2,326.292	3.155	0.660	6.050	1.456
Germany	1,383.937	580.560	2,323.540	483.084	3.067	0.660	5.830	1.347
Greece	1,153.445	268.700	2,555.86	651.238	6.103	0.660	20.000	5.322
Hungary	1,120.097	122.280	2,633.13	699.931	12.366	5.290	23.450	6.256
Ireland	8,180.751	2,486.510	17,781.750	3,786.668	4.108	0.660	11.010	2.787
Italy	4,529.861	1,772.150	8,205.560	1,726.834	3.038	0.66	5.280	1.328
Luxembourg	512.284	192.500	960.080	215.878	3.104	0.437	6.520	1.421
Netherlands	3,737.393	1,283.640	6,400.900	1,332.423	2.997	0.385	5.283	1.362
Poland	125.213	28.130	284.420	65.179	4.043	0.660	13.630	2.867
Portugal	257.083	94.010	507.350	97.454	3.780	0.660	9.320	2.306
Spain	622.820	158.400	1,212.86	286.957	3.823	0.480	9.170	2.061
Sweden	3,415.055	878.070	6,563.210	1,684.232	4.563	0.590	7.580	2.069
UK	19,908.790	8,117.490	32,136.890	6,803.258	4.833	0.660	7.000	1.823

Notes: The table reports basic statistics, i.e. mean, min, max and standard deviation for a number of financial variables, i.e. market capitalization, the P/E ratio, stock prices, and short-term interest rates.

Table 3. Club Convergence – Earnings-Price Ratios

All Countries	First Convergence Club	Second Convergence Club	Not Converging Countries
Austria	Austria	Belgium	Denmark
Belgium	Cyprus	Germany	
Cyprus	Czech	Ireland	
Czech Rep.	Finland	Netherlands	
Denmark	France		
Finland	Greece		
France	Hungary		
Germany	Italy		
Greece	Luxembourg		
Hungary	Poland		
Ireland	Portugal		
Italy	Spain		
Luxembourg	Sweden		
Netherlands	UK		
Poland			
Portugal			
Spain			
Sweden			
UK			
logt = -1.660*	logt = -0.527***	logt = 7.601***	

*Notes: The first column reports the results of the full convergence logt test, i.e., convergence across all sample countries. The results from implementing the club clustering algorithm are reported in columns two and three, while the last column reports the non-diverging country group. ***, ** and * indicate significance at the 1, 5 and 10 percent levels respectively.*

Table 4. Club Convergence – Market Capitalization

All Countries	First Convergence Club	Second Convergence Club	Third Convergence Club	Not Converging Countries
Austria	Austria	Belgium	Germany	Netherlands
Belgium	Hungary	Cyprus	UK	
Cyprus	Poland	Czech Rep.		
Czech Rep.		Denmark		
Denmark		Finland		
Finland		France		
France		Greece		
Germany		Ireland		
Greece		Italy		
Hungary		Luxembourg		
Ireland		Portugal		
Italy		Spain		
Luxembourg		Sweden		
Netherlands				
Poland				
Portugal				
Spain				
Sweden				
UK				
logt = -19.733	logt = 2.123***	logt = 2.519***	logt = 2.134***	

*Notes: The first column reports the results of the full convergence logt test, i.e., convergence across all sample countries. The results from implementing the club clustering algorithm are reported in columns two to four. The last column reports the non-converging club. *** indicates significance at the 1 percent level.*

Next, we repeat the analysis by separating our full sample into a small markets group and a large markets group in terms of market capitalization. The median of such capitalization is the criterion for dividing all capital markets under study as small or large capital markets. The results are reported in Table 5.

The null hypothesis of full convergence is rejected at the 5% level in both sizes. For large capitalization countries (Panel A), the countries under examination belong exactly to two convergence clubs. In particular, the point estimate of *logt* t-statistic over the relevant period is -51.75, which rejects the full convergence hypothesis. The first club includes Spain and Sweden and the second club includes Belgium, France,

Germany, Italy and the UK, while the Netherlands displays a non-converging behavior. In terms of the small capitalization countries (Panel B), the results also point out to an overall diverging pattern, since the *logt* t-statistic is -10.681, which strongly rejects the convergence hypothesis. The analysis in this group reveals the presence of two clubs. More specifically, the first club includes Austria, Hungary and Poland, while the second club includes Cyprus, Czech Republic, Greece, Ireland, Luxemburg and Portugal. The results imply that the market size is definitely a factor giving rise to a divergent behavior.

Table 5. Club Convergence – Market Capitalization: Large vs Small Capitalization Countries

Panel A: Large Capitalization Countries				**Panel B: Small Capitalization Countries**			
Countries	**First Convergence Club**	**Second Convergence Club**	**Not Converging Countries**	**Countries**	**First Convergence Club**	**Second Convergence Club**	**Not Converging Countries**
Belgium	Spain	Belgium	Netherlands	Austria	Austria	Cyprus	
France	Sweden	France		Cyprus	Hungary	Czech rep.	
Germany		Germany		Czech Rep.	Poland	Greece	
Italy		Italy		Denmark		Ireland	
Netherlands		UK		Finland		Luxembourg	
Spain				Greece		Portugal	
Sweden				Hungary			
UK				Ireland			
				Luxembourg			
				Poland			
				Portugal			
logt = -51.75	logt = 1.08***	logt = 12.62***		logt = -10.681	logt = 1.31***	logt = 9.64***	

*Notes: Panel A reports the market capitalization convergence results across the large capitalized countries, while Panel B repeats the analysis for the small capitalized countries. The first column of each panel reports the results of the full convergence logt test, i.e., convergence across all sample countries, while the results from implementing the club clustering algorithm are reported in columns two to three in each panel. Finally, column four in each panel reports the non-converging club. *** indicates significance at the 1 percent levels.*

Next, three macroeconomic variables are examined: inflation, industrial production, and the real interest rate. According to Lamont (1998), earnings contain information about returns that is not contained in other variables. Because earnings are correlated with business conditions, earnings have predictive power for returns. Guenther and Young (2000) make a similar argument, claiming that earnings are associated with real economic activity. First, we examine inflation, a factor attributed to the economic environment of the firm. More specifically, Modigliani and Cohn (1979) and French and Poterba (1991) argue that inflation is a source of difference between accounting and economic earnings. The smaller the inflation differences between markets, the smaller their differences in corresponding earnings-price ratios. The results are reported in Panel A of Table 6. As the findings in Panel A of Table 6 indicate, there is a full convergence club for which the point estimate of *logt* t-statistic is 2.419. Only the Netherlands displays a non-converging country behavior, which does not alter the convergence results. These empirical findings are in accordance with the expectations that the formation of the EU (involving also the countries participating in the European Monetary Union area) has acted as a converging propelled mechanism for inflation convergence. Next, we report (Panel B) the industrial production convergence results. Industrial production is used as proxy for real economic activity. For the empirical purposes of this study we make use of the growth rate of industrial production. The findings show that the null hypothesis of full convergence is rejected, since the value of *logt* t-statistic is -90.633. The results of the club clustering procedure identify five converging clubs. These findings document the diverging behavior of the European Southern countries in terms of their real economic activity. In other words, the results point out the different economic structure of EU economies, an argument supporting the view of the asymmetric behavior of EU economies and the recent crisis events.

Table 6. Club Convergence – Inflation and Industrial Production

All Countries	**First Convergence Club**	**Not Converging Countries**
Panel A: Inflation		
Austria	-------	Netherlands
Belgium		
Cyprus		
Czech Rep.		
Denmark		
Finland		
France		
Germany		
Greece		
Hungary		
Ireland		
Italy		
Luxembourg		
Netherlands		
Poland		
Portugal		
Spain		
Sweden		
UK		
logt = 2.419***		

Panel B: Industrial Production					
All Countries	**First Convergence Club**	**Second Convergence Club**	**Third Convergence Club**	**Fourth Convergence Club**	**Fifth Convergence Club**
Austria	Hungary	Austria	Cyprus	France	Greece
Belgium	Ireland	Belgium	Denmark	Portugal	Italy
Cyprus	Poland	Czech Rep.	Germany	Spain	UK
Czech Republic		Finland	Luxembourg		
Denmark			Netherlands		
Finland			Sweden		
France					
Germany					
Greece					
Hungary					
Ireland					
Italy					
Luxembourg					
Netherlands					
Poland					
Portugal					
Spain					
Sweden					
UK					
logt=-90.63	logt=5.764***	logt=7.970***	logt=4.394***	logt=6.38***	logt=0.923***

*Panel A in this table reports the results for inflation. The first column reports the results of the full convergence logt test, i.e., convergence across all sample countries, while column 3 reports the non-converging club results. Panel B reports the results for the definition of industrial production. The first column reports the results of the full convergence logt test, i.e., convergence across all sample countries, while the results from implementing the club clustering algorithm are reported in columns 2 to 6. *** indicates significance at the 1 per cent.*

The final factor capturing the impact of the macroeconomic environment is real interest rates. A number of studies document a significant diverging behavior in real interest rates across countries over time (Mishkin, 1984; Darby, 1986; Chowdhry and Titman, 2001). Frankel (1991) and Chowdhry and Titman (1991) have discussed the potential association between real interest rates and earnings-price ratios across countries. The principal mechanism that governs such an association is the impact of real interest rates on the productivity of traded goods. In particular, changes in productivity lead to changes in the cost of producing capital goods and, thus, to changes in their prices. The latter affect investors' behavior and consequently the rental rate of such capital. However, the price to rental ratio is fundamentally the same as the price-earnings ratio, and thus, we can document the association between real interest rates and price-earnings ratios. Moreover, industries involved in the production of capital goods usually make use of either labor-intensive production patterns or capital-intensive production technology, implying an additional

mechanism that explains the above association. Table 7 reports the results of the formal convergence test obtained from the clustering algorithm for the interest rates. Full convergence is strongly rejected for both short- (Panel A) and long-term (Panel B) interest rates, since the logt statistic takes the values of -3.178 and -5.125, respectively. Moreover, the application of the club convergence procedure on short-term interest rates indicates the presence of four convergence clubs, while two countries in our sample, namely, Denmark and Hungary, fail to converge to any of these clubs, i.e. they form a distinctive club. In terms of long-run interest rates, the results indicate the presence of three convergence clubs.

Table 7. Club Convergence – Real Interest Rates

All Countries	**First Convergence Club**	**Second Convergence Club**	**Third Convergence Club**	**Fourth Convergence Club**	**Not Converging Countries club**
Panel A: Short-term Real Interest Rates					
Austria	Greece	Czech Rep.	Ireland	Austria	Denmark
Belgium	Sweden	Italy	Poland	Belgium	Hungary
Czech Rep.		Portugal		Finland	
Denmark		Spain		France	
Finland		UK		Germany	
France				Luxembourg	
Germany				Netherlands	
Greece					
Hungary					
Ireland					
Italy					
Luxembourg					
Netherlands					
Poland					
Portugal					
Spain					
Sweden					
UK					
logt = -3.178	logt = -1.539**	logt = 1.536***	logt = -0.933***	logt = 1.529***	logt = -6.779

Panel B: Long-Term Real Interest Rates			
All countries	**First Convergence Club**	**Second Convergence Club**	**Third Convergence Club**
Austria	Greece	Hungary	Austria
Belgium	Poland	Ireland	Belgium
Cyprus	Portugal		Czech Republic
Czech Republic			Denmark
Denmark			Finland
Finland			France
France			Germany
Germany			Italy
Greece			Luxembourg
Hungary			Netherlands
Ireland			Sweden
Italy			UK
Luxembourg			
Netherlands			
Poland			
Portugal			
Spain			
Sweden			
UK			
Logt = -5.125	Logt = 0.647***	Logt = 1.663***	Logt = -1.598**

Notes: Panel A in this table reports the results for the short-term real interest rates. The first column reports the results of the full convergence logt test, i.e., convergence across all sample countries. The results from implementing the club clustering algorithm are reported in columns 2 to 5, while column 6 reports the non-converging club results. Panel B reports the results for the long-term real interest rates. The first column reports the results of the full convergence *logt* test, i.e., convergence across all sample countries. The results from implementing the club clustering algorithm are reported in columns 2 to 4. *** and ** indicate significance at the 1 and 5 per cent, respectively.

Finally, we turn to dividend yields, which we use as a proxy for the convergence of accounting standards. These yields are closely linked to the firm's cost of capital. In terms of convergence, changes in dividend yields should be a reliable measure of changes in the extent of convergence, while lower dividend yields

imply stronger convergent markets. Bekaert and Harvey (2000) measure the impact of liberalization on financial market convergence by examining the role of the cost of equity capital. According to them, dividend yields capture the permanent price effects following changes in the cost of capital much better than noisy returns can. As a result, dividend yields decline after liberalizations. Land and Lang (2002) claim that dividends are the cash resources paid out to stockholders. If earnings change, say, due to changes in accounting practices, then changes in earnings are very unlikely to be associated with changes in dividends (Aron, 1991; French and Poterba, 1991). While stock prices are computed as discounted future flows of dividends, it is highly likely that dividend yields could change along with earnings-price ratios due to changes in accounting practices. By contrast, La Porta *et al.* (2000) document dividends are low in countries where legal systems do not strongly protect minority shareholders from inside expropriation. Given the high heterogeneity of such relevant institutional factors around the globe, dividends are not expected to act as potential proxies for financial, i.e., earnings-price ratios, convergence. Only by opting into more protective legal regimes or radically changing the legal environment could higher convergence processes be generated.

Results from the formal econometric convergence tests are reported in Table 8 (Panel A), which show that the null hypothesis of full convergence is accepted for the time period under consideration, since the value of *logt* t-statistic is 14.018. These findings document a strong convergence pattern in accounting practices.

Table 8. Club Convergence – Dividend Yields and Market Efficiency

Panel A: Dividend Yields			**Panel B: Market Efficiency**		
All Countries	**First Convergence Club**	**Not Converging Countries**	**All Countries**	**First Convergence Club**	**Not Converging Countries**
Austria	---------	----------	Austria	Austria	Czech Rep.
Belgium			Belgium	Belgium	Portugal
Cyprus			Cyprus	Cyprus	
Czech Rep.			Czech Rep.	Denmark	
Denmark			Denmark	Finland	
Finland			Finland	France	
France			France	Germany	
Germany			Germany	Greece	
Greece			Greece	Hungary	
Hungary			Hungary	Ireland	
Ireland			Ireland	Italy	
Italy			Italy	Luxembourg	
Luxembourg			Luxembourg	Netherlands	
Netherlands			Netherlands	Poland	
Poland			Poland	Spain	
Portugal			Portugal	Sweden	
Spain			Spain	UK	
Sweden			Sweden		
UK			UK		
logt = 14.018***			logt = -20.219	logt = 5.818***	logt = -7.454

*Notes: Panel A in this table reports the results for dividend yields. The first column reports the results of the full convergence logt test, i.e., convergence across all sample countries, while Panel B reports the results for the definition of market efficiency. The first column reports the results of the full convergence logt test, i.e., convergence across all sample countries, while the results from implementing the club clustering algorithm are reported in column 2. Finally, column 3 reports the non-converging club results. *** indicates significance at the 1 per cent.*

In addition, the final factor examined in this paper is market efficiency in its weak form. In particular, if prices follow a random walk (RW) pattern, then returns turn out to be described as IID processes and the variance of q-period returns equals q times the variance of the one-period returns. Following Lo and MacKinlay (1988) and Campbell *et al.* (1997), we define a measure of relative efficiency based on variance ratios (VRs). Under the RW hypothesis, VRs should be equal to one. By contrast, in an inefficient market VRs ratios are above unity if returns are positively correlated and below unity in the case returns exhibit negative autocorrelations. We call our measure of efficiency the absolute variance ratio (AVR), which is defined as:

$$AVR_t = \left| \frac{\widetilde{\sigma}_t(q)}{\widetilde{\sigma}_t(1)} - 1 \right|,$$

where the index t refers to the quarter t and $\widetilde{\sigma}_t(q)$ is the sample variance of the q-period holding returns calculated using daily returns in the quarter t. In the present paper we set q=5. We also perform the analysis for q=2, 3, and 4. The results are qualitatively similar to those presented here and are available from the authors upon request. Table 8 (Panel B) reports that the logt statistic takes the value of -20.219, which rejects the null hypothesis of full convergence. However, we do have the presence of a single club, while two countries, i.e. the Czech Republic and Portugal form a non-converging group. In other words, we can potentially claim that the EU countries form virtually a single club in terms of their market efficiency issue.

CONCLUDING COMMENTS

This paper tested for earnings-price ratios convergence among 19 EU countries over the period 1994:q1-2012:q2 and attempted to explain the contribution of certain economic and accounting factors, such as inflation, market capitalization, dividend yields, stock prices, short- and long-term interest rates, industrial production and a measure of market efficiency in the countries' converging or non-converging behavior. To serve both objectives, the novel methodology of Phillips and Sul (2007) has been used. This methodology used a non-linear factor model with a common and an idiosyncratic component – both time-varying, which allows for technical progress heterogeneity across countries.

The empirical findings suggested that the countries in our sample did not form a homogeneous convergence club in terms of their earnings-price ratios, but their convergence patterns varied across markets, a fact that was mainly attributed to inter-country differences in economic factors affecting the behavior of earnings-price ratios. In particular, EU countries were characterized by different economic conditions, such as growth opportunities or business cycles synchronizations.

The implications of our findings are that for a country or a group of countries which do not follow the convergence pattern, more coordinated growth strategies are required to facilitate a stronger capital market convergence process. Furthermore, the implications seem crucial for policy makers as well once they are concerned with the impact of the convergence issue on capital flows and the volatility of financial aggregates. Finally, potential avenues for further research will be the consideration of a sample of countries that includes more developed as well as emerging capital markets.

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INVESTOR REACTION IN STOCK MARKET CRASHES AND POST-CRASH MARKET REVERSALS

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ABSTRACT

We study investor overreaction using data for five major stock market crashes during the 1987-2008 period. We find some evidence of investor overreaction in all five stock market crashes. The prices of stocks investors bid down more than the average during crashes tend to increase more than the average in post-crash market reversals. In line with CAPM, we find that high beta stocks lose more value in crashes and gain more value in post-crash market reversals relative to low beta stocks. We further find that smaller firms and those with a low market-to-book ratio lose more value in stock market crashes. However, they do not gain more value in post-crash market reversals, implying that investor reaction against these firms in stock market crashes is not an overreaction. In examining industry-specific behavior, our results indicate that investors overbid down the prices of high-tech stocks in the 1997 crash and manufacturing stocks in the 2008 crash relative to other stocks. However, the prices of stocks in these industries increased more than other stocks in the post-crash market reversals, implying investor overreaction for these industries in these stock market crashes.

JEL: G00, G01, G10, G14

KEYWORDS: Stock Market Crash, Post-Crash Market Reversal, Determinants of Stock Returns, Investor Overreaction

INTRODUCTION

In this paper, we study investor overreaction using data for five major stock market crashes during the 1987-2008 period. A stock market crash is commonly defined as a sudden dramatic decline of stock prices across a significant cross-section of a stock market. There is no generally accepted threshold for duration or magnitude for the decline in stock prices. Wang et al. (2009) and Gulko (2002) define a stock market crash as 5% or greater decrease in stock prices in a single trading day. In this paper, we study the stock market crashes with a minimum of 9.8% cumulative decline in stock prices in several consecutive trading days.

Stock market crashes are generally followed by several days of sharp market reversal. If there is an overreaction towards stocks with certain financial characteristics during a crash, the reaction is reversed with a sharp market correction during the post-crash market reversal period. For instance, Wang et al. (2013) find that investors overreacted to the technical insolvency risk and bankruptcy risk characteristics of firms by bidding down their stock prices sharply in the 2008 crash. These stocks gained more value relative to other stocks in the post-crash market reversal. In this study, following the methodology used in Wang et al. (2013), we compare the crash and post-crash market reversal periods to determine if there was any investor overreaction in the five most important stock market crashes of the 1987-2008 period. The crash and post-crash market reversal periods included in the study are presented in Table 1.

Table 1: Crash and Recovery Event Characteristics

Event	Crash dates	Crash return (%)	Recovery dates	Recovery return (%)
October 1987	Oct 14 – Oct 19	-28.51	Oct 20 – Oct 21	+14.92
October 1997	Oct 22 – Oct 27	-9.80	Oct 28	+5.12
August 1998	Aug 26 – Aug 31	-12.41	Sep 1	+3.86
April 2000	Apr 10 – Apr 14	-10.54	Apr 17 – Apr 18	+6.27
October 2008	Oct 1 – Oct 10	-22.90	Oct 13	+11.58

This table details the five crash and recovery events that are the subject of this study. Included are the dates for crash and recovery, as well as the S&P 500 index returns for each.

Our research makes several important contributions to the literature. We document a consistent pattern of investor overreaction in a large cross-sectional sample across five of the most significant stock market crashes of the past three decades. We also find that different stock characteristics had varying impact on the magnitude of overreaction among the events included in our study. The paper is organized as follows: The next section examines prior related literature. We then provide information about our data and methodology, and follow with a presentation of our results on stock market crashes during the crash and recovery periods. In the final section, we conclude the paper and note suggestions for future research.

LITERATURE REVIEW

Stock market crashes have received considerable attention in finance literature. Arshanapalli and Doukas (1993) and Lau and McInish (1993) study the co-movements of the world's stock markets before and after the 1987 stock market crash. Roll (1988) and Pan et al. (2001) study the effects of emerging markets in stock market crashes. Wang et al. (2009 and 2010) study the determinants of stock returns in stock market crashes. De Bondt and Thaler (1985, 1987) argue that investors tend to overreact to economic events. Chopra et al. (1992), Rozeff and Zaman (1998), Bauman et al. (1999) and others provide empirical evidence for investor overreaction. In a recent article, Wang et al (2013) demonstrate that investors overreacted to the bankruptcy risk and technical insolvency risk characteristics of firms in the 2008 stock market crash. There are many possible explanations for this 'investor overreaction' phenomenon, from behavioral sentiment issues (Baker, Wurgler, 2006; Barberis, et al., 1998), to herding (Puckett, Yan, 2008) to market microstructure constraints (Park, 1995; Kaul, Nimalendran, 1990; Atkins, Dyl, 1990), to appropriate response to changing risk (Brown et al., 1993).

Dreman and Lufkin (2000) conclude that investor overreaction is psychological. Amini et al. (2013) present an overview of the literature on price reversals. Analysis of price reversals accompanying issue-specific public news or lack thereof on a shorter time frame tends to find evidence of overreaction (e.g. Chan, 2003; Larson, 2003; Bremer, Sweeney, 1991). Chopra et al. (1992) examine a longer time frame and also conclude there is evidence of overreaction. While some researchers attempt to measure investor sentiment directly (Baker, Wurgler, 2006), in this study we follow most prior research and focus on the price movements and company financials to investigate the issue of investor overreaction in general stock market crashes with data for the five most important stock market crashes during the 1987-2008 period.

DATA AND METHODOLOGY

The daily stock trading prices, used in the calculation of daily returns, are obtained from the Center for Research in Security Prices (CRSP) database. The 'crash return' is defined as the cumulative return over several consecutive daily price decreases in the S&P 500 index during the crash event. The 'recovery return' is defined as the cumulative return over several consecutive daily price increases in the S&P 500 index immediately following the crash event. We use the event study methodology and calculate the

cumulative stock returns during the crash and recovery windows using trading price data. We compute the CAPM betas of the stocks using the daily stock returns for the past 90 calendar days and the CRSP-provided returns on a value-weighted index which includes NYSE, NASDAQ, and ARCA securities. Firms with missing trading prices on key event dates and those with fewer than 30 trading quotes in the past 90 calendar days are excluded from the sample. Following Wang et al. (2009), we also exclude firms with a trading price of less than one dollar. We use the Research Insight (COMPUSTAT) quarterly database to collect balance-sheet information on the individual securities. For each security and each event, we select the latest available COMPUSTAT quarterly observation within the year prior to the start of the event. Firms with missing data are excluded from the sample.

The study by Wang et al. (2013) finds that investors overreacted to the technical insolvency risk and bankruptcy risk characteristics of firms in the 2008 stock market crash. The current ratio measures the ability of a firm to meet its maturing obligations and is a standard measure of technical insolvency risk (see: Wang et al, 2013). The debt ratio is commonly used in empirical studies as a measure of firm bankruptcy risk (see, e.g., Mitton, 2002; Baek et al., 2004; Wang et al., 2013). The current ratio and the debt ratio are the two key financial ratios used in Ohlson's (1980) bankruptcy prediction model. We use these two financial ratios in our empirical tests to study investor reaction to technical insolvency risk and bankruptcy risk characteristics of firms in the five stock market crashes and post-crash market reversals included in the paper. Sharpe's CAPM has been generally tested with data for normal time periods. Beta has not been studied sufficiently as a determinant of stock returns in stock market crashes and post-crash market reversals. We use beta as a control variable in our empirical tests and we study if it was a significant determinant of stock returns in the stock market crashes and post-crash market reversals included in the paper. In their three-factor CAPM, in addition to beta, Fama and French (1992, 1993) use firm size and the market-to-book ratio as determinants of stock returns. In our empirical tests, we also use these two variables as controls and investigate if they were significant determinants of stock returns in the stock market crashes and post-crash market reversals included in the study.

Industry dummy variables are commonly used in cross-sectional studies of stock returns (see, e.g., Mitton, 2002; Baek at al., 2004; Wang et al., 2009). To control for the industry effect, we construct five broad industry portfolios (French, 2008) based on SIC codes. The portfolios are 'cnsmr', including consumer durables, nondurables, wholesale, retail, and some services; 'manuf', including manufacturing, energy, and utilities; 'hitec', including business equipment, telephone and television transmission; 'hlth', including healthcare, medical equipment, and drugs; and 'allother', which includes mines, construction, building materials, transportation, hotels, business services, entertainment, and finance. Utility firms' financial decisions are affected by regulation, while financial firm financial ratios are not comparable to those of other firms. Therefore, following Fama and French (2001, 2002), Gadarowski et al. (2007), and Wang (2009 and 2013), we exclude utilities (SIC code 4900-4999) and financial firms (SIC code 6000-6999) from our sample. The data items used in the study from the CRSP and COMPUSTAT databases are presented in Table 2. We list the variables constructed with the data in Table 3. After excluding observations with missing values, we winsorize extreme values using robust median-based measures of center and scale. At the end, we have 2591 observations for 1987, 4642 observations for 1997, 4443 observations for 1998, 4442 observations for 2000, and 3277 observations for 2008, a total of 19395 observations for the entire sample.

The descriptive statistics for the sample are presented in Table 4. The statistics in the table show a pattern of growing firm size over time, both in terms of total assets and market cap. Mean total assets gradually increases from 968 million in 1987 to 5,184 million in 2008, which is expected given the general growth of the economy as well as dollar inflation over the time period. The mean current ratio steadily grows from 2.8 in 1987 to 3.06 in 2000, then drops back down to 2.77 in 2008. The mean debt-to-equity ratio starts out high at 1.34 in 1987, decreases to 1.25 in 1997, then gradually increases to 1.29 by 2008. The crash and recovery returns are highly variable within each event sample showing that, even during

significant overall market moves, there is wide variation in the performance of individual stocks.

Table 2: Data Items Used in the Study

Variable	Source	Description
Indtret	CRSP	Combined index return, including NYSE, Nasdaq, ARCA
Adjtprc	CRSP	Adjusted trading price, present only if trades occurred during the day
Tprc	CRSP	Trading price, present only if trades occurred during the day
SIC	CRSP	SIC code
TCap	CRSP	Market cap, in thousands (rescaled to millions for convenience)
ACTQ	COMPUSTAT	Total current assets, in millions
LCTQ	COMPUSTAT	Total current liabilities, in millions
LTQ	COMPUSTAT	Total liabilities, in millions
ATQ	COMPUSTAT	Total assets, in millions

This table lists all the data items used in this study, along with their descriptions and source databases.

Table 3: Constructed Variables Used in the Study

Variable	Description
crash.return	Simple total return for the 'crash' period for each event.
recovery.return	Simple total return for the 'recovery' period for each event.
beta	Stock beta calculated over the past 90 calendar days of daily stock data.
mkbk	Market to book equity ratio. TCap / (ATQ – LTQ)
dr	Debt to equity ratio. LTQ / (ATQ – LTQ)
cr	Current ratio. ACTQ / LCTQ
cnsmr	Dummy variable set to 1 for the 'consumer' industry portfolio.
manuf	Dummy variable set to 1 for the 'manufacturing' industry portfolio.
hitec	Dummy variable set to 1 for the 'high technology' industry portfolio.
hlth	Dummy variable set to 1 for the 'healthcare' industry portfolio.

This table lists all the constructed variables used in this study, along with their detailed descriptions.

We use the following multivariate regression model for each of the five crash events with the dependent variable as the crash return:

$$crash.return = a_0 + a_1 beta + a_2 TCap + a_3 mkbk + a_4 dr + a_5 cr + a_6 cnsmr + a_7 hitec + a_8 hlth + a_9 manuf + \varepsilon \quad (1)$$

where a_0 is a constant (the intercept term), ε is the error term, and $a_1, a_2, \ldots a_9$ are the regression coefficients. The independent variables in the model are beta (*beta*), size (*TCap*), market-to-book ratio (*mkbk*), debt-to-equity ratio (*dr*), current ratio (*cr*), and the dummy variables for the industry portfolios (*cnsmr*, *hitec*, *hlth*, and *manuf*). The effect of the fifth portfolio, *allother*, is left in the intercept. We use the following multivariate regression model for each of the five post-crash market reversal events with the dependent variable as the recovery return:

Table 4: Summary Statistics

Variable	Min	1st Qu.	Median	Mean	3rd Qu.	Max	StDev
Panel 1: October 1987 crash (N = 2591)							
ATQ	1.149	31.832	91.966	816.51	316.44	70011	3563.8
LTQ	0.0780	12.045	41.785	453.38	160.12	37783	2000.2
TCap	1.464	30.892	83.050	770.58	320.23	90131	3393.4
cr	0.0087	1.556	2.234	2.913	3.418	10.280	2.134
dr	0.0106	0.4826	0.9761	1.306	1.636	5.978	1.221
mkbk	0.0058	1.337	1.999	2.578	3.109	11.733	2.024
beta	-2.388	0.1750	0.5484	0.5738	0.9545	4.553	0.6293
crash.return	-0.6316	-0.2609	-0.1905	-0.1931	-0.1228	0.5532	0.1020
recovery.return	-0.4576	-0.0837	-0.0282	-0.0130	0.0428	0.6667	0.1146
Panel 2: October 1997 Crash (N = 4642)							
ATQ	1.441	36.252	108.82	1051.8	449.40	97123	4393.2
LTQ	0.0260	9.947	40.838	610.93	230.78	58537	2701.8
TCap	1.832	54.739	165.41	1480.1	617.79	167169	7473.1
cr	0.0271	1.452	2.229	2.998	3.572	10.280	2.369
dr	0.0006	0.344	0.7755	1.213	1.552	5.978	1.310
mkbk	0.0029	1.710	2.800	3.747	4.768	11.733	2.932
beta	-5.023	0.0694	0.4545	0.4323	0.8180	4.833	0.6817
crash.return	-0.5000	-0.1299	-0.0798	-0.0862	-0.0385	0.4884	0.0768
recovery.return	-0.3824	-0.0134	0.0155	0.0237	0.0558	0.4444	0.0618
Panel 3: August 1998 Crash (N = 4443)							
ATQ	1.605	43.629	137.94	1217.4	561.02	93216	4661.2
LTQ	0.0470	12.086	51.172	712.19	289.75	58535	2839.1
TCap	1.482	44.312	139.24	1650.2	530.50	280297	9682.6
cr	0.0580	1.485	2.257	3.051	3.648	10.280	2.397
dr	0.0011	0.3489	0.7932	1.229	1.576	5.978	1.319
mkbk	0.0030	1.184	1.965	2.888	3.512	11.733	2.667
beta	-4.390	0.3863	0.7696	0.7988	1.179	3.778	0.6648
crash.return	-0.6826	-0.2048	-0.1349	-0.1439	-0.0769	0.5476	0.1011
recovery.return	-0.4326	-0.0192	0.0141	0.0227	0.0583	1.500	0.0852
Panel 4: April 2000 Crash (N = 4442)							
ATQ	1.786	51.508	176.84	1709.5	685.69	244192	8340.4
LTQ	0.0230	13.711	59.798	960.33	333.68	114871	4524.9
TCap	1.379	57.436	228.02	3019.5	915.03	490266	18498
cr	0.0447	1.397	2.210	3.161	3.798	10.280	2.619
dr	0.0008	0.3078	0.7918	1.230	1.594	5.978	1.334
mkbk	0.0011	1.120	2.364	3.928	5.645	11.733	3.698
beta	-3.641	0.1954	0.5736	0.6757	1.095	4.989	0.7540
crash.return	-0.7903	-0.2777	-0.1309	-0.1643	-0.0366	0.5000	0.1627
recovery.return	-0.5000	-0.0166	0.0250	0.0509	0.1021	1.250	0.1279
Panel 5: October 2008 Crash (N = 3277)							
ATQ	1.680	138.62	507.64	4962.0	2150.9	342679	19707
LTQ	0.0090	38.369	200.83	2812.1	1155.7	200770	11704
TCap	0.5337	113.40	414.16	3425.6	1568.4	403366	14974
cr	0.0000	1.343	2.077	2.828	3.366	10.280	2.294
dr	0.0003	0.3633	0.8211	1.252	1.556	5.978	1.338
mkbk	0.0021	0.9472	1.629	2.410	2.912	11.733	2.425
beta	-1.328	0.5052	0.8322	0.8328	1.138	3.076	0.5048
crash.return	-0.8080	-0.3386	-0.2441	-0.2542	-0.1697	0.4119	0.1338
recovery.return	-0.4110	0.0461	0.0960	0.1061	0.1532	1.100	0.0982
Panel 6: All Data (N = 19395)							
ATQ	1.149	46.394	160.85	1869.6	691.11	342679	9739.2
LTQ	0.0090	13.787	60.454	1065.0	339.91	200770	5709.7
TCap	0.5337	52.760	181.69	2105.6	736.12	490266	12388
cr	0.0000	1.443	2.206	3.007	3.594	10.280	2.396
dr	0.0003	0.3575	0.8221	1.239	1.581	5.978	1.311
mkbk	0.0011	1.233	2.167	3.210	3.970	11.733	2.955
beta	-5.023	0.2434	0.6337	0.6586	1.043	4.989	0.6793
crash.return	-0.8080	-0.2375	-0.1398	-0.1600	-0.0652	0.5532	0.1313
recovery.return	-0.5000	-0.0180	0.0242	0.0387	0.0862	1.500	0.1048

This table shows the summary statistics of the data samples for the five individual crash events in Panels 1-5, in chronological order, and for all data aggregated together in Panel 6. Included are the minimum, first quartile, median, mean, third quartile, maximum, and standard deviation. All variables are as defined earlier in Tables 2 and 3.

$$recovery.return = b_0 + b_1 crash.return + b_2 beta + b_3 TCap + b_4 mkbk + b_5 dr + \\ + b_6 cr + b_7 cnsmr + b_8 hitec + b_9 hlth + b_{10} manuf + e \quad (2)$$

where b_0 is a constant (the intercept term), e is the error term, and $b_1, b_2, \ldots b_{10}$ are the regression coefficients. The independent variables of Model (2) are the same as in Model (1) except the crash returns (*crash.return*) variable. Crash returns are used as an independent variable in Model (2) to determine if crash returns can explain post crash returns.

RESULTS AND DISCUSSION

Crash Periods

The multivariate regression analysis results for the five crash periods using Equation (1) are presented in Table 5. The F statistics indicate that all five regressions in the table are statistically significant. The explanatory power of the model varies between the events with the adjusted R-squared ranging from a low of 8.5 percent for 1997 to 44.9 percent for 2000. For all the events, the regression coefficient of beta is significant and negative, which indicates that stocks with a higher beta lost more value in all five stock market crashes relative to lower beta stocks. This result is in line with the CAPM, which predicts that stocks with higher betas would lose more value in down markets relative to low beta stocks. The regression coefficient of the size (TCap) variable is significant with a negative sign for the 1987 crash and with a positive sign for the 1998, 2000, and 2008 crashes. It is not statistically significant for the 1997 crash. The Fama and French (1992, 1993) three-factor CAPM argues that large firms are less risky than smaller firms. Therefore, the TCap variable should have a positive sign in a stock market crash. The 1998, 2000, and 2008 results confirm the prediction of the model. However, our results indicate that larger firms lost more value compared with smaller firms in the 1987 crash. There was a major market correction in stock prices in the 1987 crash. Investors might have thought that large firm stocks were more overvalued compared with small company stocks prior to the crash.

The regression coefficient of the market-to-book (mkbk) variable is significant with a negative sign for the 1987, 1997, 1998 and 2000 crashes and with a positive sign for the 2008 crash. According to Fama and French's (1992, 1993) three-factor CAPM, the market-to-book ratio is a risk factor in capital-asset pricing. A low market-to-book ratio implies that the firm may be in financial distress. These firms are expected to perform worse compared with high market-to-book firms in stock market crashes. Our regression result for the 2008 crash confirms the theory's prediction. However, the results for the 1987, 1997, 1998, and 2000 crashes do not support the theory's prediction. Wang et al. (2013) determine that bankruptcy risk was a serious concern for investors in the 2008 crash. Therefore, we find that firms with a low market-to-book ratio and greater risk lost more value in the 2008 crash. In the 1987, 1997, 1998, and 2000 crashes, however, perhaps investors considered firms with a high market-to-book ratio to be overvalued prior to crash and they simply bid down their prices more relative to low market-to-book ratio firms during the crash.

Our findings with the debt-to-equity (dr) variable also confirm our conclusion with the size (TCap) and mkbk variables. Bankruptcy risk was a significant concern for investors in the 2008 crash. Therefore, firms with a smaller size, lower mkbk, and higher dr lost more value in the 2008 crash. However, the dr variable is significant with a positive sign in the 1997, 1998, and 2000 crashes. Firms with a high dr performed better in these crashes relative to low dr firms. Since technical insolvency risk and bankruptcy risk were significant concerns for investors in the 2008 crash, like the dr variable, the current ratio (cr) variable is also significant for the 2008 crash with a positive sign. Firms with a higher cr (i.e., firms with a better ability to meet their maturing obligations) lost less value in the 2008 crash. We find a similar result for the 1987 crash. However, the sign of the cr variable is significant but negative for the 1998 and 2000 crashes. Firms with more investment in less profitable liquid assets lost more value in these crashes.

The regression coefficients for the industry dummy variables indicate that there is significant variation ithe industry effect between the five crash events. With the exception of the 1987 crash, it appears that the consumer goods industry segment (cnsmr) generally performed better during crashes with positive and significant regression coefficients. Firms in the 'hitec' industry group performed better than the average in the 2008 crash and worse than the average in the 1997, 1998 and 2000 crashes. The healthcare (hlth) industry regression coefficient is significant only for the 1987 and 2000 crashes with firms in this industry underperforming the average in the former and outperforming the average in the latter event. The regression coefficient for the manufacturing (manuf) industry segment is positive and significant for 1997, 1998, and 2000, and negative for 2008, indicating that manufacturing firms performed better than the average in the 1997,1998, and 2000 crashes and worse than the average in the 2008 crash. We observe a pattern of opposition between the effects of the hitec and manuf industry sectors in stock market crashes. Whenever the regression coefficient of one is positive the other is negative, and vice versa. The cnsmr sector appears to perform consistently better than the average in crashes.

Table 5: Multivariate Regression Analysis Results for the Crash Events

	Dependent Variable: Crash.Return									
	1987		**1997**		**1998**		**2000**		**2008**	
	Result	**VIF**	**Result**	**VIF**	**Result**	**VIF**	**Result**	**VIF**	**Result**	**VIF**
Intercept	-0.1473 *** (-21.271)		-0.0737 *** (-19.509)		-0.1069 *** (-20.962)		-0.0685 *** (-11.165)		-0.2271 *** (-29.607)	
beta	-0.0513 *** (-16.916)	1.072	-0.0147 *** (-9.027)	1.053	-0.0327 *** (-14.722)	1.051	-0.0438 *** (-15.922)	1.313	-0.0528 *** (-11.641)	1.081
TCap	-0.0024 *** (-4.274)	1.059	0.0000 (0.3134)	1.060	0.0006 *** (3.745)	1.063	0.0008 *** (7.863)	1.050	0.0003 ** (2.170)	1.029
mkbk	-0.0063 *** (-6.709)	1.078	-0.0036 *** (-9.197)	1.161	-0.0046 *** (-7.804)	1.197	-0.0133 *** (-22.522)	1.449	0.0057 *** (5.742)	1.191
dr	-0.0013 (-0.7928)	1.210	0.0061 *** (6.485)	1.302	0.0027 ** (2.203)	1.279	0.0106 *** (6.894)	1.275	-0.0109 *** (-5.647)	1.376
cr	0.0026 *** (2.637)	1.278	-0.0003 (-0.5933)	1.349	-0.0014 ** (-2.024)	1.275	-0.0089 *** (-10.937)	1.395	0.0041 *** (3.783)	1.271
cnsmr	-0.0020 (-0.329)	1.168	0.0135 *** (3.652)	1.216	0.0181 *** (3.677)	1.187	0.0292 *** (4.695)	1.356	0.0277 *** (3.914)	1.171
hitec	-0.0074 (-1.164)	1.168	-0.0163 *** (-4.481)	1.216	-0.0142 *** (-2.952)	1.187	-0.0569 *** (-9.313)	1.356	0.0197 *** (2.954)	1.171
hlth	-0.0257 *** (-3.127)	1.168	0.0048 (1.070)	1.216	-0.0009 (-0.1613)	1.187	0.0188 ** (2.516)	1.356	0.0018 (0.2137)	1.171
manuf	0.0015 (0.2521)	1.168	0.0095 ** (2.520)	1.216	0.0118 ** (2.348)	1.187	0.0519 *** (7.998)	1.356	-0.0342 *** (-4.832)	1.171
N	2591		4642		4443		4442		3277	
Adj. R-squared	0.1530		0.0853		0.1002		0.4492		0.1142	
F statistic	52.974 *** (0.0000)		49.092 *** (0.0000)		55.981 *** (0.0000)		403.36 *** (0.0000)		47.922 *** (0.0000)	

*This table shows the results of multiple regression analysis on the data for the individual events. The dependent variable is the crash return, and the independent variables are as listed in the table in the leftmost column. The regression specification is as follows: crash.return = a_0 +$a_1$beta + a_2TCap + $a_3$mkbk + $a_4$dr + $a_5$cr + $a_6$cnsmr + $a_7$hitec + $a_8$hlth + $a_9$manuf + ε Each column shows the regression results for one of the events, left to right in chronological order, as labeled. Shown are the coefficients with significance indicators, and t-statistics below in parentheses. All variables are as defined earlier in Tables 2 and 3, but with TCap rescaled to billions. The last three rows list the number of observations, the adjusted R-squared, and the F statistic (with p-value in parentheses), for each of the regressions. The variance inflation factors (VIF) are listed to the right of each coefficient for each regression. The VIF is used to test for multicollinearity in the model. Prior literature suggests that there is no major multicollinearity associated with a variable if the VIF value is less than 10 (Belsley et al., 2009). ***, **, * indicate significance at the 1, 5, and 10 percent levels, respectively.*

Post-Crash Market Reversal Periods

The results of the regressions specified by Equation (2) are shown in Table 6. The F statistics indicate that all five regressions are statistically significant. The explanatory power of the model varies between the events but is on average higher than in the crash regressions with the adjusted R-squared ranging from a 14.7 percent for 1998 to 36.5 percent for 2000.

Table 6: Multivariate Regression Analysis Results for the Post-Crash Market Reversal Events

	Dependent Variable: Recovery.Return									
	1987		**1997**		**1998**		**2000**		**2008**	
	Result	**VIF**	**Result**	**VIF**	**Result**	**VIF**	**Result**	**VIF**	**Result**	**VIF**
Intercept	-0.1407 *** (-19.266)		-0.0188 *** (-6.435)		-0.0226 *** (-5.159)		-0.0214 *** (-3.749)		-0.0043 (-0.7503)	
crash.return	-0.6303 *** (-32.912)	1.185	-0.3346 *** (-30.672)	1.095	-0.2936 *** (-23.842)	1.114	-0.3016 *** (-21.848)	1.819	-0.3111 *** (-26.534)	1.132
beta	-0.0024 (-0.7661)	1.190	0.0062 *** (5.097)	1.072	0.0097 *** (5.214)	1.103	0.0156 *** (5.997)	1.388	0.0304 *** (9.808)	1.126
TCap	0.0054 *** (9.809)	1.066	0.0003 *** (2.995)	1.060	0.0003 *** (2.775)	1.067	0.0002 ** (2.281)	1.064	0.0003 *** (3.346)	1.031
mkbk	0.0004 (0.3911)	1.097	0.0017 *** (5.613)	1.183	0.0008 * (1.693)	1.213	0.0026 *** (4.615)	1.615	0.0022 *** (3.267)	1.203
dr	-0.0007 (-0.4099)	1.210	-0.0012 * (-1.784)	1.314	-0.0028 *** (-2.751)	1.280	-0.0026 * (-1.803)	1.288	-0.0020 (-1.566)	1.389
cr	-0.0012 (-1.304)	1.282	-0.0002 (-0.4664)	1.349	-0.0004 (-0.6623)	1.276	0.0006 (0.7674)	1.432	-0.0006 (-0.7816)	1.277
cnsmr	0.0148 ** (2.514)	1.174	0.0084 *** (3.075)	1.246	-0.0020 (-0.4876)	1.206	0.0015 (0.2597)	1.495	-0.0127 *** (-2.666)	1.207
hitec	-0.0045 (-0.7287)	1.174	0.0103 *** (3.808)	1.246	-0.0055 (-1.387)	1.206	0.0079 (1.395)	1.495	0.0072 (1.609)	1.207
hlth	-0.0024 (-0.2971)	1.174	0.0056 * (1.689)	1.246	0.0023 (0.4889)	1.206	-0.0207 *** (-3.008)	1.495	0.0058 (1.037)	1.207
manuf	0.0144 ** (2.446)	1.174	0.0029 (1.037)	1.246	-0.0063 (-1.534)	1.206	0.0104 * (1.735)	1.495	0.0175 *** (3.677)	1.207
N	2591		4642		4443		4442		3277	
Adj. R-squared	0.3655		0.2203		0.1473		0.2476		0.2604	
F statistic	150.22 *** (0.0000)		132.14 *** (0.0000)		77.748 *** (0.0000)		147.17 *** (0.0000)		116.34 *** (0.0000)	

This table shows the results of multiple regression analysis on the data for the individual events. The dependent variable is the recovery return, and the independent variables are as listed in the table in the leftmost column. The regression specification is as follows: recovery.return = b_0 + $b_1$crash return + $b_2$beta + b_3TCap + $b_4$mkbk + $b_5$dr + $b_6$cr + $b_7$cnsmr + $b_8$hitec + $b_9$hlth + b_{10}manuf + e Each column shows the regression results for one of the events, left to right in chronological order, as labeled. Shown are the coefficients with significance indicators, and t-statistics below in parentheses. All variables are as defined earlier in Tables 2 and 3, but with TCap rescaled to billions. The last three rows list the number of observations, the adjusted R-squared, and the F statistic (with p-value in parentheses), for each of the regressions. The variance inflation factors (VIF) are listed to the right of each coefficient for each regression. The VIF is used to test for multicollinearity in the model. Prior literature suggests that there is no major multicollinearity associated with a variable if the VIF value is less than 10 (Belsley et al., 2009).
****, **, * indicate significance at the 1, 5, and 10 percent levels, respectively.*

For all the events, the regression coefficient of the crash return variable is significant with a negative sign indicating that firms that experience a larger negative return during the crash period make up for it with a larger positive return in the post-crash market reversal. It implies investor overreaction during the crash period with a significant market correction in the post-crash market reversal.

The regression coefficient of beta is significant with a positive sign for all events except the 1987 post-crash market reversal event. The 1997, 1998, 2000, and 2008 results are in line with the prediction of the CAPM. The model predicts that stocks with higher betas earn higher returns relative to low beta stocks in up markets. The regression coefficient of size (TCap) is significant with a positive sign for all post-crash market reversal events (i.e., large company stocks outperform small company stocks in all post-crash market reversals). This result is in line with the prediction of the Fama-French (1992, 1993) three-factor CAPM. The regression coefficient of the market-to-book ratio (mkbk) is positive and significant for all post-crash market reversal events except the 1987 event. The 1997, 1998, 2000, and 2008 results are in line with the prediction of the Fama-French model. The model predicts that high market-to-book stocks would outperform low market-to-book stocks in up markets.

The regression coefficient of the debt-to-equity ratio (dr) is negative and significant for the 1997, 1998, and 2000 market reversal events. Since these coefficients are positive and significant during the crash, it implies that investors overreacted by bidding down the prices of low dr firm stocks too much in these crashes, which resulted in a significant market correction after the crash. The regression coefficients for the current ratio (cr) variable are statistically insignificant in all five post-crash market reversal events. It implies that there was no investor overreaction to the cr variable in the market crashes which would have resulted in a significant market correction in the post-crash market reversal. The signs and significance of the regression coefficients for the industry portfolios vary between the events. The industry effect appears to be generally less significant in the post-crash market reversal period than in the crash period. The results imply investor overreaction in the hitec industry in the 1997 crash and in the manuf industry in the 2008 crash. Stocks that lost more value than the average in the hitec industry in the 1997 crash and in the manuf industry in the 2008 crash gained more value than the other stocks in the post-crash market reversal.

Combined Data for All Five Crashes

Although the five stock market crash events have a number of distinct characteristics, running regressions with the combined sample may provide some useful insights about the overall mean effects of the variables across all crashes. We present our regression results with the entire data set for all five market crash and post-crash market reversal events in Table 7. The F statistics indicate that both the crash regression and the post-crash market reversal regression are statistically significant at the 1-percent level.

The regression coefficient of the crash return variable is significant with a negative sign for the post-crash market reversal. It indicates that stocks that lose more value in crashes tend to gain more value in post-crash market reversals. It implies investor overreaction in stock market crashes. The regression coefficient of beta is significant in both regressions and it has a negative sign for the crash and a positive sign for the post-crash market reversal. It implies that stocks with higher betas lose more value in crashes and they gain more value in post-crash market reversals relative to low beta stocks. This finding is in line with the prediction of the CAPM that high beta stocks lose more value in down markets and gain more value in up markets relative to low beta stocks. The regression coefficient of size (TCap) is significant with a positive sign in both regressions. It implies that stocks with larger market capitalization perform better compared with stocks with smaller market capitalization both in crashes and in post-crash market reversals. This finding is in line with the Fama-French (1992, 1993) three-factor CAPM, which argues that large firms are less risky and investors require lower returns from these firms.

Table 7: Multivariate Regression Analysis Results with All data

	Crash		Recovery	
	Result	VIF	Result	VIF
Intercept	-0.1105 *** (-37.061)		-0.0303 *** (-13.120)	
crash.return			-0.3533 *** (-65.620)	1.175
beta	-0.0596 *** (-45.334)	1.056	0.0128 *** (12.327)	1.168
TCap	0.0003 *** (3.967)	1.039	0.0005 *** (8.519)	1.040
mkbk	-0.0043 *** (-13.529)	1.173	0.0020 *** (8.409)	1.184
dr	0.0027 *** (3.682)	1.262	-0.0022 *** (-3.918)	1.263
cr	-0.0024 *** (-5.855)	1.311	-0.0008 ** (-2.508)	1.313
cnsmr	0.0258 *** (8.855)	1.197	0.0000 (-0.0178)	1.214
hitec	-0.0131 *** (-4.573)	1.197	0.0032 (1.490)	1.214
hlth	0.0108 *** (3.048)	1.197	0.0004 (0.1694)	1.214
manuf	0.0165 *** (5.553)	1.197	0.0036 (1.609)	1.214
N	19395		19395	
Adj. R-squared	0.1488		0.2487	
F statistic	377.62 *** (0.0000)		643.08 *** (0.0000)	

This table shows the results of multiple regression analysis on the entire data sample. In the first column, labeled 'Crash', the dependent variable is the crash return, and the independent variables are as listed in the table in the leftmost column: crash.return = a_0 +$a_1$beta + a_2TCap + $a_3$mkbk + $a_4$dr + $a_5$cr + $a_6$cnsmr + $a_7$hitec + $a_8$hlth + $a_9$manuf + ε For the second column, labeled 'Recovery', the dependent variable is the recovery return, and the independent variables are as listed in the table in the leftmost column: recovery.return = b_0 + $b_1$crash return + $b_2$beta + b_3TCap + $b_4$mkbk + $b_5$dr + $b_6$cr + $b_7$cnsmr + $b_8$hitec + $b_9$hlth + b_{10}manuf + e Shown are the coefficients with significance indicators, and t-statistics below in parentheses. All variables are as defined earlier in Tables 2 and 3, but with TCap rescaled to billions. The last three rows list the number of observations, the adjusted R-squared, and the F statistic (with p-value in parentheses), for each of the regressions. The variance inflation factors (VIF) are listed to the right of each coefficient for each regression. The VIF is used to test for multicollinearity in the model. Prior literature suggests that there is no major multicollinearity associated with a variable if the VIF value is less than 10 (Belsley et al., 2009).
****, **, * indicate significance at the 1, 5, and 10 percent levels, respectively.*

When the regression coefficient of a variable is significant with different signs in the crash and post-crash market reversal periods, it implies investor overreaction during the crash. The regression coefficient of the mkbk variable is significant in both regressions and it has a negative sign for the crash and a positive sign for the post-crash market reversal. It implies that investors consider high mkbk stocks to be overvalued prior to crashes and they bid down their prices more relative to low mkbk stocks in stock market crashes. However, there is a significant market correction for the prices of high mkbk stocks in the post-crash reversals implying investor overreaction towards these stocks during the crash.

The regression coefficient of the debt-to-equity ratio (dr) variable is significant in both regressions and it has a positive sign for the crash and a negative sign for the post-crash market reversal implying investor overreaction during the crash period. The result implies that the stocks of firms with higher debt ratios generally perform better in crashes (excluding the 2008 crash when investors had a serious concern with bankruptcy risk) but they perform worse in post-crash market reversals compared with the stocks of firms with lower debt ratios. The regression coefficient of current ratio (cr) is significant with a negative sign in

both regressions. It implies that the stocks of firms with more investment in less profitable current assets generally perform worse both in crashes and in post-crash market reversals. However, this is an aggregate result for all crashes. Because technical insolvency risk was a major concern for investors, low cr firms lost more value relative to high cr firms in the 1987 and 2008 stock market crashes (see Table 6). All regression coefficients for the industry dummy variables are statistically significant for the crash period. The sign of the regression coefficient for the cnsmr, hlth, and manuf industries is positive. It implies that the stocks of firms in these industries generally perform better than the average in stock market crashes. The sign of the regression coefficient for the hitec industry is negative in the crash regression. It implies that the stocks in this industry generally perform worse than the average in crashes. The regression coefficients of all four industries are insignificant for the post-crash market reversal period. This implies that the stocks in all four industries generally perform similarly in post-crash market reversals with no major market correction for any industry to correct an overreaction during the crash periods.

CONCLUSION

In this paper, we study the determinants of stock returns in five major stock market crashes and post-crash market reversals during the 1987-2008 period to investigate if there was any investor overreaction in these crashes. Using daily closing prices we calculate cumulative returns for the crash and reversal periods for the events listed in Table 1, and regress crash and reversal returns on a number of firm characteristics. The regression coefficient of the crash return variable is statistically significant with a negative sign in all post-crash market reversal regressions. This result implies that there is investor overreaction in stock market crashes. Stocks that lose more value in crashes tend to gain more value after the crash with a significant market correction in the post-crash market reversal. Sharpe's CAPM predicts that high beta firms lose more value in down markets and gain more value in up markets compared with low beta firms. As predicted by the theory, in this paper, we find that high beta companies lose more value in stock market crashes and gain more value in post-crash market reversals.

In the Fama-French (1992, 1993) three-factor CAPM, in addition to beta, firm size and market-to-book ratio are also market risk factors and determinants of stock returns. The model argues that smaller firms and those with lower market-to-book ratios are riskier. Therefore, investors would require a higher rate of return with a larger risk premium when valuing these firms. As predicted by the theory, we find that smaller firms and those with lower market-to-book ratios lose more value in stock market crashes. However, the sign of the regression coefficients for these variables does not change in the post-crash market reversals. It implies that investor reaction against smaller and lower market-to-book ratio firms in stock market crashes is not an overreaction. The regression coefficient of the debt ratio (dr) variable is significant with a positive sign in the 2008 crash. Since bankruptcy risk was a serious concern for investors, high-dr firms lost more value relative to low-dr firms in the 2008 crash. However, the regression coefficient for the dr variable is not significant in the post-crash market reversal. It implies that investors' bidding down the prices of high dr firms was not an overreaction in the 2008 crash.

Our crash regressions show that, because technical insolvency risk was an important concern for investors, firms with a higher current ratio (cr) and thus greater ability to meet their maturing obligations, lost less value relative to lower cr firms in the 1987 and 2008 crashes. However, the regression coefficient for the cr variable is not significant in the post-crash market reversal regressions. It implies that investors' bidding down the prices of low cr firms was not an overreaction in the 1987 and 2008 crashes.

The industry dummy variables indicate that there is no specific pattern of industry effect in stock market crashes. However, investors appear to have overreacted against high tech stocks in the 1997 crash and against manufacturing stocks in the 2008 crash with a significant market correction in the values of these stocks in the post-crash market reversals. The present research has several limitations. First, in this study we have a relatively short term definition of market crash and reversal, requiring consecutive stock index

price declines for the crash, and consecutive increases for the reversal. It would be instructive to look at longer-term market crash and reversal periods. Additionally, we only look at U.S. firms; it is possible that international markets would exhibit different patterns of investor behavior. Investigating these issues in greater detail should be fertile ground for future research.

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INFLATION TARGETING AS A POSSIBLE MONETARY FRAMEWORK FOR NIGERIA

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ABSTRACT

One of the issues facing Nigeria today is the choice among two nominal anchors: exchange rate pegging or inflation targeting. The incessant increase in interest rates, exchange rates, money supply and domestic credit have all accumulated, leading to persistent inflation in Nigeria. At this instant, it is pertinent to look for another nominal anchor to keep inflation in check because the present exchange rate pegging seems useless. This groundbreaking study, in an effort to do this, examines inflation targeting as a possible monetary framework for Nigeria, using time series data and with the aid of Granger Causality test and impulse response functions. The empirical results show evidence that inflation is highly sensitive to exchange rate and interest rate while economic growth is highly sensitive to exchange rate and inflation in Nigeria. Further, the causation from real exchange rate to economic growth is stronger than the causation from inflation to economic growth, meaning exchange rate determines economic growth in Nigeria more than inflation does. Therefore, inflation targeting will be less preferable to exchange rate targeting in Nigeria as a policy alternative. This unexpected finding has important implications for monetary policy conduct in Nigeria.

JEL: E31, E52, E44, E58

KEYWORDS: Inflation Targeting, Monetary Policy, Inflation

INTRODUCTION

Inflation targeting, as an economic policy, is an attempt to direct inflation towards an expected, or "target" inflation rate using monetary tools such as interest rate changes (Coy, 2005). Under the policy, the actions of the central bank become more transparent. Investors, knowing what the central bank estimates as the target inflation rate, can easily factor in possible interest rate changes in their investment sets, leading to better economic stability. One of the issues facing Nigeria today is the choice among two nominal anchors: exchange rate pegging or inflation targeting. Volatility in price and hyperinflation are huge economic challenges, able to create financial instability and tumble economies. Many industrialized economies, after experiencing persistent inflation rates for decades, have reduced inflation to extremely low levels recently with the aid of inflation targeting. Nigeria can do the same. In fact, inflation targeting frameworks have regularly and successfully been adopted in economies suffering from chronically high inflation. Can inflation targeting regime work effectively in Nigeria which employs the exchange rate stability objectives? Bakradze and Billmeier's (2007) study observed that the rising number of countries embracing inflation targeting and its success are inducements for countries that employ monetary or exchange rate targeting to consider a change to inflation targeting.

Thus, countries like Nigeria need to consider earnestly such a change now or in the near future, bringing us to the question if Nigeria is ready for inflation targeting now, later or maybe never. According to Mishkin (2000), for inflation targeting to successfully raise output growth, lower unemployment, increase external competitiveness -- through monetary policy, there must exist a strong institutional commitment to make price stability the primary goal of the central bank. This is particularly important in an emerging market

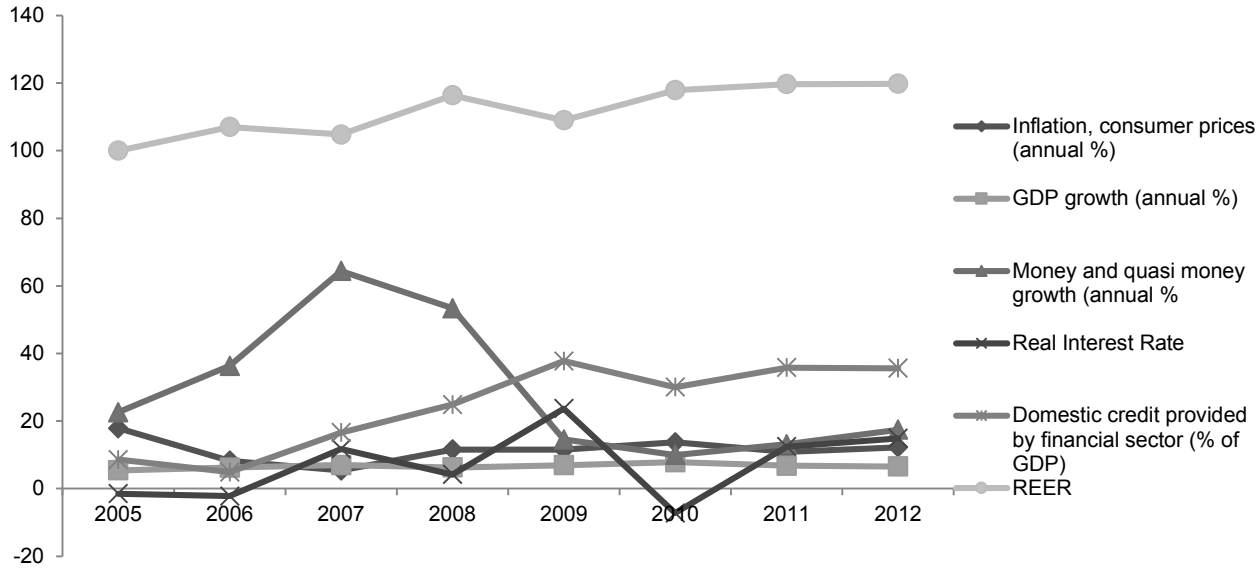
country such as Nigeria which has often had a past history of monetary mismanagement. Table 1 and Figure 1 show the evolution of the monetary policy outcomes in Nigeria from 2005 till 2012. It shows the quantum of monetary policy in Nigeria. Price stability is not accorded the highest priority in Nigeria, to the detriment of the aforementioned "institutional commitment to price stability which requires that the central bank be given a mandate to have price stability as its primary goal, making it clear that when there is a conflict with other goals, such as exchange rate stability or promotion of high employment, price stability should be accorded the highest priority" (Mishkin, 2000, pp. 3). A look at Table 1 and Figure 1 shows that interest and inflation rates in Nigeria have been on a double-digit value averagely over the period of 2005-2012. In recent years, inflation in Nigeria has been steadily above 10%, except 2007. As well, interest rate has been high, especially excessively highest in 2011 and 2012. The real exchange rate against the US dollar rises over the years. In the same vein, the consumer price index (CPI) is not left out in the steady increase. Within the same period, money and quasi money growth has reduced while Domestic credit provided by financial sector (% of GDP) has increased remarkably. The increase in interest rates, exchange rates, persistent growth in money supply and domestic credit have all accumulated, leading to persistent inflation in Nigeria.

Table 1: Monetary Policy Outcomes (2005-2012)

	2005	2006	2007	2008	2009	2010	2011	2012
Inflation, consumer prices (annual %)	17.863	8.240	5.382	11.578	11.538	13.720	10.841	12.217
GDP growth (annual %)	5.400	6.211	6.972	6.270	6.934	7.840	6.791	6.531
Money and quasi money growth (annual %)	22.604	36.351	64.417	53.360	14.543	9.969	13.142	17.416
Real Interest Rate	-1.513	-2.214	11.764	4.190	23.707	-7.231	12.416	14.870
Domestic credit provided by financial sector (% of GDP)	8.600	4.909	16.575	24.891	37.772	30.009	35.800	35.617
Real Exchange rate	100.0	107.0	104.8	116.4	109.0	117.9	119.7	119.8

This table shows the evolution of the monetary policy outcomes in Nigeria from 2005 till 2012. It shows the quantum of monetary policy in Nigeria.

Figure 1: Monetary Policy Outcomes (2005-2012)



This figure shows the evolution of the monetary policy outcomes in Nigeria from 2005 till 2012. It shows the size and trend of the variables, thus depicting the thrust of monetary policy in the recent years.

Thus, at this instant, it is pertinent to look for another nominal anchor to keep inflation in check because the present exchange rate pegging seems useless. This groundbreaking study, in an effort to do this, examines inflation targeting as a possible monetary framework for Nigeria. The remainder of this paper is structured as follows: section 2 reviews the literature. Section 3 presents the data and methodology. Sections 4 give the results. Section 5 concludes.

LITERATURE REVIEW

A popular choice since the early 1990s, inflation targeting has gained adherence from more than twenty countries in developed and emerging-market economies. Countries have adopted inflation targeting under varying conditions, ranging from the answer to a currency crisis (e.g. United Kingdom) to a planned switch from a completely different policy regime (e.g. Canada and New Zealand). Likewise inflation targeting has been practiced with varying verve and under diverse institutional arrangements (Bamidele, 2007). Forged in 1990 in New Zealand, inflation targeting is now in use by the central banks of Canada (Bank of Canada), United Kingdom (Bank of England), Australia (Reserve Bank of Australia), Iceland (Central Bank of Iceland) South Korea (Bank of Korea), Egypt, and Brazil (Brazilian Central Bank) South Korea (Bank of Korea), and Brazil (Brazilian Central Bank) and South Africa (South African Reserve Bank), among others, and empirical evidence shows that it does what its proponents claim (Coy, 2005). Only two countries, in Sub-Saharan Africa, have officially embraced inflation targeting: Ghana and South Africa (Hajj et al., 2013). Till date, the outcome of prior studies on the performance of inflation targeting has been diverse.

The first set of empirical studies finds no significant improvement in the economies between pre- and post-inflation targeting or between the economies of inflation-targeting countries and non- inflation targeting countries (for examples Cecchetti & Ehrmann, 2000; Honda, 2000; Ball & Sheridan, 2005; Berument & Yuksel, 2006). The second set finds meaningful improvement as inflation targeting causes improvement in economic structure and inflationary path (for examples Garcia, 2000; Pétursson, 2004 and so on). Yet, a set of studies, constructing indicators to measure the impact of inflation targeting, evaluates the performance of inflation targeting through disinflation cost and observations of country-specific data (e.g. Pétursson, 2004). Some authors assess the impact of inflation targeting from the perspective of the cost of disinflation (i.e. the ratio of loss in output divided by the fall in inflation). The studies, evaluating the impact of inflation targeting employing the cost of disinflation include Senda & Smith, 2008 Tunali, 2008; Goncalves & Carvalho, 2006 and so on. All this previous works have shown that the performance of inflation targeting differs across dimensions, countries and over time (Ramos-Francia & Capistran, 2007; Mishkin & Schmidt-Hebbel, 2007). As well, evidence suggests that the credibility of the central bank and the economic structure are factors in the various outcomes of inflation targeting. For instance, Fraga, Goldfajn & Minella's (2003) study demonstrates that inflation targeting is more successful in developed economies compared to emerging market economies in terms of reduced volatility in output, inflation, exchange rate and interest rate. Additional factors shaping the performance of inflation targeting include type of demand or supply shock (Lai & Chang, 2001) and exchange rate (Bleaney, 2000 and Brenner & Sokoler, 2006).

Mthuli Ncube and Eliphas Ndou (2012), using a Bayesian VAR sign restriction approach, derives the inflation equation to comb for a plausible transmission channel between the inflation rate, real interest rate, exchange rates and real output growth rate. The empirical findings indicate that the real interest rate responds negatively to inflation rate shocks; in the long run the Fisher effect holds. They demonstrate that strict inflation targeting is incompatible with significant output growth. Conversely, a flexible inflation-targeting framework which places importance on real effective exchange rates leads to a significant real output growth. Hajj, Dufrénot, Sugimot and Wolf's (2013) study examines the monetary policy actions with which Sub-Saharan African central banks have sought to reduce or eliminate the negative consequences of the shocks confronting their economies. Comparing two types of monetary policy regimes: a currency board regime in the CFA zone countries and an inflation targeting policy regime in Ghana and

South Africa, they found that both policies are unsuitable for economies exiting from the impacts of negative demand shocks. However, both policies are vital when negative shocks to primary balance arise. Alvaro Angeriz and Philip Arestis (2006), using intervention analysis on structural time series models of ten countries, investigates the empirical aspects of inflation targeting. The outcome shows that if the initial impacts of inflation targeting are taken into consideration, central banks, that have followed this strategy, have been unsuccessful. Sek (2006) evaluates inflation targeting in three emerging East-Asian economies: Korea, Philippines and Thailand by comparing the changes in the economy between the pre- and post-inflation targeting periods. Applying a bivariate GARCH (1,1) model to study the relationship between inflation and output gap, he detects lower inflation rate in the post-IT period and no significant correlation between inflation and output gap was found. He concludes that inflation targeting has bettered the economies of those countries. Besides, inflation targeting and the exchange rate flexibility have a close relationship. Exchange rate, a vital instrument in an open economy like Nigeria, plays as a transmission channel for monetary policy and simultaneously as an influencer of the real economy. Undue volatility of exchange rate can be injurious to trade and growth. Thus, responding to inflation and exchange rate variability conjointly in the policy function can lead to the risk of tradeoff or compromise between inflation and exchange rate variability.

If, for example the central bank sees impending increase in inflation; to tighten the price of tradable goods, interest rates are raised. As soon as the inflation is under control, interest rates are lowered and the exchange rate depreciates. Thus the fall in inflation variability has brought about the rise in the volatility of exchange rate. Exchange rate stability is inconsistent with inflation targeting regime; inflation targeting regime certainly necessitates exchange rate flexibility (Debelle, 2000). Foreign exchange intervention policy and inflation targeting, according to Brenner & Sokoler (2006), cannot coexist because there is conflict between the two policies. Taguchi and Kato's (2011) assessment of inflation targeting in some East-Asian economies shows that flexibility in exchange rate is a prerequisite to the success of inflation targeting regime. Then again, there are contradictory views that intermediate regimes would be good for inflation targeting. In fact, the case of Chile and Israel show that exchange rate objectives is containable within an inflation targeting regime (Debelle, 2000). Finally, there is an argument that inflation targeting cannot work well in emerging markets, like Nigeria, as emerging markets are deficient of the preconditions for a proper operation of inflation targeting. According to Kadioğlu et. al (2000), the prerequisites for the success of inflation targeting consist of sound economic structure; exchange rate flexibility; central bank independence; the institutional set-up; political commitment; a great deal of transparency and accountability of the Central Bank; absence of fiscal dominance; a single, clear inflation target; a sound inflation forecasting model; virile financial markets. The absence or inadequacy of these prerequisites may pose huge challenges for emerging markets like Nigeria trying to embrace inflation targeting.

DATA AND METHODOLOGY

Model

After a meticulous review of foregoing studies and refining upon the theoretical postulates explicated above, the two models for this study are expressed as follows:

$$INFLATION_t = \alpha_0 + \alpha_1 GROWTH_t + \alpha_2 MONEY_t + \alpha_t INTEREST_t + \alpha_4 CREDIT_t + \alpha_5 EXCHANGE_t + \alpha_6 EXPENDITURE_t + \alpha_7 OIL_t + \xi_t \quad (1)$$

$$GROWTH_t = \beta_0 + \beta_1 INFLATION_t + \beta_2 MONEY_t + \beta_t INTEREST_t + \beta_4 CREDIT_t + \beta_5 EXCHANGE_t + \beta_6 EXPENDITURE_t + \beta_7 OIL_t + \xi_t \quad (2)$$

Where:

Inflation	=	Inflation, Consumer Prices (Annual %)
Growth	=	Gdp Growth (Annual %)
Money	=	Money And Quasi Money Growth (Annual %)
Interest	=	Real Interest Rate
Credit	=	Domestic Credit Provided By Financial Sector (% Of Gdp)
Exchange	=	Real Exchange Rate
Expenditure	=	Recurrent Government Expenditure
Oil	=	Oil Revenue
Trade	=	Trade Openness
Capital	=	Gross Capital Formation

Vector Autoregressive Model

The vector autoregressive model (VAR) is used to analyse the variables' system. Each endogenous variable of the system is a function of the lagged values of the endogenous variables.
The VAR model is as follows:

$$y_t = A_1 y_{t-1} + A_2 y_{t-2} + ... + A_p y_{t-p} + c + \varepsilon_t \quad (3)$$

Where:

y_t is a vector of n endogenous variables,
x_t is a vector of m exogenous variables,
A_1, A_2..., A_p are matrices of the parameters being estimated
c is the constant term
ε_t is a vector of terms produced by a white noise process with these proprieties:

$$E[\varepsilon_t] = 0 \quad \forall t$$
$$E[\varepsilon_t \varepsilon'_t] = \begin{cases} \Omega & s = t \\ 0 & s \neq t \end{cases} \quad (4)$$

This shows that the ε's are serially uncorrelated.

Granger Causality

The Granger (1969) approach is used to investigate how much of the current y is explained by the lagged values of y and if, after adding past values of x we can increase the explanation of the model. Succinctly, we say "x Granger causes y" if the coefficients of the lagged variables of x are statistically significant.
The Granger causality entails the estimation of two regressions like the following:

$$\begin{aligned} y_t &= \alpha_0 + \alpha_1 y_{t-1} + ... + \alpha_l y_{t-l} + \beta_1 x_{t-1} + ... + \beta_l x_{t-l} + \varepsilon_t \\ x_t &= \alpha_0 + \alpha_1 x_{t-1} + ... + \alpha_l x_{t-l} + \beta_1 y_{t-1} + ... + \beta_l y_{t-l} + u_t \end{aligned} \quad (5)$$

for all the possible values of the series (x,y).

The Granger causality test implies the F Wald test for the joint hypotheses $\beta_1 = \beta_2 = ... = \beta_l = 0$ for each equation. The hypotheses are expressed as:

H_0: 'x does not Granger cause y', in one equation, and
H_1: 'y does not Granger cause x', in the other.
This test statistic can be expressed as:

$$F = \frac{(SQEr - SQEnr)/m}{SQEnr/(n-k)} \quad (6)$$

Where:

m is the number of lagged terms of Y
k is the number of parameters estimated without restrictions,
SQEr is the sum of squared errors in the restraint regression (when H_0 is true) and
SQEnr is the sum of squared errors with the unrestricted regression.
With m and n-k degrees of freedom, this statistic follows the F-distribution.

Impulse Response Function

Modeled in the framework of a vector autoregression, the impulse response functions are used to describe how the economy responds over time to shocks or exogenous impulses. Succinctly, it is used to describe the reaction of endogenous variables such as inflation, GDP growth, money growth, interest rate, and exchange rate at the time of the shock and over succeeding points in time.

Data Bank

The data covers annual data between the years between 1980-2012. The values employed in the empirical analysis were mined from the data banks of the World Bank and the Central Bank of Nigeria. They are published on their sites, www.cenbank.org/documents/data.asp and data.worldbank.org/Indicators. They are Inflation, consumer prices (annual %); GDP growth (annual %); money and quasi money growth (annual %); real interest rate; domestic credit provided by financial sector (% of GDP); real exchange rate; growth rate of recurrent government expenditure, growth rate of oil revenue, growth rate of trade openness and growth rate of capital formation.

Table 2: Augmented Dickey Fuller Unit Root Test (Trend and Intercept)

Series	ADF Test Statistic	5% Critical Values	10% Critical Values	Order	Remarks
Growth	-4.948	-3.556	-3.211	I(1)	Stationary
Money	-6.949	-3.556	-3.211	I(1)	Stationary
Interest Rate	-7.934	-3.556	-3.211	I(1)	Stationary
Credit	-5.314	-3.556	-3.211	I(1)	Stationary
Exchange Rate	-6.904	-3.556	-3.211	I(1)	Stationary
Expenditure	-5.119	-3.556	-3.211	I(1)	Stationary
Oil Revenue	-5.951	-3.556	-3.211	I(1)	Stationary
Trade	-5.897	-3.556	-3.211	I(1)	Stationary
Capital	-5.843	-3.556	-3.211	I(1)	Stationary
Inflation	-5.788	-3.556	-3.211	I(1)	Stationary

This table shows that all the time series are I(1). They have the same order of integration, as the Augmented Dickey Fuller Unit Root Test (Trend and Intercept) shows.

RESULT AND DISCUSSION

Unit-Root and Cointegration Tests

The first thing is to determine the order of integration of the individual time series using Augmented Dickey Fuller Unit Root Test (Trend and Intercept) and Phillips-Perron Unit Root Test (Trend and Intercept) as shown in Table 2 and Table 3.

Table 3: Phillips-Perron Unit Root Test (Trend and Intercept)

Series	P Test Statistic	5% Critical Values	10% Critical Values	Order	Remarks
Growth	-4.917	-3.556	-3.211	I(1)	Stationary
Money	-7.311	-3.556	-3.211	I(1)	Stationary
Interest Rate	-8.933	-3.556	-3.211	I(1)	Stationary
Credit	-5.316	-3.556	-3.211	I(1)	Stationary
Exchange Rate	-7.973	-3.556	-3.211	I(1)	Stationary
Expenditure	-6.386	-3.556	-3.211	I(1)	Stationary
Oil Revenue	-7.377	-3.556	-3.211	I(1)	Stationary
Trade	-7.540	-3.556	-3.211	I(1)	Stationary
Capital	-7.704	-3.556	-3.211	I(1)	Stationary
Inflation	-7.867	-3.556	-3.211	I(1)	Stationary

This table shows that all the time series are I(1). They have the same order of integration using Phillips-Perron Unit Root Test (Trend and Intercept).

Since most of the time series have the same order of integration, we tested and saw they are cointegrated, using Johansen's methodology as shown in Table 4 below. Both the maximum eigenvalue and trace tests reject the null hypothesis of no cointegration (r = 0), at both 5 per cent and 10 per cent levels of significance. Consequently the results accept the alternative hypothesis of r =1. This implies that they have cointegration relation and that there is only one cointegrating vector. Thus, the VAR model is set up in the levels of the data. 2 is the maximum lag length for the variables in the VAR, based on the AIC. The VAR is well-specified; we ensure that no serial correlation exists in the residuals for a reliable result. Straight interpretation of VAR model is very longwinded and can lead to poor conclusions. Instead, this study interprets the Granger causality test and impulse response functions (IRF).

Table 4: Johansen Multivariate Co-Integration Test

Unrestricted Cointegration Rank Test (Trace)				
Hypothesized No. of CE(s)	**Eigenvalue**	**Trace Statistic**	**0.05 Critical Value**	**Prob.****
None *	0.4792	53.208	47.856	0.0145
At most 1	0.2717	27.764	29.797	0.0843
At most 2	0.2473	15.396	15.494	0.0517
At most 3 *	0.1047	4.3157	3.8414	0.0378
Trace test indicates 1 cointegrating eqn(s) at the 0.05 level				
Unrestricted Cointegration Rank Test (Maximum Eigenvalue)				
Hypothesized No. of CE(s)	**Eigenvalue**	**Max-Eigen Statistic**	**0.05 Critical Value**	**Prob.****
None	0.4792	25.443	27.584	0.0917
At most 1	0.2717	12.368	21.131	0.5118
At most 2	0.2473	11.080	14.264	0.1502
At most 3 *	0.1047	4.3157	3.8414	0.0378
Max-eigenvalue test indicates 1 cointegration at the 0.05 level				

This table shows that the variables are cointegrated, using Johansen's methodology. The maximum eigenvalue and trace tests reject the null hypothesis of no cointegration (r = 0), at both 5 per cent and 10 per cent levels of significance.

Granger Causality Tests

The Granger (1969) test is employed to test for causality as shown in Table 5, providing an evaluation of causation from one variable to the other. This table suggests that exchange rate determines economic growth in Nigeria more than inflation does. Thus, inflation targeting will be less preferable to exchange rate targeting in Nigeria as a policy alternative.

Table 5: Granger Causality Wald Tests

	Excluded	chi2	df	Prob > chi2
INFLATION	Growth	4.052	2	0.132
	Money	12.616	2	0.002***
	Interest Rate	0. 142	2	0.931
	Credit	15.764	2	0.000***
	Exchange Rate	17.019	2	0.000***
	Expenditure	7.014	2	0.030**
	Oil Revenue	0.338	2	0.844
	Trade	1.303	2	0.521
	Capital	1.345	2	0.510
	All	91.822	18	0.000***
GROWTH	Inflation	5.894	2	0.049**
	Money	3.563	2	0.168
	Interest Rate	2.154	2	0.341
	Credit	3.754	2	0.153
	Exchange Rate	6.710	2	0.035**
	Expenditure	1.110	2	0.574
	Oil Revenue	1.233	2	0.540
	Trade	3.549	2	0.617
	Capital	.9667	2	0.020**
	All	32.301	18	0.020**

*note: *** and ** denote rejection of the exclusion at the 1 and 5 per cent level. the table shows* and provides an evaluation of causation from one variable to the other using granger (1969) test.

From the results in Table 5, the following causalities have been detected: Causation from MONEY to INFLATION suggesting that the amount of money in the economy determines the quantum of inflation Causation from the CREDIT to INFLATION suggesting that the amount of credit in the economy determines the quantum of inflation. Causation from the REAL EXCHANGE RATE to INFLATION suggesting that the level of real exchange rate determines the quantum of inflation.

Causation from GOVERNMENT RECURRENT EXPENDITURE to INFLATION suggesting that the amount of government recurrent expenditure in the economy determines the quantum of inflation. Causation from INFLATION to ECONOMIC GROWTH suggesting that the amount of inflation in the economy determines the quantum of inflation. Causation from REAL EXCHANGE RATE to ECONOMIC GROWTH suggesting that the level of real exchange rate determines the quantum of inflation. Causation from GROSS CAPITAL FORMATION to ECONOMIC GROWTH suggesting that the amount of gross capital formation in the economy determines the quantum of inflation

The causation from REAL EXCHANGE RATE to ECONOMIC GROWTH is stronger than the causation from INFLATION to ECONOMIC GROWTH suggesting that exchange rate determines economic growth in Nigeria more than inflation. Thus, inflation targeting will be less preferable to exchange rate targeting in Nigeria as a policy alternative.

Impulse Response Functions

The impulse response functions (IRF) shows that Economic Growth responds positively to innovations (i) in its own impulses over the 10 years period and (ii) in real exchange rate over the 10 years period. It responds negatively to innovations (i) in inflation over the 10 years and (ii) in real interest rate over the 10 years. The figure shows that the intensity of economic growth reaction is greater for exchange rate, interest and inflation, than for money and the reaction stays steady and durable. As well, the impulse response

functions (IRF) shows that inflation responds positively to innovations (i) in its own impulses over the first 7 years period and (ii) in money over the first 6.5 years. It responds negatively to innovations (i) in the interest rate over the first 4 years, (ii) in economic growth over the first 3.5 years, and (iii) in the exchange rate over the 5 years. The figure shows that the intensity of inflation reaction is greater for exchange rate and interest, than for money and economic growth and the reaction stays steady and durable. Summarily, the impulse response function shows that inflation is highly sensitive to exchange rate and interest rate while economic growth is highly sensitive to exchange rate and inflation.

CONCLUDING COMMENTS

The paper has examined inflation targeting as a possible monetary framework for Nigeria. Particularly, the paper has focused on the relative causality and responses between economic growth, inflation, exchange rate, interest rate, money and credit. The empirical methodology uses Granger causality and impulse response functions. The data (1980-2012) were mined from the data banks of the World Bank and the Central Bank of Nigeria. The variables are Inflation, consumer prices (annual %); GDP growth (annual %); money and quasi money growth (annual %); real interest rate; domestic credit provided by financial sector (% of GDP); real exchange rate; growth rate of recurrent government expenditure, growth rate of oil revenue, growth rate of trade openness and growth rate of capital formation. As to the empirical results of the paper, we have evidence that inflation is highly sensitive to exchange rate and interest rate while economic growth is highly sensitive to exchange rate and inflation in Nigeria. Further, the causation from real exchange rate to economic growth is stronger than the causation from inflation to economic growth, meaning exchange rate determines economic growth in Nigeria more than inflation does. Therefore, inflation targeting will be less preferable to exchange rate targeting in Nigeria as a policy alternative. This finding is in line with Ncube & Ndou's (2012) study which show that strict inflation targeting is incompatible with significant output growth, but a flexible inflation-targeting framework which places importance on real effective exchange rates leads to a significant real output growth. In fact, Taguchi and Kato's (2011) assessment of inflation targeting in some East-Asian economies shows that flexibility in exchange rate is a prerequisite to the success of inflation targeting regime.

Moreover, in interpreting the findings of this study, special emphasis is placed on the argument that inflation targeting cannot work well in emerging markets, like Nigeria, as emerging markets are deficient of the preconditions for a proper operation of inflation targeting. According to Kadioğlu et. al (2000), the prerequisites for the success of inflation targeting consist of sound economic structure; exchange rate flexibility; central bank independence; the institutional set-up; political commitment; a great deal of transparency and accountability of the Central Bank; absence of fiscal dominance; a single, clear inflation target; a sound inflation forecasting model; virile financial markets. The absence or inadequacy of these prerequisites may pose huge challenges for emerging markets like Nigeria trying to embrace inflation targeting. A lot of questions remain unanswered. The monetary authorities in Nigeria in Q4:2014 reacted to the sharp fall in crude oil price by increasing MPR by 100 basis points from 12% to 13% and moved the midpoint of the official exchange rate from N155/US$ to N167/US$.

These changes has nothing to do with difference between the forecast inflation rate and the target inflation. At best the changes are quick fix response to speculative activities in the foreign exchange market. These leaves us with more questions regarding the most appropriate method to track inflation targeting as the next best alternative for Nigeria. How can inflation targeting be best tested or modelled in a country where exchange rate is inflexible and interest rate is literarily managed by the monetary authority? Assuming the target inflation variable is known and the monetary authorities have a well-developed inflation forecasting model, what level should interest rate be raised for it to influence key macroeconomic variables? What is the interest rate threshold level that tracks inflation and economic growth in Nigeria? How effective is inflation targeting in a country that has a strong presence of fiscal dominance and institutional constraints? Perhaps, future studies will provide fuller answers to these questions.

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CURRENCY-ADJUSTED STOCK INDEX CAUSALITY AND COINTEGRATION: EVIDENCE FROM INTRADAY DATA

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ABSTRACT

Currency adjusted stock indices consider the impact of both stock value changes and underlying currency value changes on total wealth changes. This paper explores causality and cointegration of currency-adjusted indices using intraday data. This paper examines tick-by-tick data for seven currently available stock indexes, the Philadelphia Housing Index and the Dollar Index for the period 2002-2013. Results show cointegrating relationships between each combination of series examined. The analysis reveals a higher level of causality than found in previous research. The results show bidirectional Granger causality for every index pairwise combination examined.

JEL: F15, G11, D14

KEYWORDS: Cointegration, Stock Index, Currency-Adjusted Stock Index, Dow Jones Industrial Average

INTRODUCTION

Currency adjusted indexes control for two impacts on investor wealth. Standard stock indexes aggregate stock prices for many stocks into an index. Currency adjusted stock indexes simultaneously consider the impact of stock prices and underlying currency values. By considering both elements, currency adjusted indexes provide a measure of total portfolio value change. Many individuals live in one country but invest in another. Some authors suggest individuals do this to achieve international diversification benefits (Christoffersen, Errunza, Jacobs and Langlois, 2012; Berger, Pukthuanthong and Yang, 2011). Other motivations can drive this behavior as well. This investor behavior may be done for investment familiarity reasons, to retain an original country return option, or because of investment or tax regulations that limit the ability to relocate the funds or make such a relocation costly.

Investors living internationally but investing domestically, convert domestic investment earnings into domicile country currency for consumption. Two factors affect the purchasing power of these individuals. Purchasing power depends on performance of their investments and the exchange rate at which earnings convert to domicile country currency. Currency adjusted indices measure the combined effect of stock and currency changes on individual wealth.

Other individuals also face exposure to international markets. Individuals who use investment earnings to travel internationally, face exposure to both investment and currency risk. They wish to know the international purchasing power of their investments. Other individuals purchase items manufactured in non-domicile countries. These individuals desire to know the extent to which their domestic investments afford the international purchase. Currency adjusted indices more precisely measure the extent to which individuals achieve their goals.

This paper examines time-series properties of currency adjusted stock indices developed in Jalbert, 2012, 2014, 2015 and 2016. These earlier works make use of varying starting indexes, index observation frequencies, statistical techniques and cover different time periods. This work makes advances on the statistical sophistication of the analysis and examines a higher frequency dataset. The paper utilizes cointegration and Granger causality analyses to examine tick-by-tick trading data on a different series of indexes than examined in earlier studies. Several previous articles in this series examines close of day data. While close-of-day research is useful, news outlets regularly report index levels throughout the day. This paper provides an examination of stock indexes in a minute-by-minute setting thereby providing additional precision in the analysis.

The remainder of the paper begins with a review of the relevant literature. Next, the paper provides a description of the data and methodology used in the paper. The following section presents empirical results. The paper closes with some concluding comments.

LITERATURE REVIEW

Currency value adjusted stock indexes were first introduced by Jalbert (2012). He developed indices based on day-end closing values of existing stock indexes. He used United States Federal Reserve currency value data to adjust existing index values to reflect changes in the U.S. dollar value. Federal Reserve currency index data tracks U.S. dollar value against an international currency basket. His examination reveals large differences between original index and currency value adjusted index returns. For example, in 1981, the Standard and Poor's 500 Index produced a -10.237 percent return. However, the currency-adjusted version of the same index produced a -0.608 percent return. Similarly, in 2007, the Standard and Poor's 500 Index produced a 3.469 percent return, but the currency-adjusted version of the same index produced a -7.024 percent return. In as many as 15.49 percent of all daily returns, the original and currency-adjusted indices indicate different signs. Thus, one index indicates the market has increased but the other indicates the market has declined. He finds mixed evidence on index volatility. Currency adjusting indexes increased volatility in some instances but decreased volatility in other instances.

Jalbert (2014) utilizes the Dollar Index (DXY) as an alternative measure of U.S. dollar value. This index has advantages over the Federal Reserve indexes used in his earlier study. His analysis uses close-of-day data. He finds that currency value changes explain a larger portion of wealth changes than identified in Jalbert (2012). He finds currency level changes explain as much as 14.9 percent of wealth changes compared to 8.4 percent in Jalbert (2012). Jalbert (2015) examines intraday data from 2002-2013 including over one million tick-by-tick data observations for each index examined. His results show significant differences between the within-tick spreads for adjusted and raw indices. His results also show that Dollar Index changes explain more of total wealth changes than identified in his earlier works. Jalbert (2016) examined cointegration and causality between currency-adjusted indexes using close-of-day data. His results show cointegration exists between each pairwise combination of currency-adjusted indexes. His results further show that about half the pairwise combinations of currency-adjusted and raw indexes examined display bidirectional Granger causality.

The relationship between stock prices and exchange rates is relevant for the study at hand. Chien-Chung and Cheng-Few (2001) examine stock prices and exchange rates using data for G-7 countries. They find no long-run relationship between stock prices and exchange rates. They further find the relationship between stock prices and dollar values cannot be reliably used to predict future stock prices. Lin (2012) examines comovement between exchange rates and stock prices. She finds that comovement between stock prices and exchanges rates becomes stronger during crisis periods. Chkili and Nguyen (2014) examine linkages between exchange rates and stock market returns for BRICS (Brasil, Russia, India, China and South Africa). They find that stock market returns impact exchange rates, but exchange rates do not impact

stock market returns. Caporale, Hunter and Ali (2014) find univariate Granger causality from stock returns to echange rates in the U.S and U.K, but find an opposite effect for Canada.

The extant literature examines relationships between stock indexes of various exchanges and countries. Madaleno and Pinho (2012) examine correlation among four stock indexes of different countries. They find that geographically and economically closer markets display higher correlation and more short-run comovements. Wang, Podobnik, Horvatic and Stanley (2011) develop a model in which all index returns fluctuate in response to a single global factor. They find that a significant amount of index cross corelations are explainable by the global factor. Lee and Rui (2002) examine the relationship between trading volume, returns and volatility. They find trading volume does not Granger cause market returns. They further find that U.S. trading volume contains considerable predictive power for financial market variables of other countries.

Hammes and Wills (2005) examined the 1970's Oil Price Shock around the end of the Brenton Woods Agreement which implied fixed exchange rates. The end of the Brenton Woods Agreement resulted in a floating U.S. dollar value relative to other currencies. Oil prices were denominated in U.S. dollars before and after dissolution of the Brenton Woods Agreement. The authors argue these events, combined with a decline in dollar value, imply the 1970's oil price shock was a normal reaction. The shock simply maintained oil prices at approximately an equal amount of gold.

DATA AND METHODOLOGY

Data for seven U.S. Stock Indices, the Philadelphia Housing Index and the U.S. Dollar Index were obtained from Pi Trading (Pi Trading, 2013). Pi Trading provides open, high, low and close data for each minute throughout the trading day. Data for this study covers the period August 5, 2001 through April 5, 2013. Dollar Index data and stock index data were matched for each trading tick. The Dollar Index is based on U.S. dollar value relative to a basket of six major currencies. The dollar index started in March of 1973 with value of 100. Higher index levels imply a stronger dollar relative to March of 1973. The Dollar Index level displays considerable variation over time. The Index level ranges from a high of 164.72 in February 1985 to a low of 70.698 in March 2008. The index is widely quoted in the popular press. This research analyzes seven stock indices: The Dow Jones Industrial Average (DJIA), The NASDAQ Composite (COMPX), NASDAQ 100 (NDX), S&P 500 (SPX), S&P 400 (MID), Russell 3000 (RUA), and Russell 1000 (RUI). The paper also examines the Philadelphia Housing Index (HGX) which measures real estate values. The final dataset includes as many as 1,101,000 observations for each series.

Consider a stock index with level, RI_t at time *t.* Consider also the Dollar Index with base level of 100 and level, DI_t at time *t.* Then the currency adjusted stock index, CAI_t, at time *t* equals:

$$CAI_t = RI_t \, X \, \frac{DI_t}{100} \qquad (1)$$

Consider an unadjusted index with a level of 1,000 occurring at the same time the Dollar Index equals 110. Equation 1 produces an adjusted index level equaling 1,100. A Dollar index of 100 results in equal raw and adjusted indices. A Dollar Index level above 100, implies currency-adjusted index levels exceed raw index levels. Similarly, Dollar Index levels below 100 imply raw index levels exceed the currency-adjusted index level.

This paper examines tick-by-tick index levels, index changes and index returns. Consider a currency-adjusted stock index level at time t, CAI_t, and previous period level, CAI_{t-1}. Computation of index changes, and returns follow Equations 2 and 3 respectively:

$$Index\ Change\ =\ CAI_t - CAI_{t-1} \quad (2)$$

$$Index\ Return\ =\ ln(\frac{CAI_t}{CAI_{t-1}}) \quad (3)$$

The analysis involves calculating index changes and returns using closing data for each trading minute. The examination analyzes these figures using cointegration and Granger causality techniques.

RESULTS

Augmented Dickey-Fuller statistics (Dickey and Fuller 1979 & 1981) provide a method for examining time-series stationairity. The process tests for unit roots in the data. The analysis here allows for a maximum of 32 lags. Shwartz (1978) Information Criteria methodology is used to determine exact lag length. Augmented Dickey-Fuller results for index levels are presented in Table 1, Panel A. Level form results indicate a failure to reject the presence of a unit root for any of the stock series or the Philadelphia Housing series. It does indicate rejection of a unit root for the Dollar Index. The results show rejection of a unit root for each data series in first difference form.

Table 1: Unit Root Tests on Currency Adjusted Stock Indexes

	Level Form			First Differences		
Variable	**Lag Length**	**T-Statistic**	**Prob.**	**Lag Length**	**T-Statistic**	**Prob.**
Panel A: Index Levels						
Currency Adjusted DJIA	4	-2.1578	0.2221	3	-502.69	0.0001***
Currency Adjusted NASDAQ Composite	2	-1.4391	0.5646	1	-681.11	0.0001***
Currency Adjusted NASDAQ 100	4	-1.0379	0.7418	3	-512.05	0.0001***
Currency Adjusted S&P 500	2	-2.3057	0.1702	1	-672.80	0.0001***
Currency Adjusted S&P 400	6	-0.5456	0.8799	5	-383.22	0.0001***
Philadelphia Housing	6	-1.9525	0.3083	5	-382.57	0.0001***
Currency Adjusted Russell 3000	5	-1.8483	0.3573	4	-434.67	0.0001***
Currency Adjusted Russell 1000	5	-1.8434	0.3597	4	-436.96	0.0001***
Dollar Index	1	-3.0389	0.0314**	0	-1037.7	1.0000
Panel B: Index Returns						
Currency Adjusted DJIA	10	-311.24	0.0001***	121	-160.15	0.0001***
Currency Adjusted DJ Transport	1	-680.89	0.0001***	121	-159.01	0.0001***
Currency Adjusted DJ Utilities	6	389.60	0.0001***	121	-159.97	0.0001***
Currency Adjusted NASDAQ 100	3	-494.60	0.0001***	121	-160.98	0.0001***
Currency Adjusted NYSE Composite	5	-384.22	0.0001***	120	-153.42	0.0001***
Currency Adjusted S&P 500	14	-249.64	0.0001***	119	-144.36	0.0000***
Currency Adjusted Russell 3000	4	-435.73	0.0001***	120	-158.53	0.0001***
Currency Adjusted Russell 1000	3	-487.44	0.0001***	120	-155.87	0.0001***
Dollar Index	0	-1,036.6	1.000	121	-159.06	0.0001***

*This table shows Augmented Dickey-Fuller unit root test results on currency adjusted stock indexes and the Philadelphia Housing Index. *** indicates significance at the 1 percent level.*

Table 1, Panel B shows results for currency-adjusted index returns. The results indicate rejection of a unit root, for each series, in both level form and first difference form. To confirm result robustness, the analysis is completed with and without an intercept term. The results of the two methodologies are not discernably different. Phillips (1987) and Phillips and Perron (1988) developed an alternate test specification for detecting unit roots in data. Tests results using the Augmented Dickey Fuller test and the Phillips and Perron test produce similar results, further attesting to the robustness of the findings. Based on these results, the remaining analysis examines index level changes and index returns without adjustment.

Table 2: Johanson Cointegration Analysis on Currency Value Adjusted Index Level Changes

Index 1	Index 2	Hypothesized Relations	Eigenvalue	Trace Statistic	P-Value
DJ Industrials	NASDAQ Comp	None	0.1930	400,820	1.0000
		At Most 1	0.1591	179,108	0.0000***
DJ Industrials	S&P 400	None	0.2018	388,610	1.0000
		At Most 1	0.1429	157,855	0.0000***
DJ Industrials	NASDAQ 100	None	0.1883	400,654	1.0000
		At Most 1	0.1638	184,924	0.0000***
DJ Industrials	Phil. Housing	None	0.1762	347,839	1.0000
		At Most 1	0.1507	159,068	0.0000***
DJ Industrials	Russell 3000	None	0.2141	419,059	1.0000
		At Most 1	0.1524	170,524	0.0000***
DJ Industrials	Russell 1000	None	0.2188	427,900	1.0000
		At Most 1	0.1549	173,444	0.0000***
DJ Industrials	S&P 500	None	0.2271	444,258	1.0000
		At Most 1	0.1579	177,766	0.0000***
NASDAQ Comp	S&P 400	None	0.2095	399,809	1.0000
		At Most 1	0.1440	159,134	0.0000***
NASDAQ Comp	NASDAQ 100	None	0.1996	396,156	1.0000
		At Most 1	0.1482	165,893	0.0000***
NASDAQ Comp	Phil. Housing	None	0.1711	342,838	1.0000
		At Most 1	0.1515	160,078	0.0000***
NASDAQ Comp	Russell 3000	None	0.2021	405,937	1.0000
		At Most 1	0.1544	173,063	0.0000***
NASDAQ Comp	Russell 1000	None	0.1989	402,745	1.0000
		At Most 1	0.1555	174,205	0.0000***
NASDAQ Comp	S&P 500	None	0.1981	405,629	1.0000
		At Most 1	0.1575	177,245	0.0000***
S&P 400	NASDAQ 100	None	0.2048	393,813	1.0000
		At Most 1	0.1441	159,214	0.0000***
S&P 400	Phil. Housing	None	0.1621	323,013	1.0000
		At Most 1	0.1460	152,282	0.0000***
S&P 400	Russell 3000	None	0.2015	387,983	1.0000
		At Most 1	0.1434	158,079	0.0000***
S&P 400	Russell 1000	None	0.1996	384,458	1.0000
		At Most 1	0.1428	157,251	0.0000***
S&P 400	S&P 500	None	0.1970	383,222	1.0000
		At Most 1	0.1434	158,476	0.0000***
NASDAQ 100	Phil. Housing	None	0.1768	349,477	1.0000
		At Most 1	0.1514	159,954	0.0000***
NASDAQ 100	Russell 3000	None	0.1954	396,749	1.0000
		At Most 1	0.1540	172,480	0.0000***
NASDAQ 100	Russell 1000	None	0.1937	396,362	1.0000
		At Most 1	0.1557	174,487	0.0000***
NASDAQ 100	S&P 500	None	0.1935	401,569	1.0000
		At Most 1	0.1590	179,175	0.0000***
Phil. Housing	Russell 3000	None	0.1647	335,197	1.0000
		At Most 1	0.1517	160,088	0.0000***
Phil. Housing	Russell 1000	None	0.1661	336,165	1.0000
		At Most 1	0.1515	159,642	0.0000***
Phil. Housing	S&P 500	None	0.1703	341,784	1.0000
		At Most 1	0.1513	159,898	0.0000***
Russell 3000	Russell 1000	None	0.2778	506,729	1.0000
		At Most 1	0.1532	171,348	0.0000***
Russell 3000	S&P 500	None	0.2614	484,587	1.0000
		At Most 1	0.1534	171,894	0.0000***
Russell 1000	S&P 500	None	0.3020	545,842	1.0000
		At Most 1	0.1562	175,086	0.0000***

*This table shows Johanson cointegration test results for currency-adjusted index changes. The first and second columns indicate the two series under examination. The third column indicates the hypothesized number of cointegrating relations. The fourth column shows the ordered eigenvalues. The fifth and sixth columns show the Trace statistic and P-value respectively. ***, ** and * indicate significance at the one, five and ten percent levels respectively.*

Table 3: Johanson Cointegration Analysis on Currency Adjusted Index Returns

Index 1	Index 2	Hypothesized Relations	Eigenvalue	Trace Statistic	P-Value
DJ Industrials	NASDAQ Comp	None	0.1997	409,782	1.0000
		At Most 1	0.1594	179,501	0.0000***
DJ Industrials	S&P 400	None	0.2129	403,516	1.0000
		At Most 1	0.1434	158,444	0.0000***
DJ Industrials	NASDAQ 100	None	0.1934	407,865	1.0000
		At Most 1	0.1643	185,614	0.0000***
DJ Industrials	Phil. Housing	None	0.1784	354,288	1.0000
		At Most 1	0.1539	162,838	0.0000***
DJ Industrials	Russell 3000	None	0.2224	430,769	1.0000
		At Most 1	0.1530	171,269	0.0000***
DJ Industrials	Russell 1000	None	0.2264	438,634	1.0000
		At Most 1	0.1554	174,117	0.0000***
DJ Industrials	S&P 500	None	0.2322	451,768	1.0000
		At Most 1	0.1585	178,457	0.0000***
NASDAQ Comp	S&P 400	None	0.2141	406,640	1.0000
		At Most 1	0.1448	160,054	0.0000***
NASDAQ Comp	NASDAQ 100	None	0.2007	398,890	1.0000
		At Most 1	0.1492	167,166	0.0000***
NASDAQ Comp	Phil. Housing	None	0.1760	351,581	1.0000
		At Most 1	0.1541	163,042	0.0000***
NASDAQ Comp	Russell 3000	None	0.2062	412,015	1.0000
		At Most 1	0.1550	173,807	0.0000***
NASDAQ Comp	Russell 1000	None	0.2033	409,073	1.0000
		At Most 1	0.1560	174,787	0.0000***
NASDAQ Comp	S&P 500	None	0.2034	413,083	1.0000
		At Most 1	0.1580	177,914	0.0000***
S&P 400	NASDAQ 100	None	0.2074	398,135	1.0000
		At Most 1	0.1449	160,193	0.0000***
S&P 400	Phil. Housing	None	0.1695	333,034	1.0000
		At Most 1	0.1473	153,768	0.0000***
S&P 400	Russell 3000	None	0.2140	404,819	1.0000
		At Most 1	0.1440	158,799	0.0000***
S&P 400	Russell 1000	None	0.2115	400,445	1.0000
		At Most 1	0.1434	157,967	0.0000***
S&P 400	S&P 500	None	0.2094	399,882	1.0000
		At Most 1	0.1440	159,230	0.0000***
NASDAQ 100	Phil. Housing	None	0.1802	357,335	1.0000
		At Most 1	0.1547	163,781	0.0000***
NASDAQ 100	Russell 3000	None	0.1978	401,094	1.0000
		At Most 1	0.1549	173,679	0.0000***
NASDAQ 100	Russell 1000	None	0.1997	401,016	1.0000
		At Most 1	0.1565	175,447	0.0000***
NASDAQ 100	S&P 500	None	0.1967	406,841	1.0000
		At Most 1	0.1600	180,303	0.0000***
Phil. Housing	Russell 3000	None	0.1684	343,025	1.0000
		At Most 1	0.1548	163,605	0.0000***
Phil. Housing	Russell 1000	None	0.1692	343,492	1.0000
		At Most 1	0.1548	163,395	0.0000***
Phil. Housing	S&P 500	None	0.1731	349,188	1.0000
		At Most 1	0.1548	163,907	0.0000***
Russell 3000	Russell 1000	None	0.3002	540,264	1.0000
		At Most 1	0.1540	172,380	0.0000***
Russell 3000	S&P 500	None	0.2809	513,145	1.0000
		At Most 1	0.1541	172,758	0.0000***
Russell 1000	S&P 500	None	0.3166	568,357	1.0000
		At Most 1	0.1568	175,826	0.0000***

*This table shows Johanson cointegration test results for currency-adjusted index returns. The first and second columns indicate the two series under examination. The third column indicates the hypothesized number of cointegrating relations. The fourth column shows the ordered eigenvalues. The fifth and sixth columns show the Trace statistic and P-value respectively. *** indicates significance at the one percent level.*

This paper conducts Johanson cointegration analysis (Johanson, 1991). The methodology involves pairwise cointegration tests of each two-index pairwise currency adjusted index combination. The test specification incorporates EViews software options of intercept, trend in CE intercept in VAR and Lag intervals equaling 1 4. The analysis uses MacKinnon, Haug and Michelis (1999) p-values. Table 2 shows results for tick-by-tick currency-adjusted index level changes. The results reveal two cointegrating equations for each index combination. For each index combination the data rejects the test for at most one cointegrating relationship. Table 3 shows results of cointegration analysis on currency adjusted tick-by-tick index returns. The results here also show two cointegrating equations for each pairwise combination of indexes.

The next analysis involves conducting Granger causality tests between pairwise combinations of adjusted indices (Granger 1969). The test specification permits up to ten lags. Table 4, shows results of tests on currency adjusted index level changes and returns. The tests determine if the index listed in column, Index 1, Granger causes the index listed in column, Index 2. The analysis involved conducting fifty-six pairwise tests. The results show bi-directional causality for each index pair. These findings differ from those of Jalbert (2016), who found, using daily data and a different combination of indexes, that about half of all index combinations display bi-directional causality. The combined results show more pronounced elements of Granger causality in higher frequency series. This causality exists in both index-level changes and index returns.

Table 4: Granger Causality Test Results on Currency Adjusted Stock Indexes

Index 1	Index 2	Index Changes	Index Returns	Index 1	Index 2	Index Changes	Index Returns
		F Statistic	F Statistic			F Statistic	F Statistic
NASDAQ Cp	DJ Industrials	341.30***	433.37***	Phil. Housing	S&P 400	4,749.02***	6,567.20***
DJ Industrials	NASDAQ Cp	687.02***	886.09***	S&P 400	Phil. Housing	91.23***	48.05***
S&P 400	DJ Industrials	101.46***	63.35***	Russell 3000	S&P 400	7,009.48***	8,116.25***
DJ Industrials	S&P 400	9,167.36***	11,086.3***	S&P 400	Russell 3000	319.42***	267.50***
NASDAQ 100	DJ Industrials	381.20***	485.71***	Russell 1000	S&P 400	6,599.67***	7,712.67***
DJ Industrials	NASDAQ 100	459.43***	577.03***	S&P 400	Russell 1000	280.02***	215.47***
Phil. Housing	DJ Industrials	46.89***	62.14***	S&P 500	S&P 400	5,463.24***	6,785.80***
DJ Industrials	Phil. Housing	548.66***	712.20***	S&P 400	S&P 500	256.32***	182.00***
Russell 3000	DJ Industrials	16.50***	9.953***	Phil. Housing	NASDAQ 100	101.34***	111.07***
DJ Industrials	Russell 3000	4,410.22***	5,011.87***	NASDAQ 100	Phil. Housing	303.41***	475.97***
Russell 1000	DJ Industrials	25.61***	16.55***	Russell 3000	NASDAQ 100	31.39***	44.95***
DJ Industrials	Russell 1000	4657.36***	5,291.46***	NASDAQ 100	Russell 3000	1426.45***	1,520.70***
S&P 500	DJ Industrials	43.70***	44.19***	Russell 1000	NASDAQ 100	22.72***	28.26***
DJ Industrials	S&P 500	5,738.73***	6,101.28***	NASDAQ 100	Russell 1000	1,381.03***	1,511.11***
S&P 400	NASDAQ Cp	39.36***	25.06***	S&P 500	NASDAQ 100	13.83***	22.86***
NASDAQ Cp	S&P 400	5,616.26***	5,825.10***	NASDAQ 100	S&P 500	1,830.83***	19,33.36***
NASDAQ 100	NASDAQ Cp	250.78***	288.91***	Russell 3000	Phil. Housing	86.88***	183.13***
NASDAQ Cp	NASDAQ 100	19.70***	28.92***	Phil. Housing	Russell 3000	246.68***	385.35***
Phil. Housing	NASDAQ Cp	204.72***	247.88***	Russell 1000	Phil. Housing	95.07***	176.45***
NASDAQ Cp	Phil. Housing	246.16***	426.49***	Phil. Housing	Russell 1000	212.47***	355.82***
Russell 3000	NASDAQ Cp	91.75***	133.35***	S&P 500	Phil. Housing	73.60***	154.42***
NASDAQ Cp	Russell 3000	1,384.11***	1,523.54***	Phil. Housing	S&P 500	252.88***	327.12***
Russell 1000	NASDAQ Cp	74.50***	100.77***	Russell 1000	Russell 3000	312.50***	334.58***
NASDAQ Cp	Russell 1000	1276.06***	1,462.80***	Russell 3000	Russell 1000	265.13***	441.15***
S&P 500	NASDAQ Cp	45.48***	77.53***	S&P 500	Russell 1000	224.47***	313.10***
NASDAQ Cp	S&P 500	1,635.04***	1,804.79***	Russell 1000	S&P 500	1,087.25***	1,167.34***
NASDAQ 100	S&P 400	5,675.94***	5,704.49***	S&P 500	Russell 3000	210.98***	319.44***
S&P 400	NASDAQ 100	41.069***	33.70***	Russell 3000	S&P 500	1,319.91***	1.245.54***

*This table shows Granger causality test results. The table shows results for each pairwise combination of currency value adjusted indexes. The results labeled Index Changes show the results for tests on index level changes as specified in Equation 2. Results labeled Index Returns shows results of tests based on Equation 3. ***, ** and * indicate significance at the one, five and ten percent levels respectively.*

CONCLUDING COMMENTS

This paper examines currency-adjusted stock indexes. The construction of these indexes begins with an existing index and adjusts it to reflect changes in underlying currency value. The methodology involves

determining the U.S. dollar value as compared to a basket of six currencies using the DXY Dollar Index. This paper extends the analysis by using cointegration and Granger causality statistical techniques and by using tick-by-tick trading data. This paper examines tick-by-tick data for seven stock indexes, the Philadelphia Housing Index and the Dollar Index from 2001-2013 including more than 1,000,000 observations for each series.

In level form, for all series examined, the data fails to reject the presence of a unit root. But, in first difference form, the data rejects a unit root for each series. In contrast, the data rejects the presence of a unit root for each index return series without adjustment. The results show cointegrating relationships between pairwise index combinations of index changes and returns. Granger causality tests show bi-directional causality between each pairwise combination of indexes.

Currency value adjusting stock indexes and other asset values provides opportunities for additional research. Examining index arbitrage opportunities created by currency value adjustments might lead to interesting insights. In addition, currency adjusted examinations of real estate values may provide interesting insights.

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ON THE QUANTITY THEORY OF MONEY, CREDIT, AND SEIGNIORAGE

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ABSTRACT

This paper argues the predictive power of the sectoral approach towards a quantity theory of credit is weak. A quantity theory of commercial-bank-seigniorage approach is proposed in its place. It suggests that the financial system may be held responsible for price and output fluctuations to the extent commercial bank seigniorage alters the stock of money in circulation. If not, the financial sector can become the source of instability by influencing profitability in the real sector through a Goodwin-type interaction. These trends could be countered by an interest rate rule based on deposit habits and on the deposit rate, and supplemented perhaps by a policy of influencing these habits and manipulating the deposit rate.

JEL: E3, E4, E5

KEYWORDS: Quantity Theory, Credit Theory, Commercial Bank Seigniorage, Instability

INTRODUCTION

Commercial banks enjoy the privilege of using their liabilities as a form of money; their lending decisions based on private interest affect in the aggregate the stock of money and thus social welfare. To maximize profit, banks tend to minimize their level of deposit reserve holdings, loaning out as much of their funds as possible at the market rate of interest. This has provoked repeatedly market instability and fluctuations in the money supply as documented, for example, by Friedman and Schwartz (1963) for the United States. Banks create money, "out of thin air" as the saying goes; money, chasing a given volume goods, decreasing afterwards individuals' buying power, no differently than what would have been done with forged money.

The Austrian School of Economics claims that this is a win-win state of affairs for the banks and the government: The government tolerates the lending practices of banks in exchange for banks' commitments to buy debt from the government to fund government's ambitions (Myers, 2012). But, according to anthropologist Graeber (2011), for most of human history, money has been widely understood to represent debt. So, this is the environment in which the institution of commercial banking had to develop. The point is that nowadays most of the money in circulation is credit money. It is income earned by banks by expanding lending under fractional reserve banking, and labelled through the term "commercial bank seigniorage"; £ 50 billion in the UK, and about € 240 billion in the Euro area in 2013: "At 82–97% (depending on country and monetary aggregate) bank money today represents the lion's share of the money supply in public circulation." (Huber, 2014, 1).

And, as already noted, credit is powerful in disturbing economic activity at least in the short- and medium-run (see also Iacoviello, 2015). Theoretically, the credit or debt approaches to money that have been advanced fall into a broader category of work which postulates that monetary creation is endogenous (Wray, 2004). Their origin is traced by Schumpeter (1954) back to Plato, and their modern formulation owes to Macleod (1889) and Mitchell-Innes (1914). Recently, Professor Werner (1992, 1997, 2009, 2012) tried to

make the credit viewpoint of money compatible with the quantity theory of money. The purpose of this paper is to look at this attempt closer and take it a step further if possible.

LITERATURE REVIEW

Professor Werner has been recommending a quantity theory of credit as a means of answering questions about *inter alia* defining money, the declining velocity, and financial crises. Fisher's (1911) equation of exchange is broken down into one part referring to the real sector of the economy, and into a second part concerning the financial sector, as proposed originally by Keynes (1930). Werner is doing so focusing in essence on the propagation of the business cycle by the financial system, which was a matter of similar concern by the early Chicago version of the quantity theory of money, (see e.g. Simons, 1936). And, his approach has been found to have enough explanatory power.

Nevertheless, although breaking down the equation of exchange as proposed not only by Keynes but by Fisher himself too, can be empirically important, such a break-down is shown herein to be of limited predictability. The correctness, for instance, of the quantitative easing prescribed by Werner against the current recession, does not suggest predictive power since, this prescription follows directly from the current conditions of insufficient demand (see e.g. Soldatos, 2015): Give people money to spend and spur recovery.

From the viewpoint of the quantity theory of money, the real issue is the decomposition between central and commercial bank money rather than between a real and a monetary sector. The real issue is what happens within the monetary sector given the "Classical Dichotomy" but also given the correct version of this dichotomy; the version which distinguishes between the invariance of real variables to money changes and the invariance in the demand of goods: To hold the "demand-invariance" view is shown to be simply wrong (Hahn, 2002).

What happens within the monetary sector does affect the demand for goods. And, from the viewpoint of the credit theory of money, what happens depends mostly on commercial bank money. So, a successful marriage between the quantity theory of money and the debt theory of it requires the development of a quantity theory of commercial bank seigniorage. This task is undertaken in the next Section. Section 3 concludes this work by putting the topic under investigation into further theoretical perspective.

A QUANTITY THEORY OF SEIGNIORAGE

In this section, we first show the limited predictability entailed by Werner's Sectoral Quantity Theory, we then proceed with a hypothesis based on central and commercial bank seigniorage decomposition, and we finish with some policy considerations.

The Sectoral Quantity Theory

Let total nominal transactions, Y, be the sum of the transactions Y_1 and Y_2 in sectors 1 and 2 of the economy, respectively:

$$Y \equiv Y_1 + Y_2 \quad (1)$$

And, let the total stock of money, M, consist of the stocks employed in these two sectors:

$$M \equiv M_1 + M_2 \quad (2)$$

Following Keynes (1930), the overall velocity of the economy, $V = Y/M$, is the sum of the true, sectoral velocities, $V_1 = Y_1/M_1$ and $V_2 = Y_2/M_2$, with the former velocity being weighted by $m = M_1/M$ and the latter by $1 - m = M_2/M$:

$$V = mV_1 + (1 - m)V_2 \quad (3)$$

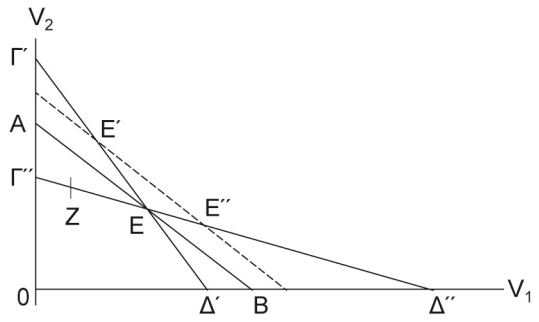
The total differential of (3) is: $dV = mdV_1 + (1 - m)dV_2$. Setting $dV = 0$ and solving for dV_2/dV_1, one obtains that:

$$dV_2/dV_1 = -m/(1 - m) = -M_1/M_2 \quad (4)$$

which is depicted as the "iso-velocity" line AB in the space $V_1 - V_2$ of Figure 1; the slope of this line is:

$\varphi \equiv -m/(1 - m) = -M_1/M_2$.

Figure 1: The Sectoral Decomposition Diagrammatically



Next, substituting the sectoral velocities, $V_1 = Y_1/M_1$ and $V_2 = Y_2/M_2$, in (2) yields:

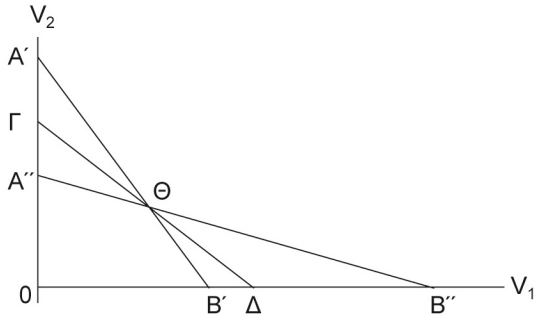
$$M = (Y_1/V_1) + (Y_2/V_2)$$

whose total differential gives when $dM = 0$ that:

$$dV_2/dV_1 = -(1 - y)/y = -Y_2/Y_1 \quad (5)$$

where $y = Y_1/Y$ and $1 - y = Y_2/Y$. The "iso-money" line $\Gamma\Delta$, having slope $\psi \equiv -(1 - y)/y = -Y_2/Y_1$, is thus obtained in Figure 1. It may coincide with AB, or cut it from above, as $\Gamma'\Delta'$ does, or below, as $\Gamma''\Delta''$ does. Two points need to be made now: Firstly, given $\Gamma\Delta$, its intersection point with AB, point E, is the point at which both $dV = 0$ and $dM = 0$. Suppose, for instance, that AB, $\Gamma'\Delta'$, and Z is the case. Z is on a $\Gamma\Delta$ below $\Gamma'\Delta'$; that is, given the V connected with AB, the stock of M required to sustain the volume of transactions associated with $\Gamma'\Delta'$ is too small. Increasing this stock, the iso-money line passing through Z will start shifting upward, sliding from Z to E. The second point needed to be made is that ψ gives the sectoral composition of Y, and that a single φ can be consistent with many ψ's as Figure 1 illustrates by having $\Gamma'\Delta'$ and $\Gamma''\Delta''$ passing through the same point E on AB. Figure 2 illustrates that a single ψ can be consistent with various φ's, too.

Figure 2: The Vagueness of the Sectoral Decomposition



Therefore, a change in the sectoral velocities with unchanged the overall one, V, i.e. a pivot of AB centered at E in Figure 1, provides no information about a would-be change in the sectoral composition and volume of total transactions. Similarly, a pivot of $\Gamma\Delta$ centered at Θ in Figure 2, offers no insight as to would-be changes in V and in sectoral velocities. Also, a change, an increase, say of V given its sectoral composition, i.e. a parallel shift of AB upwards, cutting $\Gamma'\Delta'$ and $\Gamma''\Delta''$ at E' and E'', respectively, can be taken to mean one out of four things. Given the composition of transactions, it might mean either that $\Gamma'\Delta'$ has shifted to the right passing now through E'', or that $\Gamma''\Delta''$ has shifted upwards passing now through E'. And, given the overall volume of transactions, either that its composition has changed from that described by $\Gamma\Delta$ to the one captured by $\Gamma''\Delta''$, or the opposite. These are the four eventualities in case of a shift in AB including a change in slope as well. Analogous remarks can be made about the shifts of $\Gamma\Delta$ given AB in Figure 2, all pointing to the conclusion that without guidance from empirical observation towards the identification of that "eventuality" which is empirically relevant,… anything goes.

The Seigniorage Decomposition Quantity Theory

There do exist a few empirical investigations of sectoral velocities, notably by Selden (1961), McGouldrick (1962), Garvey and Blyn (1969), and Ireland (1991). To our knowledge, they are the only ones, and they too reflect implicitly or explicitly that the ultimate concern is the financial instability coming out of commercial bank seigniorage, in the spirit always of the early Chicago tradition. Indeed, the nexus between such seigniorage and real economic activity is what prompts in the first place the interest in disaggregation, in a quantity theory of credit. But, disaggregation, differentiation of the source of instability, is one thing, and decomposition, the explicit sectoral modeling, another. The focus is on the interaction between these sources and not on the sources *per se*. Towards this direction, note that the distinction between central and commercial bank seigniorage does reflect such a source differentiation. Consequently, once a quantity theory of seigniorage is advanced, once central and commercial bank seigniorage are incorporated directly in the equation of exchange as a sum, disaggregation becomes built in this sum. It is an approach to instability, which will not have to cope eventually with the vagueness characterizing sectoral modeling results as follows:

In the absence of commercial banks or the same, under a 100% reserve system, and central bank only seigniorage, S, we have that, $S = i(H/P)$, where H is the monetary base, P is the price level, and i is the nominal interest rate. From the real-sector quantity-theory equation, $H = kPQ$ and hence, $S = i(kPQ/P) \Rightarrow S = ikQ$, where Q is real income and $1/k = V$ is the velocity of circulation. In the presence of a commercial banking system benefiting from commercial bank seigniorage, $\mathcal{S} = [i(1-\rho) - r_n]D/P$, under a required reserve ratio $\rho \leq 1$ on deposited money $D/P = \lambda Q$ and under a deposit rate r_n, total seigniorage is the sum $S + \mathcal{S} = (ik - \lambda r_n)Q$, given that $H = F + D$, where F is cash (see e,g. Baltensperger and Jordan, 1997). Now, note that what banks do mostly is producing bank money out of their own bank

money given a token of D. That is, $P\mathcal{S}$ is the output and money of the banking system, or in terms of the sectoral notation above, Y_2 and M_2, implying that $V_2 = 1$, $dV_2 = 0$, and $dV = dV_1$: An additional unit of bank money can be produced instead of rotating an already existing one, and any observed change in velocity comes out not from change in the financial sector *per se*, but from the impact of financial change on the real sector.

So, letting $M \equiv S + \mathcal{S} = (ik - \lambda r_n)Q$ the overall quantity equation, $MV = PQ$ becomes:
$(ik - \lambda r_n)QV = PQ \Rightarrow [i(1/V) - \lambda r_n]V = P \Rightarrow (i - V\lambda r_n) = P \Rightarrow i = P + V\lambda r_n$

which in conjunction with the definition that $i = r + \pi$, implies that: $P + V\lambda r_n = r + \pi \Rightarrow$

$$V = (r + \pi - P)/\lambda r_n \tag{6}$$

where r is the real rate of interest and π is the inflation rate. Setting $dV = 0$ in the total differential of (6), yields that: $dr + d\pi - dP = (r + \pi - P)[(1/\lambda)d\lambda + (1 + r_n)dr_n]$. In view of this expression and given that from (6), $r + \pi - P = V\lambda r_n$, the velocity, V^*, which is consistent with $dV = 0$ is:

$$V^* = (dr + d\pi - dP)/(r_n d\lambda + \lambda dr_n) \tag{7}$$

Noting that $\pi = (P - P_{-1})/P_{-1}$ and $\pi_{-1} = (P_{-1} - P_{-2})/P_{-2}$, (7) may be rewritten as follows:

$$PP_{-2} - P_{-1}^2 + P_{-1}P_{-2}[dr - V^*(r_n d\lambda + \lambda dr_n)] = 0 \tag{8}$$

where subscripts "-1" and "-2" denote time lags. We have clearly a complicated cubic equation in prices, capturing price instability with constant V at V^*. Setting $P = P_{-1} = P_{-2}$ in (8), which would be the case of price stability, the condition that: $dr - V^*(r_n d\lambda + \lambda dr_n) = 0$ has to be satisfied. It is a necessary but not sufficient condition, because zeroing the bracketed term in (8) gives that $PP_{-2} - P_{-1}^2 = 0$, which again is a necessary but not sufficient condition by itself for price stability. Stability presupposes the satisfaction of both conditions. Setting for simplicity but plausibly $P_{-2} = 1$ in (8), the latter condition becomes: $P = P_{-1}^2$; it is a condition for geometric price reduction.

Next, setting $dM = 0$ and $dV = 0$ in the total differential of $M = [i(1/V) - \lambda r_n]Q$, one obtains that $Qdi + idQ = V^*(r_n d\lambda + \lambda dr_n)Q + V^*\lambda r_n dQ$, or letting $dQ = Q - Q_{-1}$ and $Q/Q_{-1} = 1 + g$:

$$Q\{(1 + g)[i + di - V^*(r_n d\lambda + \lambda dr_n) - V^*\lambda r_n] - (i - V^*\lambda r_n)\} = 0 \tag{9}$$

According to (9), given M and V, the composition of M alone does provoke output instability too, as expected, because keeping M and V constant, price changes should be offset by output changes. This can be seen by solving (8) for $V^*(r_n d\lambda + \lambda dr_n)$ and inserting the resulting expression of P's in (9). Note that under price stability, that is, under $dr - V^*(r_n d\lambda + \lambda dr_n) = 0$ and $d\pi = 0$, and given that $di = dr + d\pi$, (9) becomes: $gQ(i - V^*\lambda r_n) = 0$, which implies that steady growth under price stability presupposes an i equaling to $V^*\lambda r_n$. This interest rate rule that $i = V^*\lambda r_n$, takes the place of the condition of geometric price decline, needed for price stability beyond the condition that $dr - V^*(r_n d\lambda + \lambda dr_n) = 0$. V, λ, and r_n are not quantities that change every day. V and λ reflect *inter alia* consumer habits while r_n changes sporadically. Consequently, in practice, the interest rate rule is really one about interest rate stability; price stability would be indeed corroborated by interest rate stability. This is what the condition $dr - V^*(r_n d\lambda + \lambda dr_n) = 0$ is about too, because if in practice $d\lambda = 0$ and $dr_n = 0$, this condition amounts to $dr = o$, which when coupled with $d\pi = 0$, implies that $di = dr + d\pi = 0 + 0 = 0$. It appears that in practice, one should be the policy rule, namely that $di = 0$.

It also appears that the stricter, Friedman's rule that $i = 0$ would apply only if $r_n = 0$; both S and $\mathcal{S}$ would be zero in this case. This of course would be in the spirit under which this rule was advanced initially assuming away commercial banking: namely, elimination of central bank seigniorage. Incorporating commercial banks into the discussion, the elimination of commercial bank seigniorage too, would come as the natural extension of Friedman's rule. But, would such an extended Friedman rule be sensible policy-wise beyond the weaker requirement that $di = 0$? According to (8) and (9), failure to abide by this weaker rule does produce price and output instability even if M and V are held constant; but much more so when M changes since $M = (ik - \lambda r_n)Q$. And, responsible for this change should be held the financial system given that the bulk of M is easily expandable commercial bank seigniorage. It seems to the authors that under these circumstances the extended Friedman rule should be a must if of course it was decided to be followed. Nevertheless, the *k*-percent rule advanced by Friedman (1960) too, is more practical and a policy of $di = 0$ should be seen as its natural companion.

Instability and Monetary Policy

Bank money has own life for which discretionary monetary policy cannot do much. For example, a policy keeping M constant under a Goodwin-type interaction between bank and firm profitability, would be ineffective as follows: Total firm revenues are $PQ \equiv Y \equiv MV$, and are made possible through lending Π_b to pay capital and labor expenses in such a manner that:

$$d\Pi_f/dt = \Pi_f(Y - \beta\Pi_b) \tag{10}$$

where f and b designate the firms and the bank sector, respectively, t is time, β is some positive constant reflecting the availability of lending, and Π captures profits. That is, the rate of change of the profit of firms depends on whether lending is enough to keep sustaining a given volume of transactions. Variations in lending that leave Y constant by manipulating prices and quantities on the part of firms, become fully mapped into profit variations.

The profit of the banking sector is the bank money loaned to the firms, and based on $D = \zeta d\Pi_f$, i.e. $(i - r_n)\zeta d\Pi_f$, in such a manner that: $(i - r_n)\zeta d\Pi_f = \beta\Pi_b + \xi$, and that:

$$d\Pi_b/dt = \Pi_b\left[(i - r_n)\zeta d\Pi_f - \xi\right] \tag{11}$$

The term $\xi = (1 - \beta)\Pi_b$ is a constant, reflecting normal profit by the banks, covering their opportunity cost; it is the profit that would keep them in operation under the worst of circumstances. In effect, bank money, bank lending, and commercial bank seigniorage become synonymous to supernormal bank profit. Equations (10) and (11) are Lotka-Volterra ones, having firms being the prey of predating banks for bank profit beyond the normal one. The two critical points for stationariness are: $\Pi_f^* = \xi/(i - r_n)\zeta$ and $\Pi_b^* = Y/\beta$ in connection with a cycle of period equal to $2pi/\sqrt{Y\xi}$ and with the cycle of the profit of firms leading by (1/4)th of this period, where pi=3.14159… Now, having S changing over the cycle to be keeping M constant would not alter this course of things given that $(i - r_n)\zeta d\Pi_f = \mathcal{S}$. Such a policy would be useful only to the extent that V does not respond adequately to the cycle as it did happen with the Great Crash in 1929.

CONCLUDING COMMENTS

To sum up, Werner's sectoral approach towards a quantity theory of credit is too vague in its predictions. A quantity theory of commercial bank seigniorage approach is proposed in its place, arriving at the conclusion that the financial system may be held responsible for price and output fluctuations to the extent

commercial bank seigniorage alters the stock of money in circulation considerably. If not, the financial sector can become the source of instability by affecting profitability in the real sector through a Goodwin-type interaction. These trends could be countered by an interest rate rule based on deposit habits and on the deposit rate, and supplemented perhaps by a policy of influencing these habits and manipulating the deposit rate.

Many will have noticed a contradiction in the term "quantity theory of (commercial bank) seigniorage". The quantity theory of money wants money to be exogenous, with independent causal role, whereas commercial bank seigniorage refers to money endogeneity. Such has been the standard doctrinal standpoint. Thirlwall (1999) distinguishes three versions of the quantity theory: the equation of exchange, the income quantity theory, and the cash balance equation. But, none of these is actually a theory. They are all equations, alternative expressions of a single analytical tool adapted to the needs of the particular hypothesis explored each time. A single tool, when money is seen from the viewpoint of financial products within the accounting structure of a banking firm; money defined according to flows of income and expense and stocks of assets and liabilities as, for instance, Fixler and Zieschang (1998) do.

And, a tool does not form a hypothesis, a theory by itself; it helps instead establish one. Keynes and monetarism, for example, wanted it to be a hypothesis *per se*, but Old Chicago used it to rationalize "A Program for Monetary Reform" (see e.g. Douglas et al., 1939). The source of the confusion has been Hume's (1987 [1742], II.IV.2) view that: "Were all the gold in England annihilated at once, and one and twenty shillings substituted in the place of every guinea, would money be more plentiful or interest lower? No surely: We should only use silver instead of gold." As Diaz-Gimenez and Kirkby (2013, 2) note, the keyword in this expression is "at once", which only in the long-run is expected to hold as a reflection of an average trend: "When central banks conduct monetary policy, changes in the quantity of money are not introduced evenly and 'at once'… [since] money is injected into… typically the banking system, and it spreads out gradually from there." And, as a matter of fact, there can be no short- and medium-run theory of money at all, since: "Money is means to effect transactions and savings; what will be used as money depends upon such a miscellany of factors that no 'theory' can be expected to emerge" (Rashid, 2001, 1).

Finally, note that the view of commercial bank seigniorage employed herein is in line with the Fixler-Zieschang accounting approach to the definition of credit money: Money stemming from the microeconomic theory of financial firms and of household consumption of financial asset services, by taking "the total sales of financial institutions to be the net interest income on "produced" asset and liability products" (Fixler and Zieschang, 1998, 1): This consistent with any of the three quantity equations.

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FINANCIAL RISK AND ISLAMIC BANKS' PERFORMANCE IN THE GULF COOPERATION COUNCIL COUNTRIES

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ABSTRACT

This study examines the relationship between financial risk and performance of Gulf Cooperation Council Islamic banks and the relative importance of the most common types of risk. The study covers 11 of the 47 Islamic banks of the Gulf Cooperation Council region from 2000 to 2012, based on the availability of data. Data were obtained from the Bankscope database. For bank performance, the two most common measures, ROA and ROE, were alternatively used and for risk measures. Four types of financial risk were used, namely credit risk, liquidity risk, operational risk, and capital risk. Regression analysis indicate there exists a significant negative relationship between the Gulf Cooperation Council Islamic banks' performance, capital risk and operational risk. The results also confirm a significant negative relationship between Gulf Cooperation Council Islamic banks' performance. Furthermore, the results indicate that the most important type of risk is capital risk, followed by operational risk.

JEL: G20, G21

KEYWORDS: Bank Performance, Financial Risk, Gulf Cooperation Council Islamic Banks

INTRODUCTION

Islamic banking has experienced significant growth worldwide during the last three decades. Currently, there are large number of Islamic banks and Islamic units spreading throughout the world. The UK, for example, represents one of the leading centers for Islamic banking globally. International giant banks such as HSBC (HSBC Amanah), Citi Bank (Citi Islamic), and Standard Chartered have established Islamic units. The Gulf Cooperation Council Countries (GCC) operate under dual banking systems: conventional and Islamic banking. 1n 2012, total assets of Islamic banking in the GCC region were 34% of assets of Islamic banks worldwide. Iran's Islamic banking assets were 42.7% of total global Islamic banking assets and Malaysia contributed 10.0% (Zawya, 2013) of the total. Islamic banking is gaining popularity after the 2007–2008 financial crisis. The Islamic banking industry has witnessed a radical change in terms of the number of banks, branches, Islamic units, and Islamic financial instruments. For some experts the main reason for the financial crisis was the collapse in credit supply (e.g., the collapse of Lehman Brothers in September 2008). In the search for solutions there has been interest in the model of interest-free banking used in the Islamic banking industry since Islamic law prohibits usury, the collection and payment of interest.

The objective of this study is to examine the relationship between financial risk and performance of the GCC Islamic banks and the relative importance of the most common types of risk. As mentioned above, Islamic Banking has been growing significantly worldwide during the last three decades and has grown faster in the GCC region. In the UAE, for example, where the first Islamic bank (Dubai Islamic Bank) was

established in 1975, there are now 8 Islamic banks with 283 branches constituting 34% of the total number of bank branches in the country. Total assets of Islamic banks have increased from AED 182.6 billion (about US$ 49.6 billion) in 2008 to AED 358 billion (about US$ 97.3 billion) in 2012. The proportion of UAE Islamic banks' assets increased from 14.9% of UAE banking sector's total assets in 2008 to 20% in 2012.

The structure of the paper is as follows. In the first section, a literature review of the most recent studies is provided. The second section deals with the research questions and hypotheses, followed by an exposition of research methods and data collection. The fourth section is devoted to discussion of the empirical findings. In the final section a brief summary of the paper and conclusions concerning the main results are provided.

LITERATURE REVIEW

In this section we summarize the main findings of selected empirical studies. To the best of the author's knowledge, there is no empirical study that deals with the same issues addressed in this paper, for GCC Islamic banks, or those in other areas of the world. The majority of available empirical studies deal with only one type of risk such as liquidity risk, operational risk, credit risk or capital risk. The following is an attempt to highlight the results of some empirical studies that address one or more of the four types of risk considered in the current study.

Wasiuzzaman and Gunasegavan (2013) compared Islamic and conventional banks in Malaysia. They found that liquidity and operational risks were highly significant in affecting profitability (performance). Febianto (2012) attempted to answer the question: Why are Islamic banks reluctant to indulge in mudharabah and musharakah financing? The main conclusion of the study was that risk management can give Islamic banks guidance on how to manage risk attributed to profit and loss sharing arrangements through Mudharabah and Musharakah contracts. Febianto indicated that this guidance can motivate Islamic banks to be more participative in profit and loss sharing arrangements, especially on their asset side.

Hidayat et al. (2012) investigated the level liquidity risk management effectiveness of Islamic banks in Bahrain. They used a questionnaire which was distributed to a sample of 50 depositors who have active relationships with Islamic banks. The 50 employees chosen were managers and supervisors of Islamic Banks in the Kingdom of Bahrain. The main finding was the positive perception on the status of equity-based financing, which was believed to be an effective part of liquidity risk management. The findings also indicated no significant difference in perception between the employees and depositors on the level of liquidity risk management effectiveness in terms of deposit portfolio and equity financing.

Liquidity risk and banking system performance in Pakistan were examined by Zulfiqar and Anees (2012). The sample period covered was 2004-2009 and includes 22 banks. These banks constitute the main part of the Pakistani banking system. They found that liquidity risk significantly affects bank profitability. Kumaran (2012) examined risk management and mitigation techniques in Islamic finance. The study reveals that Islamic financial institutions face the same risks as conventional banks, namely credit risk, market risk, liquidity risk, and operational risk. However, due to Sharī'ah compliance the nature of these risks change. Abu Hussein and Al-Ajmi (2012) examined risk management practices of conventional and Islamic banks in Bahrain. They found that credit, liquidity and operational risk are the most important risks facing both conventional and Islamic banks. They also found that the levels of risks faced by Islamic banks are significantly higher than those faced by conventional banks.

Ramadan (2011) studies bank-specific determinants of Islamic banks' profitability. He finds that well capitalized banks, efficient management, and higher credit risk lead to higher return on assets, which is used to measure bank performance. He also finds that credit risk positively and significantly affects the profit margin of Jordanian Islamic banks, which is another measure of bank performance. Bank

performance of Islamic banks compared with conventional banks in Indonesia has been investigated by Ika et.al (2011). They conclude that Islamic banks are generally more liquid compared to conventional banks. Boumediene (2011) attempted to answer the question: Is credit risk really higher in Islamic banks? The results indicate that credit risk is indeed higher in Islamic banks compared to conventional banks. Abdullah et al. (2011) investigated operational risk in Islamic banks. The main finding of their study is that risk measurement and risk management practices still need specific adaptations to Islamic banks' operational characteristics. They also highlighted the unique characteristics of Islamic banks and raised serious concerns regarding the applicability of the Basel II methodology for Islamic banks.

Rahman (2010) investigated determinants of the three-factor capital asset pricing model (CAPM) for Malaysian commercial banks. The main findings of this study were: different types of risk exposure have different determinants; the market risk exposure for Islamic banks is lower than for conventional banks; the merger program is fruitful because it reduces interest rate risk exposure, total risk exposure, and unsystematic risk exposure; and the banks under study had higher total and unsystematic risk exposures during the 1997 Asian financial crisis.

Sensarma and Jayadev (2009) examined the relationship between returns on bank stocks and risk management. They found that bank risk management capabilities have been improving over time and returns on bank stocks appear to be sensitive to the risk management capability of banks. Rahman (2010) examined the financing structure and insolvency risk exposure of Islamic banks. The main findings of this study were that an increase in real estate financing decreases insolvency risk, but increasing concentration of financing structure increases insolvency risk. Furthermore, increasing stability of the financing structure reduces risk in the short term. The study recommended the regulatory bodies, policymakers, and market players in the Islamic banking industry should take appropriate action to manage the insolvency risk of Islamic banks.

Volatility of the returns and expected losses of Islamic bank financing were investigated by Ismal (2010). He calculates value at risk (VaR) on the volatility of returns and expected losses of bank financing. He finds that risk of investment and expected losses are well managed. This conclusion was mainly based on the assumption that equity and debt-based financing produce sustainable returns of bank financing. Marcellina (2011) examined credit scoring and risk assessment, and was able to confirm that financial ratios are good predictor variables of bank performance and can be used for classifying and evaluating the bank's customers. Consequently, the bank can reduce its non-performing loans and its credit risk exposure.

Siddiqui (2008) investigated financial contracts, risk, and performance of Islamic banking. The results indicate a good performance of the Islamic banks covered, measured by returns on assets and equity. These banks also demonstrated better risk management and maintained adequate liquidity. Saiful and Mohammad (2008) examined the relationship between risk and return for Islamic bank investment deposits and shareholders' fund. The findings indicate that deposit yields in conventional banks were lower than return on equity (ROE), as a result of the contractual differences between fixed deposit and bank capital. The findings also indicate that Islamic bank deposit yields and ROEs do not reflect their risk-taking properties.

Turk and Sarieddine (2007) highlighted some challenges facing Islamic banks in implementing capital adequacy guidelines. For instance, more complications arise when attempting to measure Sharī'ah compliance risk. Islamic banks are exposed to a significant liquidity risk, as currently Islamic banks tend to rely on short-term Murabahat which is not sufficient for liquidity purposes. Therefore, more work is needed to better account for liquidity risk exposure and risk-weighted assets that do not include assets funded by profit-sharing investment accounts.

How et al. (2005) investigated whether Islamic financing can explain three important bank risks in a country (i.e., Malaysia) with a dual banking system. The three risks examined were credit risk, interest-rate risk,

and liquidity risk. They found that commercial banks with Islamic financing have significantly lower credit and liquidity risks but significantly higher interest-rate risk than banks without Islamic financing. Bank performance and risk has been investigated by John and Courington (1993). They examined the causes of variation in loan performance among banks located in the energy-producing states of Louisiana, Oklahoma, and Texas. The results indicate that a substantial portion of the variation in troubled assets can be attributed to differences in local economic conditions as well as to unusually poor performance of particular industries like energy and agriculture. The results also indicate that excessive risk-taking played a critical role in the loan problems experienced by many of the region's banks and was a contributing force to the diversity in loan performance throughout the region.

METHODOLOGY

The purpose of this study is to examine the relationship between risk and performance of GCC Islamic banks and the relative importance of common types of risk. Thus, we formulate the following hypotheses:

H1: There is a significant relationship between credit risk, liquidity risk, capital risk and operational risk and Islamic banks' performance in the GCC.

H2: There are significant differences in the impact of each type of financial risk on Islamic banks performance in the GCC.

The logic behind the first hypothesis may indicate a positive relationship based on the most common relationship between risk and return. The higher the risk, the more the profit (i.e., improvement in performance) and vice versa. However, an opposite relationship may be also logical if, for example, liquidity risk is increased, which means cash is not sufficient or not available for borrowers and depositors. This lack of liquidity may negatively affect revenues or profit (i.e., performance is getting worse). The purpose of the second hypothesis is to test the relative importance of each type of risk and how it explains the behavior of Islamic bank performance.

Data and the Empirical Model

Based on the previous empirical studies mentioned above, the model used in this study includes some variables previously used in the literature. For example, in evaluating overall bank performance, two ratios are commonly used: return on equity (ROE) and return on assets (ROA). In this study, ROE and ROA are used alternately with four independent variables. The following are brief justifications for use of the independent variables selected here.

The first independent variable is credit risk (CRK) measured by total loans/total assets or loan losses/total loans. It is well established in the literature that there exists a positive relationship between credit risk and profit (see for example Alam et al., 2012). Alam et al. highlighted that banks which have higher loans to total asset ratios tend to take on lower risk. However, if for some reason banks face default or collection problems, the positive relationship between credit risk and profit might not exist.

The second independent variable is liquidity risk measured by total loans/total deposits. The UAE Central Bank determines this ratio as 1:1. Based on this, UAE commercial banks are not allowed to provide loans exceeding their total deposits. However, 1:1 means the maximum limit, as it is risky for banks to use all deposits for lending purposes because they need cash to meet their short-term commitments. The third independent variable is capital risk (CAPR) measured by equity capital/total assets. As capital is used as a cushion, the higher this ratio, the better (less risk) and vice versa.

The fourth variable is operational risk (OPR) measured by the proxy measure cost/income. Ross (2012) explains operating risk as follows: "uncertainty regarding a financial firm's earnings due to failures in computer systems, errors, misconduct by employees, floods, lightning strikes and similar events or risk of loss due to unexpected operating expenses." On the other hand, the European Commission, in line with the Basel II regulations, defines operational risk as "the risk of a change in value caused by the fact that actual losses, incurred for inadequate or failed internal processes, people and systems, or from external events (including legal risk), differ from the expected losses." (see http://en.wikipedia.org/wiki/Operational_risk)

However, as it is difficult to access one measure reflecting the causes of operating risk mentioned in the Ross explanation, cost/income is used as a proxy measure of operating risk. It is worth mentioning here that there are other types of risk, but because of unavailability of data, the study used the above mentioned four types of risk. Examples of other types of risks that were not included are: market risk, interest rate risk, legal risk, foreign exchange risk, and off-balance sheet risk.

The data used in the study covered 11 Islamic banks out of the 47 in the GCC region for the period of 2000 to 2012. Data used in this study were obtained from the Bankscope database. The 11 banks consist of three banks from Kuwait, two banks from Bahrain, three banks from the United Arab Emirates, two banks from Qatar, and one bank from Saudi Arabia. The selection of banks was entirely based on the availability of data.

The name of the banks included were: Kuwait Finance House, Dubai Islamic Bank plc, Abu Dhabi Islamic Bank, Qatar Islamic Bank SAQ, Islamic Development Bank, Qatar International Islamic Bank, Sharjah Islamic Bank, Kuwait International Bank, Arcapita Bank B.S.C., Gulf Finance House BSC, and International Investor Company. The regression model used in this study is as follows:

$$PERF = f(CRK, LIQR, CAPR, OPR) \quad (1)$$

Where:

PERF represents two alternative performance measures for the GCC commercial banks. These two measures are ROA and ROE;
CRK is credit risk = Total loans/total assets or loan losses/total loans;
LIQR is liquidity risk = Total loans/total deposits;
CAPR is capital risk = Equity capital/total assets;
OPR is operational risk = Cost/income.

Three control variables were used. The first control variable is inflation as there exists an inverse relationship between inflation and performance. This relationship has been reported by N'cho-Oguie et al (2011). The second control variable is Bank size, measured by the natural logarithm of bank total assets. In this regard Shrieves and Dahl (1992) and Hussain and Hassan(2004) indicated that size may have an impact on bank f portfolio risk. The third control variable is the GDP growth rate, which is used as proxy for macroeconomic shocks (see Micco and Panizza, 2004 and Yanez, 2007). The macroeconomic shocks affect risk and return (performance). Table 1 indicates summary statistics of data used in the two models. The data represents the variables averages of banks included in the sample.

Table 1: Summary Statistics

Year	ROE	Credit Risk	Liquidity Risk	Capital Risk	Operational Risk	GDP Growth Rate
2000	16.65	56.0500	72.790	12.7000	8.5200	5.98
2001	15.99	68.9400	78.680	10.2200	7.5400	3.01
2003	11.65	57.1300	72.180	16.1900	8.4400	3.42
2004	22.39	61.2600	76.140	16.5300	8.2700	7.57
2005	2.31	58.2900	60.000	83.5800	11.4900	9.70
2006	25.07	59.8700	82.100	13.9200	6.5100	7.02
2007	8.85	72.3800	107.940	26.4600	12.0200	8.58
2008	8.54	63.8100	83.420	16.8900	9.4100	7.79
2009	14.18	22.0400	30.080	40.2600	12.2600	8.92
2010	1.99	34.6700	133.050	47.8600	22.4600	1.88
2011	6.00-	10.3100	45.530	80.2700	32.9600	4.98
2012	7.02	47.2300	52.850	24.2500	48.1100	5.68

This table shows summary statistics of the data.

RESULTS

Using more than one variable to examine the contribution of independent variables to the regression model may result in a multicollinearity problem among these variables. Before examining the contribution of independent variables to the regression model we examine the potential for multicollinearity. A multicollinearity test was carried out to assess the degree of correlation among variables. Table 2 provides the correlations among the independent variables. The "rule of thumb" test proposed by Anderson et al. (1990) suggests that any correlation coefficient exceeding (0.7) indicates a potential problem. The results in Table 2 suggest that correlations among variables are not statistically high enough to suggest the existence of a serious multicollinearity problem.

Table 2: The Correlation Coefficients between Independent Variables

	CRK	LIQR	CAPR	OPR
CRK	1.000			
LIQR	0.427	1.000		
CAPR	-0.605*	-0.228	1.000	
OPR	-0.526	-0.224	0 .352	1.000

*This table shows correlation among independent variables. * indicates correlation is significant at the 0.05 level (2-tailed). ** indicates correlation is significant at the 0.01 level.*

For bank performance, two measures were alternately used, ROA and ROE. In the first model, ROE was used as a measure of bank performance as better results were obtained. Table 3 shows the results of the first regression model. The table shows that the adjusted R square is 64.4%. This indicates that the four independent variables explain 64.4% of the variance of GCC Islamic bank performance. The estimated coefficients of the four independent variables were, as expected, negative and statistically significant at the 5 percent level in the case of capital risk and at 10 percent in the case of operational risk. These findings are consistent with those of Wasiuzzaman and Gunasegavan (2013). However, the estimated coefficient of the other two variables, credit risk (CRK) and liquidity risk (LIQR), were unexpectedly statistically insignificant.

Table 3: Summary of Regression Results – The First Model

	Beta	T
(Constant)		3.156
CRK	-0.009	-0.033
OPR	-0.399	-1.883*
CAPR	-0.703	-3.108**
LIQR	-0.231	-1.162*
R	0.879	
R Square	0.773	
Adjusted R Square	0.644	
Std. Error of the Estimates	5.296	

*This table shows regression estimates of the equation $PERF = f(CRK, LIQR, CAPR, OPR)$. The dependent variable is ROE (net income/equity) The independent variables are credit risk (CRK), operational risk (OPR), capital risk (CAPR), and liquidity risk (LIQR). The table reveals the coefficient values, the t-statistics and the significant level. * indicates statistically significant at the 5 percent level and ** indicates statistically significant at the 10 percent level.*

The second model was examined by considering the ROE as the dependent variable and the four independent variables used in the first model with one additional control variable, real GDP growth rate (GDPG). However, the other two control variables, were also included in the second model, but the results were meaningless. Table 4 reveals the results of the test. The adjusted R square is 59.5%. This indicates that the four independent variables explain 59.5% of the behavior of bank performance of the GCC Islamic banks. The estimated coefficient values are as expected negative except liquidity risk (LIQR). However, the variables are statistically insignificant except capital risk (CAPR) which was statistically significant at the 1 percent level.

Table 4: Summary of Regression Results – The Second Model

	Beta	T
(Constant)		2.384
CRK	-0.113	-0.372
OPR	-0.358	-0.594
CAPR	-0.788	-3.049*
LIQR	-0.128	0.522
GDPG	0.187	0.764
R	0.891	
R Square	0.793	
Adjusted R Square	0.621	
Std. Error of the Estimate	5.463	

*This table shows the regression estimates of the equation: $PERF = f(CRK, LIQR, CAPR, OPR, GDPG)$. Dependent Variable (PERF) is measured by ROE (net income/equity) and the independent variables are credit risk (CRK), operational risk (OPR), capital risk (CAPR), liquidity risk (LIQR) and GDP growth rate (GDPG). The table presents coefficient values, t-statistics and the significant level. * indicates statistically significant at the 5 percent level*

We noted that banks which have a higher loans-to-total-asset ratio tend to take on lower risk (i.e., an inverse relationship between risk and performance). However, the results were not supported by the expected negative relationship between performance and liquidity risk as the value of the coefficient was statistically insignificant in the two models.

The results of the current study are consistent with findings of Zulfiqar and Anees (2012) and Ramadan (2011). However, the results are inconsistent with those of Hayden et al. (2007) who attempted to answer the question: Does diversification lead to increased performance? As diversification affects the level of risk, the more the diversification, the lower the risk and vice versa. They found little evidence of large performance benefits associated with diversification.

The second hypothesis proposes the magnitude of the impact of each type of risk on GCC Islamic bank performance is significantly different. The results confirmed this hypothesis. The results provided in Tables

3 and 4 show the estimated coefficients were statistically significant for two independent variables, with capital risk ranked first followed by, operational risk. This finding is consistent with Rahman (2011).

CONCLUDING COMMENTS

The objective of this study was to examine the relationship between financial risk and performance of GCC Islamic banks and the relative importance of common types of risk. The study covers 11 of the 47 Islamic banks in the GCC region for the period from 2000–2012. For bank performance the two most common measures, ROA and ROE, were alternately used. Four types of risk were examined: credit risk, liquidity risk, operational risk and capital risk. The selection of these banks and types of risk was determined by data availability. By using two regression models, two performance measures were used. The results as expected indicate a significant negative relationship between GCC Islamic bank performance and two types risk, namely capital risk and operational risk. The positive relationship between risk and performance of the GCC Islamic banks was not confirmed. Furthermore, the results indicate that the most important type of risk is capital risk followed by operational risk.

For policy implementation, it is recommended that more attention be paid to capital risk, as this type of risk represents the main determinant of performance, measured by either the equity or assets components. In addition, more attention should be given to operational risk which is mainly related to uncertainty regarding a financial firm's earnings due to failures in computer systems, errors, misconduct by employees, or risk of loss due to unexpected operating expenses. Finally, more attention should also be paid to liquidity risk which represents a determinant of GCC Islamic bank performance.

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