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EXCHANGE RATE EFFECTS ON A SMALL OPEN ECONOMY: EVIDENCE FROM TAIWANESE FIRMS

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ABSTRACT

Previous empirical research discovered only mild, if any, sensitivity of firm value to exchange rate fluctuation. Chen et al. (2004) provided some insights by focusing on a small and open economy and found evidence that New Zealand that exchange rate movement affects firm value. This study reexamines firm value sensitivity to exchange rate fluctuation by focusing on individual firms as well as on three industry Taiwan sectors, high-tech, service, and manufacturing industries. By using the two-factor model with residual regression, we find consistent results that volatility of exchange rates affects the value of Taiwanese firms. The results hold regardless of the exchange rate exposure to US dollar, Japanese Yen, or Euro. In addition, the positive association between exchange rate exposure and firm value is significant and consistent for all firm samples and three industry-specific samples.

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KEYWORDS: Foreign exchange exposure, Residual regression, Exchange rate fluctuations, Firm value

INTRODUCTION

With the ever increasing tempo of economic globalization, more firms face international competition. It is widely accepted that fluctuations in exchange rates affect the competitive position of firms in the global market, and eventually impacts cash flows and firm value. Moreover, even for a domestic firm that is not actively involved in international trade, variations in exchange rates may indirectly affect firm value. This is especially true if the firm relies on imports to obtain raw materials, if its suppliers are subject to exchange rate exposure, or if the firm's competitors are from overseas.

Presumably, only unanticipated fluctuations in exchange rates affect firm value. According to the efficient market hypothesis, the expected change in exchange rate should be priced into a firm's stock price. Therefore the relation between unexpected exchange rate change and share price should be observed. However, previous studies failed to document a strong relationship between unexpected movement of foreign exchange rates and changes in share price. Many studies only discovered weak or insignificant associations between firm value and exchange rate exposure (Jorin, 1990; Gao, 2000; Griffin and Stulz, 2001; Di Iorio and Faff, 2002). Even if a significant sensitivity was found, the mean sensitivity values were far below what is predicted by theories (Booth and Rotenberg, 1990; Choi and Prasad, 1995; Frennberg, 1994).

A few empirical studies found significant exchange rate risk sensitivity under certain conditions. Bartov and Bodnar (1994) found that abnormal returns are related to lagged changes of exchange rates, which supports market inefficiency. The studies by Chow et al. (1997) and Bodnar and Wong (2000), show the association between firm value and exchange rate changes becomes significant when the time horizon is increased. According to Jorion (1990) and Shin and Soenen (1999), the magnitude of exchange rate exposure varies by firm size.

These mixed results are typically attributed to three causes. First, different models were employed and different firm sample sizes were selected. Usually different methodologies lead to divergent results.

Second, the fundamental cash flow models available might not be a good match to reality. Third, firms in an open economy may have a higher exchange rate exposure than those in a closed economy (Chen et al., 2004; Huston and Stevenson, 2010).

Chen et al. (2004) argued that there could be a serious problem in the data selection process of previous studies that focus on large capital markets. For example, in the US stock market, domestic factors dominate the pricing of consumer goods and stocks, dwarfing the international factors, such as exchange rate changes. US multinational firms could be so globally diversified that a causal relationship between firm value and exchange rate exposure is difficult to detect. In addition, Huston and Stevenson (2010) argue that a country-specific governance environment could provide incentives to enhance shareholder value in response to reducing exchange rate exposure.

We take into account both the country's economy openness and country-specific governance, in our attempt to investigate the relation between exchange rate movement and firm value. By employing a sample of firms from Taiwan, a small and open economy, our study investigates whether an association between exchange rate exposure and firm value exists. In comparison to the New Zealand market used in Chen et al. (2004), our sample of Taiwanese firms is unique in several ways: (1) Prior studies employ monthly exchange rate changes to measure exchange rate exposure. Our study uses daily exchange rate changes. The application of a shorter horizon to capture the exchange rate change is intended to provide a more powerful test setting with lower noise. (2) International trade in the Taiwan economy is essential and well known for its electronics manufacturing and high-tech industries.

Listed firms in these industries account for approximately half of Taiwan stock market as measured by market capital. (3) In general, the Taiwan market is bigger in total capital and more open than most other markets in the Asia-Pacific region. However, it has a few restrictions and the market is much smaller than major markets in the US, UK and Japan. The exchange rate in Taiwan has been volatile since the New Taiwan Dollar (NTD) began to float in 1989, when the upper/lower 2.5% limit on daily exchange rate fluctuation and other important financial regulations were lifted. The movement of exchange rates is based on market demand and supply. The Taiwan Central Bank only executes appropriate interventions on seasonal or abnormal floating when needed. (4) Most Taiwanese firms are foreign currency denominated price-takers rather than world market leaders, which implies that Taiwanese firms may be more susceptible to the influence of exchange rate fluctuations than those in large major markets.

Our study investigates firm value sensitivity to exchange rate daily fluctuations by focusing on a sample of individual firms in Taiwan from three different industry sectors, rather than from a market portfolio. We find consistent results that a positive association between foreign exchange rate exposure and firm value exists in Taiwan. Of the three industry sectors, the high-tech industry is most affected by exchange rate fluctuations.The rest of the paper is organized as follows. Section 2 reviews the recent literature. Section 3 describes the data and samples. Section 4 presents the methodology. Finally, empirical results analysis and discussion are presented in Section 5. The last two sections conclude our research and discuss the research limitation.

LITERATURE REVIEW

According to Dumas (1978), and Adler and Dumas (1984), foreign exchange exposure can be measured by the sensitivity of firm value to unexpected foreign exchange rate changes. To find the relationship between exchange rate changes and firm value, most research (such as Jorion, 1990; Bartov and Bodnar, 1994; Choi and Prasad, 1995) used stock price as the proxy of firm value and examine the connection between exchange rate fluctuations and stock prices by regression. However, early studies found limited and mixed evidence of foreign exchange exposure. Jorion (1990) investigated monthly stock returns of 287 U.S. multinationals from 1971 to 1987 and found that only some firms showed statistically significant foreign exchange exposure. In the following study, Jorion (1991) examined the exposure at the industrial level and discovered that most industry groups of stocks traded on the New York Stock Exchange do not exhibit significant exposure. Choi and Prasad (1995) also studied 409 US multinationals from 1978 and 1989 at industrial level and documented limited foreign exchange exposure for the industries. They further tested the association between foreign exchange exposure and firm-specific characteristics and found a positive relation between the scope of foreign operation and exposure. Consistent with this finding, Bartov and Bodnar (1994) suggested that in the sample selection process, those firms reporting significant impacts from changes in the value of dollars should be included. Despite this procedure they did not find significant foreign exchange exposure. In their further test, though, a significant association between abnormal stock returns and lagged dollar exchange rate changes was documented. They argue it took time for investors to price exchange exposure into stock prices.

Some other studies also target firms with likelihood of high exchange exposure. Amihud (1994) examined the 32 largest U.S. exporting firms from 1982 to 1988. They failed to document any significant contemporaneous association between stock returns and exchange rate fluctuation. Similarly, Donnelly and Sheehy (1996) selected only the largest UK exporters and found a weak contemporaneous exchange exposure. Apart from the scope of foreign operation, some other factors are associated with exchange exposure. Theoretically, exchange exposure can be due to short-term transaction exposure or long-term economic exposure. Usually, transaction exposure can be reduced to minimum by hedging while economic exposure is harder to hedge. Compared to small firms, large firms have more resources and are more capable of hedging economic exposure efficiently.

Chow et al. (1997) found that the exposure magnitude is inversely related to firm size. Also, by employing a longer horizon for stock returns and exchange rates, they found the significance of exchange exposure increases as the length of the horizon increases. The type of economy has been argued to play a role in the magnitude of foreign exchange exposure. Studies, such as Nydahl (1999), De Jong et al. (2002), Kiymaz (2003) and Moran (2005) examine several countries and document a larger number of significantly exposed firms with higher exposure coefficients than in prior studies. Chen et al. (2004) suggested that large markets like the U.S. should be less vulnerable to the exchange rate fluctuations, with domestic factors playing a predominant role in the determination of both share and consumer goods prices. Therefore, it could be difficult to identify significant exchange exposure for a large economy like the U.S. Instead, they selected 164 listed firms on the New Zealand Stock Exchange for the period from January 1994 to December 1999, and found significant relation between the change of exchange rate and firm value.

Most prior studies of exchange rate exposure are focused on the correlation of monthly stock returns and change in exchange rates. However, as Frennberg (1994) pointed out, the market is believed to react very fast to new information and the focus should be on immediate reactions to exchange rate changes. Compared to the monthly settings employed by prior research to test exchange exposure, we believe that a shorter horizon of one day to capture the variation of exchange rate provides an opportunity for a more powerful test with less noise.

METHODOLOGY

Data and Sample

Considering the characteristics of the Taiwan's economy, we selected our firm samples from two industry sectors: manufacturing and service. The manufacturing industry is the driver of Taiwan's economy, valued at $259 billion in 1999 and accounted for 98.6% of its exports in 2000. Electronics, textiles and plastics are three major exporting industries in Taiwan. Statistics shows that the electronics industry accounts for the largest share of Taiwan's manufacturing industry, totaling $101 billion in 1999. Due to the rapid growth in this subsector, Taiwan has become one of the largest producers of information technology hardware, and ranks first in the production of laptop computers, monitors and electronics related peripherals. Given the distinctions between electronics and traditional manufacturing industries and the predominant position of the electronics industry in Taiwan's economy, we set this industry aside from the other two major manufacturing industries and categorize it as the high-tech industry. In doing so, we expect to find evidence that the electronics industry may have different sensitivity to exchange rate exposure than traditional manufacturing industries. Firms newly listed or delisted during the time frame of our study between January 1, 2001 and December 31, 2002 are eliminated. From a total of 305 listed Electronics firms and 73 listed Textile or Plastics firms, the top ranked 60 firms for the former and top ranked 40 firms for the latter are selected. Excluding firms with missing data, our final sample consists of 59 companies from the high-tech industry, and 37 firms from the traditional manufacturing industries.

In addition to the test for the manufacturing sector, this study also examines whether the relationship between exchange rate movements and change in firm value still holds in the service industry. Following Fama and French (1992), we did not include financial service firms due to their distinct financial structures and unique characteristics. As an important element in Taiwan's future economic development, the tourism industry has called for investment totaling over NT$75 billion. Therefore, 14 top ranked firms from the tourism and department store industry are selected for inclusion in the service industry sample.

Eaily share returns adjusted for dividends for the selected firms listed on the Taiwan Stock Exchange (TSE) and the composite daily index for Taiwan stock market are obtained from the Global Data Stream database for the period between January 1, 2001 and December 31, 2002. The market index is used to capture the influence of general economic conditions on firm value. In addition, daily exchange rates for the New Taiwan Dollar (NTD) relative to three major world three currencies including US Dollar, Euro and Japanese Yen, are also obtained from the Global Data Stream database for the same period. After excluding the non-trading days, there are 508 days in the two-year window.

Based on the market efficiency hypothesis, the effect of unexpected changes in exchange rates should be impounded into a firm's stock price and reflect on its firm value quickly. Adler and Dumas (1984) suggest that a firm's foreign exchange exposure can be measured by the sensitivity of equity returns to exchange rate changes. In other words, a firm is subject to exchange rate exposure if exchange rate fluctuations affect future cash flows and eventually firm value. Thus, we adopt the regression models for exchange rate exposure from the prior studies (Adler and Dumas, 1984; Jorion, 1991; Choi and Prasas, 1995; Dumas and Sonik, 1995; De Santis and Gerard, 1998; Chen at el. 2004). Equation (1) is the direct model adopted from Adler and Dumas (1984), which aims to capture the average sensitivity of firm value to unexpected exchange rate fluctuations.

$$R_{it} = \alpha_{i0} + \alpha_{i1}R_{xt} + e_{it} \qquad (1)$$

where R_i is the stock return of the ith firm and R_x is the change of exchange rate. The exchange rate is expressed in the units of foreign currency per unit of domestic currency of Taiwan (NTD).

Application of the model includes the assumption that the firm valuation effect is influenced by macroeconomic conditions. Foreign exchange rates are one of these conditions. Chen et al. (2004) suggests that this model is an appropriate foreign exchange exposure measure if other general macroeconomic factors are not only (or nearly) normally distributed, but also stable over the testing period. Our study employed a short time frame of two years from 2001 to 2002, which is characterized by relatively stable economic developments in Taiwan. Model (1) is included in our study to test the sensitivity of stock returns to exchange rate movement.

To improve the specification of the model, Jorion (1991) proposed an enhanced two-factor model which is widely used in prior studies (Choi and Prasas, 1995; Moran, 2005; Huston and O'Driscoll, 2010). Augmented by the market return factor, Equation (2) controls for macroeconomic market conditions. Accordingly, the coefficient on the term of exchange rate change is to measure the risk sensitivity of idiosyncratic exchange rate exposure for individual firm.

$$R_{it} = \beta_0 + \beta_1 R_{xt} + \beta_2 R_{mt} + e_{it} \tag{2}$$

where R_m is market index return, R_i is the stock return of firm i, and R_x is the change of exchange rate expressed in the unit number of foreign currency per unit of domestic currency.

Chen et al. (2004) analyze the two-factor model and partition the effect of exchange rate movement on firm value into two components: (1) the indirect effect of exchange rate movement on firm value through its effect on the market index, and (2) the direct effect on stock returns of individual firms. The paper concludes that both effects will be reflected in the firm's market value (Chen et al., 2004). Thus model (2) attempts to capture the sensitivity of stock returns to both components of exchange rate movement and market return. However, the two-factor model doesn't take into consideration that the market index itself has substantial foreign exchange exposure. Therefore, Chen et al. (2004) derive a model by residual regression as specified in model (3), with aims to capture the total effect (including general and unique effect) of exchange rate fluctuation on the market movement.

Based on the derivation from Chen et al. (2004), the first regression is a residual model of market index.

$$R_{mt} = \delta_{m0} + \delta_{m1} R_{xt} + e_{mt} \tag{3}$$

where R_m is market index return, and R_x is the change of exchange rate expressed in the unit number of foreign currency per unit of domestic currency.

Then the residual term replaces the market index return in Jorion's (1991) model as presented in (2). Therefore, stock return is regressed on the residual of market index return and exchange rate exposure.

$$R_{it} = \gamma_{i0} + \gamma_{i1} R_{xt} + \gamma_{i2} e_{mt} + e_{it} \tag{4}$$

The component e_{mt} is the residual item of the residual model and is orthogonal to the market index return. The exchange rate coefficient γ_{i1} in Equation (4) is a measure of total effect, which is the net of general influence and unique influence (Chen et al., 2004).

We use all three models to measure the effect of foreign exchange exposure on firm value and test whether Taiwanese firms experienced a significant foreign exchange rate exposure to most common foreign currencies, including US dollars, Japanese Yen, and Euro, during 2001-2002. It is likely that the value of exchange rate coefficient will vary across different models.

ANALYSIS

Statistical Analysis for the Full Sample

Table 1 presents descriptive statistics for the full sample of 110 firms. It reports the distribution of firm size measured by total assets, total sales, and market value, respectively, as well as some firm-specific ratios. The mean (median) total assets is over 43 (15) million NTDs. The mean (median) total sales is over 27 (11) million NTDs. The mean (median) market value is over 45 (13) million NTDs. The mean (median) sales growth rate is around 25% (16%), the mean (median) gross profit ratio is 18% (16%), and the debt ratio is 0.40 on average. The mean (median) market to book ratio is 1.885 (1.558).

Table 1: Descriptive Statistics for All Firms

	Mean	Median	Max	Min	25%	75%	Std Dev
Total Assets (in thousand NTD)	43,685	15,985	465,863	740.44	7,429.4	49,694	71,813
Total Sales (in thousand NTD)	27,068	11,235	245,009	593.49	5,362.3	30,257	39,535
Market Value (in thousand NTD)	45,167	13,530	793,335	2,065.0	6,033.8	39,632	99,264
L-T Investment (in thousand NTD)	11,116	3,730.8	114,848	7.0680	1,253.6	12,145	17,838
Sale growth rate	0.2548	0.1577	3.234	-0.2798	0.0139	0.3638	0.4336
Gross profit ratio	0.1833	0.1565	0.7752	-0.3116	0.1052	0.2166	0.1367
Debt ratio	0.4027	0.4082	0.7422	0.0922	0.3090	0.4852	0.1424
Market to Book ratio	1.885	1.558	7.006	0.4248	1.003	2.160	1.316

This table presents the descriptive statistics for all firms based on the annual data in 2002. Market value is based on the multiplication of number of shares outstanding at the end of 2002 and the stock price on the last trading day of the calendar year of 2002. Debt ratio is based on total liabilities divided by total assets.

Table 2 presents the Pearson correlations between the variables for the 55,769 firm-day sample. As expected, daily stock returns and market returns are highly correlated, with the correlation coefficient of 0.5269. However, the exchange rate variables for US Dollar, Japan Yen, and Euro are not highly correlated with daily stock returns in the sample. Similar results are found in the test for Spearman's correlation. Further, we break down the full sample into industry sectors, and find that the correlation matrix for high-tech, service, or manufacturing industry is qualitatively the same as presented in Table 2.

Table 2: Pearson Correlation between Variables for All Firms

Pearson Correlation	R_{it}	R_{mt}	Euro	Yen	USD
R_{it}	1.000	0.5269 (0.0001)***	0.0978 (0.0001)***	0.0364 (0.0001)***	0.0673 (0.0001)***
R_{mt}		1.000	0.2000 (0.0001)***	0.0792 (0.0001)***	0.1673 (0.0001)***
Euro			1.000	0.2063 (0.0001)***	0.0870 (0.0001)***
Yen				1.000	-0.0246 (0.0001)***
USD					1.000

*This table presents the Pearson correlation between the variables for the full sample of 55,769 firm-days. R_{it} is the daily return adjusted for dividend for individual firms. R_{mt} is the composite daily return for Taiwan stock market. Euro is the change in daily exchange rate of New Taiwan Dollar (NTD) to Euro. Yen is the change in daily exchange rate of New Taiwan Dollar (NTD) to Yen. USD is the change in daily exchange rate of New Taiwan Dollar (NTD) to USD. The figure in parentheses in each cell is the p-value. ***, ** and * indicate significance at the 1, 5 and 10 percent levels respectively.*

Table 3 reports the regression results by model. As our hypothesis predicts, the exchange rate fluctuation should positively associate with firm value change. The test results from Equation (1) and Equation (4) support this conjecture. The coefficient on the exchange rate change for USD is 35.03 at the 1% significance level in Equation (1), and 35.03 at the 1% significance level in Equation (4). The results for Yen and Euro are similar. Equation (2) doesn't show the same results. The Euro is negatively associated with stock returns, with a coefficient of -1.290 (p-value, 0.0359). Such results are not as predicted by our hypothesis but are consistent with the results of the sensitivity and robustness tests in Chen at el. (2004), indicating that Equation (2) is the least sensitive model and measures the exchange rate exposure in a different degree.The regression analysis also supports the above argument for Table 3. The R^2 and adj-R^2

levels in Equation (2) and Equation (4) have exactly the same values across all findings in Table 3. It is noteworthy that the coefficient of market returns to USD in Equation (2) is 0.9926, identical with the residual variable coefficient in Equation (4). Both are significant at 1% level. The results suggest that Equation (4) has eliminated the mingling effect of exchange rate exposure to daily returns through market return while capturing the same effect from Equation (2).

The signs of exchange rate exposure from Equation (2), however, are very different from Equation (4). For example, the only significant exchange rate exposure in Equation (2) is from the Euro, which is -1.290 (significant at 0.0359 level). In Equation (4), however, the exposures to all major currencies are positive and significant. The only difference between these two models is the use of different market return measures. Equation (2) uses the market return itself while Equation (4) uses the residual market return, which is orthogonal to exchange rate fluctuations. The coefficients on the residual market return in Equation (4) are very similar to the respective coefficients on market return in Equation (2). Therefore, by using the residual market model, Equation (4) is able to capture most of the influence of market return on individual stock returns. After comparing the results from Equation (2) with that from Equation (4), we argue that the orthogonalization in Equation (4) does not distort our results, but only improves its precision. In summary, we find that the relation between exchange rate change and firm value is significantly positive for the Taiwanese firms in the sample.

Table 3: Regression Results for the Firms in the Full Sample on USD/Yen/Euro Exposure

All firms N=55,769		Model 1: Equation (1)		
Variable	**Prediction**	**Coefficient (p-value)**	**Coefficient (p-value)**	**Coefficient (p-value)**
Intercept	?	0.0010*** (0.0001)	0.0009*** (0.0001)	0.0010*** (0.0001)
US	+	35.031*** (0.0001)		
Yen	+		0.0594*** (0.0001)	
Euro	+			16.379*** (0.0001)
R^2		0.0076	0.0013	0.0096
Adj-R^2		0.0076	0.0013	0.0095
		Model 2: Equation (2)		
Intercept	**?**	**0.0008*** (0.0001)**	**0.0008*** (0.0001)**	**0.0008*** (0.0001)**
US	+	-0.3459 (0.8133)		
Yen	+		-0.0088 (0.1361)	
Euro	+			-1.290** (0.0359)
R_{mt}	+	0.9926*** (0.0001)	0.9931*** (0.0001)	0.9952*** (0.0001)
R^2		0.2776	0.2777	0.2777
Adj-R^2		0.2778	0.2776	0.2777
		Model 3: Equation (4)		
Intercept	**?**	**0.0010*** (0.0001)**	**0.0009*** (0.0001)**	**0.0010*** (0.0001)**
US	+	35.031*** (0.0001)		
Yen	+		0.0594*** (0.0001)	
Euro	+			16.379*** (0.0001)
Residual	+	0.9926*** (0.0001)	0.9930*** (0.0001)	0.9952*** (0.0001)
R^2		0.2776	0.2777	0.2777
Adj-R^2		0.2778	0.2776	0.2777

*This table reports the regression results for the full sample by models. Model 1 is $R_{it} = \alpha_{i0} + \alpha_{i1}R_{xt} + e_{it}$, Model 2 is $R_{it} = \beta_0 + \beta_1 R_{xt} + \beta_2 R_{mt} + e_{it}$, while Model 3 is $R_{it} = \gamma_{i0} + \gamma_{i1}R_{xt} + \gamma_{i2}e_{mt} + e_{it}$, where $R_{mt} = \delta_{m0} + \delta_{m1}R_{xt} + e_{mt}$. The dependent variable R_{it} is the daily return adjusted for dividend for individual firms. R_{mt} is the composite daily return for Taiwan stock market. R_{xt} is the change in daily exchange rate of New Taiwan Dollar (NTD) to a foreign currency, which is Euro, Yen or USD The figure in parentheses in each cell is the p-value. ***, ** and * indicate significance at the 1, 5 and 10 percent levels respectively.*

<u>Statistical Analysis for the Industry Specific Samples</u>

We divide the full sample into sub-samples by industry. The Taiwanese economy is famous for its high-tech electronics industry, which plays a major role in economic growth and accounts for most imports/exports of Taiwan. The service industry in Taiwan is highly subject to exchange rate exposure due to the significant portion of revenues from overseas customers.

There are three industry samples in the following investigation. The high-tech industry sample includes 60 firms selected with 29,912 firm-days, from the areas of semiconductor and related, PC and its peripheral and other electronics. The service industry sample includes 14 firms with 7,097 firm-days, from the areas of department stores, tourism and international trading. The manufacturing industry sample includes 37 firms with 18,758 firm-days from the areas of textiles and plastics manufacturing. We expect a positive correlation between exchange rates and firm value for the two industries of high-tech and service. The manufacturing industry in Taiwan tends to be domestic-oriented. Although we still expect a positive correlation between exchange rate movement and firm value for this industry, the magnitude of the correlation might not be as strong as for the prior two industries.

Tables 4, 5 and 6 present the regression results by industry: high-tech industry, service industry and manufacturing industry. The results for high-tech industry in Table 4 are consistent with those for the full sample. Similarly, Equation (1) and Equation (4) both provide better results in term of precision than in Equation (2). The results from Equation (2), however, show significant negative association between exchange rate exposure across all major foreign currencies. For example, the coefficient is -8.900 (p-value 0.0001) in USD in Equation (2) and -3.720 (p-value 0.0001) in Euro in Equation (2).

The high-tech industry in Taiwan is highly export-oriented and dispersed in worldwide markets. Therefore, a significant and positive relationship with exchange rate exposure is expected. However, we fail to find a positive exchange rate exposure from Yen in Equation (2). Further study of this issue is warranted. An interesting observation is that Equation (2) and Equation (4) in Table 4 share the same values of R^2 and adj-R^2, which supports that Equation (4) has the same explanatory power as Equation (2). In addition, the adj-R^2 in Table 4 in Equation (4) is always higher than that in Equation (1), suggesting that Equation (4) has a higher explanatory power than Equation (1) by adding the term of residual market return. Further, we find that high-tech industry exhibits the highest adj-R^2 among all industries. This is consistent with the fact that the high-tech firms in Taiwan have partnered with big names in the US, Japan and Europe, and “made in Taiwan” PCs, electronics components, and semiconductors are widespread across the world.

Table 5 presents the results on exchange rate exposure for firms in the service industry. We detect a positive relation between exchange rate fluctuation and firm value in this industry, in both Equation (1) and Equation (4), across all major foreign currencies.

For the manufacturing industry sector, we find consistent results in Table 6 that a positive association exists between exchange rate movement and firm value for this industry, for USD and Euro. Our results from Table 6, however, do not support that exchange rate changes of Japanese Yen directly affect firm value (p-value 0.3296) in Equation (2). The possible explanation is that the Japanese government has been restricting imports of textile products and thus the textile industry in Taiwan is not subject to a high exchange rate exposure to Japanese Yen.

Table 4: Regression Results for the Firms in Hi-tech Industry on USD/Yen/Euro Exposure

Hi-tech firms N=29,912			**Model 1: Equation (1)**	
Variable	**prediction**	**Coefficient (p-value)**	**Coefficient (p-value)**	**Coefficient (p-value)**
Intercept	?	0.0008*** (0.0001)	0.0007*** (0.0001)	0.0009*** (0.0001)
US	+	36.752*** (0.0001)		
Yen	+		0.0800*** (0.0001)	
Euro	+			19.045*** (0.0001)
R^2		0.0078	0.0022	0.0012
Adj-R^2		0.0078	0.0022	0.0012
			Model 2: Equation (2)	
Intercept	**?**	**0.0006*** (0.0001)**	**0.0006*** (0.0001)**	**0.0006*** (0.0002)**
US	+	-8.900*** (0.0001)		
Yen	+		-0.0075 (0.3123)	
Euro	+			-3.720*** (0.0001)
R_{mt}	+	1.281*** (0.0001)	1.275*** (0.0001)	1.282*** (0.0001)
R^2		0.4263	0.4259	0.4263
Adj-R^2		0.4263	0.4259	0.4263
			Model 3: Equation (4)	
Intercept	**?**	**0.0008*** (0.0001)**	**0.0007*** (0.0001)**	**0.0009*** (0.0001)**
US	+	36.752*** (0.0001)		
Yen	+		0.0800*** (0.0001)	
Euro	+			19.046*** (0.0001)
Residual	+	1.281*** (0.0001)	1.275*** (0.0001)	1.282*** (0.0001)
R^2		0.4263	0.4259	0.4263
Adj-R^2		0.4263	0.4259	0.4263

*This table reports the regression results by models for the sample of firms in the high-tech industry. Model 1 is $R_{it} = \alpha_{i0} + \alpha_{i1}R_{xt} + e_{it}$, Model 2 is $R_{it} = \beta_0 + \beta_1 R_{xt} + \beta_2 R_{mt} + e_{it}$, while Model 3 is $R_{it} = \gamma_{i0} + \gamma_{i1}R_{xt} + \gamma_{i2}e_{mt} + e_{it}$, where $R_{mt} = \delta_{m0} + \delta_{m1}R_{xt} + e_{mt}$. The dependent variable R_{it} is the daily return adjusted for dividend for individual firms. R_{mt} is the composite daily return for Taiwan stock market. R_{xt} is the change in daily exchange rate of New Taiwan Dollar (NTD) to a foreign currency, which is Euro, Yen or USD. The figure in parentheses in each cell is the p-value. ***, ** and * indicate significance at the 1, 5 and 10 percent levels respectively.*

Table 5: Regression Results for the Firms in Service Industry on USD/Yen/Euro Exposure

Service industry N=7,097			**Model 1: Equation (1)**	
Variable	**prediction**	**Coefficient (p-value)**	**Coefficient (p-value)**	**Coefficient (p-value)**
Intercept	?	0.0004 (0.2431)	0.0003 (0.3464)	0.0004 (0.2268)
US	+	24.183*** (0.0001)		
Yen	+		0.0482*** (0.0001)	
Euro	+			9.351*** (0.0001)
R^2		0.0059	0.0014	0.0051
Adj-R^2		0.0058	0.0013	0.0049
			Model 2: Equation (2)	
Intercept	?	0.0003 (0.3181)	0.0003 (0.3506)	0.0003 (0.3420)
US	+	6.575* (0.0655)		
Yen	+		0.0140 (0.3296)	
Euro	+			0.5080 (0.7349)
R_{mt}	+	0.4940*** (0.0001)	0.4979*** (0.0001)	0.4981*** (0.0001)
R^2		0.1143	0.114	0.1139
Adj-R^2		0.1141	0.1138	0.1137
			Model 3: Equation (4)	
Intercept	?	0.0004 (0.2162)	0.0004 (0.2162)	0.0004 (0.2003)
US	+	24.183*** (0.0001)		
Yen	+		0.0481*** (0.0008)	
Euro	+			9.351*** (0.0001)
Residual	+	0.4940*** (0.0001)	0.4940*** (0.0001)	0.4981*** (0.0001)
R^2		0.1143	0.1143	0.1139
Adj-R^2		0.1141	0.1141	0.1137

*This table reports the regression results by models for the sample of firms in the service industry. Model 1 is $R_{it} = \alpha_{i0} + \alpha_{i1}R_{xt} + e_{it}$, Model 2 is $R_{it} = \beta_0 + \beta_1 R_{xt} + \beta_2 R_{mt} + e_{it}$, while Model 3 is $R_{it} = \gamma_{i0} + \gamma_{i1}R_{xt} + \gamma_{i2}e_{mt} + e_{it}$, where $R_{mt} = \delta_{m0} + \delta_{m1}R_{xt} + e_{mt}$. The dependent variable R_{it} is the daily return adjusted for dividend for individual firms. R_{mt} is the composite daily return for Taiwan stock market. R_{xt} is the change in daily exchange rate of New Taiwan Dollar (NTD) to a foreign currency, which is Euro, Yen or USD. The figure in parentheses in each cell is the p-value. ***, ** and * indicate significance at the 1, 5 and 10 percent levels respectively.*

Table 6: Regression Results for the Firms in Manufacturing Industry on USD/Yen/Euro exposure

Manufacturing Industry N=18,758		**Model 1: Equation (1)**		
Variable	**prediction**	**Coefficient (p-value)**	**Coefficient (p-value)**	**Coefficient (p-value)**
Intercept	?	0.0007*** (0.0001)	0.0007*** (0.0001)	0.0007*** (0.0001)
US	+	13.935*** (0.0001)		
Yen	+		0.0050 (0.5360)	
Euro	+			6.045*** (0.0001)
R^2		0.0026	0	0.0028
Adj-R^2		0.0026	0	0.0028
		Model 2: Equation (2)		
Intercept	?	0.0007*** (0.0001)	0.0007*** (0.0001)	0.0007*** (0.0001)
US	+	1.271 (0.5102)		
Yen	+		-0.0196**	
Euro	+			-0.2878
R_{mt}	+	0.3550*** (0.0001)	0.3578*** (0.0001)	0.3567*** (0.0001)
R^2		0.078	0.0783	0.078
Adj-R^2		0.0779	0.0782	0.0779
		Model 3: Equation (4)		
Intercept	?	0.0007*** (0.0001)	0.0007*** (0.0001)	0.0007*** (0.0001)
US	+	13.925*** (0.0001)		
Yen	+		0.0050 (0.5192)	
Euro	+			6.045*** (0.0001)
Residual	+	0.3550*** (0.0001)	0.3578*** (0.0001)	0.3567*** (0.0001)
R^2		0.078	0.0783	0.078
Adj-R^2		0.0779	0.0782	0.0779

This table reports the regression results by models for the sample of firms in the manufacturing industry. Model 1 is $R_{it} = \alpha_{i0} + \alpha_{i1}R_{xt} + e_{it}$, *Model 2 is* $R_{it} = \beta_0 + \beta_1 R_{xt} + \beta_2 R_{mt} + e_{it}$, *while Model 3 is* $R_{it} = \gamma_{i0} + \gamma_{i1}R_{xt} + \gamma_{i2}e_{mt} + e_{it}$, *where* $R_{mt} = \delta_{m0} + \delta_{m1}R_{xt} + e_{mt}$. *The dependent variable* R_{it} *is the daily return adjusted for dividend for individual firms.* R_{mt} *is the composite daily return for Taiwan stock market.* R_{xt} *is the change in daily exchange rate of New Taiwan Dollar (NTD) to a foreign currency, which is Euro, Yen or USD. The figure in parentheses in each cell is the p-value. ***, ** and * indicate significance at the 1, 5 and 10 percent levels respectively.*

CONCLUSION

Based on the two-factor model of firm valuation, this study examines the sensitivity of firm value to foreign exchange rate movement using a sample of 110 listed firms in Taiwan. In contrast with most prior studies employing the market portfolio only, this study investigates the firm value sensitivity to exchange rate fluctuation by focusing mainly on individual firms as well as on three industries of high-tech, service, and manufacturing. The exchange rate exposure to the major foreign currencies, including US dollars, Japanese Yen, and Euro is examined separately. Our study finds consistent results of a positive association exists between foreign exchange rate movement and firm value in a small and open economy body of Taiwan, both for the full sample and three industries. Unlike prior studies, our study uses the daily exchange rate change. The application of a short horizon to capture the exchange rate change is expected to provide a more powerful test setting with reduced noise.

We believe that Equation (4) with residual regression is superior to two other testing models by including both direct and indirect effects of exchange rate movement on firm value. We find that, out of our three industries, the high-tech industry is most influenced by fluctuation in exchange rates. Finally, compared to Japanese Yen, the exchange rate exposure to US dollars and Euro is much prominent.

The study includes 110 major listing firms in Taiwan during 2001-2002 in the sample. The small sample size and a relatively short time frame may limit the generalizability of our results. We suggest future studies including more firms with a longer time horizon, and expanding to other small and open economy bodies, such as Hong Kong, Singapore and Korea in the investigation.

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THE EFFECTS OF EXCHANGE RATE VOLATILITY ON SOUTH AFRICA'S TRADE WITH THE EUROPEAN UNION

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ABSTRACT

In this paper we analyze the effects of the real exchange rate volatility on South Africa's trade flows with the European Union over the period 1980 to 2009. Our study uses quarterly trade flows on South Africa's exports and imports and utilizes the bounds testing approach to cointegration, and error-correction model. Our results reveal that imports depend positively on the levels of domestic economic activity and foreign exchange reserves but negatively on relative prices and exchange rate volatility. In addition, exports depend positively on the levels of foreign economic activity but negatively on relative prices and exchange rate volatility. Furthermore, the exchange volatility exerts mixed effects in the short-run and in the long-run.

JEL: F14, F31

KEYWORDS: South Africa, imports, exports, exchange rate volatility, panel cointegration

INTRODUCTION

Despite the sizeable number of studies conducted, no real consensus about the impact of exchange rate volatility on exports has emerged. While a large number of studies find that exchange rate volatility tends to reduce the level of trade, others find either weak or insignificant or positive relationships. For example, Onafowara and Owoye (2008), Byrne, Darby, and MacDonald (2008), Choudhry (2005), Bahmanee-Oskooee (2002), Arize, *et al.* (2000), Arize (1995), Chowdhury (1993), Pozo (1992), and Bahmani-Oskooee and Ltaifa (1992), find evidence for negative effects. According to these scholars, exchange rate volatility may affect exports directly through uncertainty and adjustment costs for risk-averse exporting investors. Further, it may have an indirect effect through its impact on the structure of output, investment and government policy. On the other hand, Doyle (2001), Chou (2000), McKenzie and Brooks (1997), Qian and Varangis (1994), Kroner and Lastrapes (1993), and Asseery and Peel (1991) find evidence for a positive effect for volatility on export volumes of some developed countries because exchange rate volatility makes exporting more attractive to risk-tolerant exporting firms. However, other scholars such as Aristotelous (2001), Bahmani-Oskooee and Payestch (1993), Bahmani-Oskooee (1991), and Hooper and Kohlhagen (1978) have reported no significant relationship between exchange rate volatility and exports.

Reasons for contradictory results by different studies may be due to a variety of factors, among them: different methods used to measure exchange rate volatility; the use of different price deflators; the differential use of sample data, for example, the use of aggregate export data versus sectoral export data; different time-frame periods; ignoring import dependency on intermediate and capital goods of the receiving country, as is the case with many developing countries; and the absence of complex econometric methods for studying these variations. As a result scholars stopped investigating the exchange rate volatility-export nexus by the late 1990's. However, with better access to sectoral data and the development of more sophisticated econometric models, recent studies have begun evaluating the exchange rate volatility-export connection from a sectoral perspective. The rationale behind this is that different trade sectors would be impacted differentially by exchange rate volatility, and therefore may be more revealing than aggregate studies.

This study focuses on aggregate trade flows between South Africa, a developing country, and the member countries of the European Union (EU) to uncover the nature and sensitivity of the relationship between exchange rate volatility and trade flows. We use the recently developed panel unit root and panel cointegration techniques for this purpose. Using this approach we investigate the effects of exchange rate volatility on South Africa's exports to and imports from the EU countries over a period of 30 years using quarterly data from 1980:Q1 to 2009:Q4.

To this end we provide a brief review of the literature in the next section. Thereafter, we lay the empirical framework of our study by specifying our model. In the section following that we discuss variable definitions and outline our data sources. Empirical results from the bounds testing approach to cointegration, and error-correction model estimates are presented in the penultimate section. The final section presents a summary and conclusion of the results obtained in this study.

LITERATURE REVIEW

In this section we present a brief overview of studies that examine the exchange rate volatility-trade nexus. We begin by discussing the most recent and sophisticated studies, employing cointegration techniques and error-correction models, to older, less complex studies.

Bahmani-Oskooee and Hegerty (2009) investigate the effects of exchange rate fluctuations on trade flows between the U.S. and Mexico using disaggregated, industry-level annual export and import data for 102 industries from 1962 to 2004. They analyze both the short- and long-term effects of volatility in the peso/dollar real exchange rate on Mexican-United States trade. They conclude that in the short-term increased volatility negatively affects trade flows in most industries. Long-term effects however, are significant for only one-third of the industries studied, and of this, only two-thirds are negative. They speculate that increased Mexican integration and liberalization of economic policies allow for greater adjustments in the long-term so that volatility is less of a problem in the long-term than in the short-term.

Byrne, Darby, and MacDonald (2008) analyze the impact of exchange rate volatility on the volume of bilateral U.S. trade flows using homogenized and differentiated sectoral annual data over the period 1989-2001 for a cross-section of 6 EU countries and 22 industries. Their study finds that clustering all industries together provides evidence of a negative effect on trade from exchange rate volatility, which confirms findings of other studies using aggregate data. However, when investigating sectoral trade differences, the effects of exchange rate volatility on trade is negative and significant for differentiated goods and insignificant for homogeneous goods, confirming recent studies that sectoral differences are in fact crucial to explaining the differential impact of volatility on trade. They suggest that a greater degree of disaggregation at the industry level may provide more worthwhile results, which is what we do in this study.

Bahmani-Oskooee and Kovyryalova (2008) investigate the effect of exchange rate fluctuations on trade flows between the U.S. and the United Kingdom using disaggregrated annual export and import data for 177 commodities industries from 1971 to 2003. They analyze both the short- and long-term effects of real exchange rate volatility on trade between the U.S. and the UK. Their results reveal that the volatility of the real dollar–pound rate has a short-term significant effect on imports of 109 industries and on exports of 99 industries. In most cases, such effects are unfavorable. In the long run, however, the number of significant cases is somewhat reduced: only 62 import and 86 export industries are significantly and adversely affected by exchange rate volatility. The industries affected involve both durable and non-durable goods, and include small as well as large industries, supporting findings by aggregate studies.

In another study, Bahmani-Oskooee and Mitra (2008), investigate the effects of exchange rate volatility on trade flows between the U.S. and India, an emerging economy. Using annual data from 40 industries from 1962–2004, their results demonstrate that exchange rate volatility has more short-run than long-run effects. In the short-run, 17 industries were affected on the import side and 15 on the export side. The industries affected show India's increasing ability to produce import substitutable goods. However, in the long run, only a few industries are affected because the increasing dependence on trade between India and the US cause industries to respond inelastically to exchange rate volatility.

Using both the nominal and the real exchange rate between the United States dollar and the currencies of Canada and Japan, Choudhury (2005) investigates the influence of exchange rate volatility on U.S. real exports to Canada and Japan using aggregate monthly data ranging from January 1974 to December 1998. The study uses conditional variance from the GARCH (1, 1) model as a measure of exchange rate volatility, and finds significant and mostly negative effects of exchange rate volatility on real exports.

As in the above studies, Sukar and Hassan (2001) investigate the relationship between U.S. trade volume and exchange rate volatility using cointegration and error-correction models. Their study uses quarterly aggregate data covering the period 1975Q1 – 1993Q2 and a GARCH model to measure the exchange rate volatility. Paralleling other studies, the authors find evidence for a significantly negative relationship between U.S. export volume and exchange rate volatility. However, unlike other findings, they reveal that the short-run dynamics of the exchange rate volatility -trade relationship is insignificant. They argue that this result may be due to the existence of avenues for hedging against exchange risks so as to neutralize the negative impact of exchange rate volatility. Other scholars argue that this short-run insignificant relationship may be because of the investigators' use of aggregate data, which ignores sectoral differences. For example, while one sector may exhibit a negative relationship, another may exhibit an equal but opposite effect so that they offset each other.

Arize (1995), using monthly series from February 1978 to June 1986 analyzes the effects of real exchange rate volatility on the proportions of bilateral exports of nine categories of goods from the U.S. to seven major industrial countries. The volatility measure employed is the standard deviation of the monthly percentage change in the bilateral exchange rate between the U.S. and the importing country from *t* to *t-12*. The study reveals differential effects of exchange rate volatility across different categories of exports. The study also concludes that exchange rate uncertainty has a negative effect on U.S. real exports, and that it may have a major impact on the allocation of resources to different industries depending on trade elasticities.

Lastrapes and Koray (1990) analyze the interrelationships among exchange rate volatility, international trade, and macroeconomic variables using the vector autoregression (VAR) model. The model estimates U.S. multilateral trade from 1973 to 1990 and includes a moving standard deviation measure of real exchange rate volatility. While the results reveal some evidence of a statistically significant relationship between volatility and trade, the moving average representation of the model implies a rather small quantitative effect. The study concludes that exchange rate volatility is influenced by the state of the economy, a factor ignored in a variety of other studies.

Klein (1990) is one of the first few scholars to analyze the effects of exchange rate volatility on the proportion of disaggregated bilateral exports of nine categories of goods from the U.S. to seven major industrial countries using fixed effects framework. Using monthly series data from February 1978 to June 1986, the study reveals that in six categories of exports exchange rate volatility significantly affects the volume of exports and in five of these categories the effect is positive, suggesting that real exchange rate volatility may in fact increase exports by risk-taking firms.

Koray and Lastrapes (1989) examine the relationship between real exchange rate volatility and bilateral imports from five countries, namely, the UK, France, Germany, Japan, and Canada, employing a VAR model. The study uses aggregate monthly data over a 17-year period from January 1959 to December 1985, and tests for different effects during both the fixed and the flexible exchange rate regimes. Results suggest that while the effects of volatility on imports is weak, permanent shocks to volatility experience a negative impact on imports. However, those effects are relatively more important during the flexible-rate than the fixed-rate period.

Finally, Cushman (1988) tests for real exchange rate volatility on U.S. bilateral trade flows using annual data from 1974-1983 to study the effects of the floating exchange rate regime on exchange rate volatility. The study finds evidence for significant negative effects in only two of six U.S. export flows with one export flow showing a significant positive effect, confirming other studies of a weak risk-averse effect of exchange rate volatility on exporting firms.

One major problem with most of the studies above is that the sample period includes the period prior to the end of the fixed exchange regime, so results may include the lag effects of fixed exchange rates on trade before 1973 lingering on during the transition period after the implementation of the floating exchange rate regime. The current study corrects for this potential bias by using South African quarterly aggregated trade data covering a 30-year period from 1980:Q1 to 2009:Q4. The methodology used in this study incorporates many of the recent developments in the literature, namely, panel unit roots and panel cointegration and error-correction models, which may uncover the nature and sensitivity of the exchange rate volatility-trade nexus.

METHODOLOGY

Model Specification

The objective of this study is to assess the effects of exchange rate volatility on the aggregated exports and imports flows between South Africa and the member countries of the European Union. Drawing on the existing empirical literature in this area, we specify that a standard long-run export demand function that may take the following form (see, for example, Ozturk and Kalyonku, 2009; Choudhry, 2005; Arize, 1998, 1996, 1995; and Asseery and Peel, 1991):

$$\ln X_t = \beta_0 + \beta_1 \ln Y_t + \beta_2 \ln RP_t + \beta_3 \ln RER_t + \beta_4 \ln VOL_t + \beta_5 D_t + \varepsilon_t \quad (1)$$

where X_t is real export volume of South Africa in period t, Y_t is the real income of foreign country in period t, RP_t is the relative price of exports in period t, RER_t is the real exchange rate between the South African rand and the local currency of a partner country in period t, VOL_t is a measure of exchange rate volatility between the South African rand and the local currency of a partner country in period t, D_t is a dummy variable representing the apartheid era (1980-1994) in South Africa, and ε_t is the error term bounded with classical statistical properties.

Economic theory posits that the real income level of the domestic country's trading partners would have a positive effect on the demand for its exports. Therefore, *a priori*, we would expect that $\beta_1 > 0$. On the other hand, if the relative price of exports rise (fall), domestic goods become less (more) competitive than foreign goods, causing the demand for exports to fall (rise). Therefore, *a priori*, one would expect that β_2, which measures the competitiveness of South African exports relative to the given country's domestic production, is negative. Similarly, if a real depreciation of the South African rand, reflected by a decrease in the RER, is to increase export earnings of industry i, we would expect an estimate of β_3 to be negative.

The last explanatory variable is a measure of exchange rate volatility. Various measures of real exchange rate volatility have been proposed in the literature. Some of these measures include (1) the averages of absolute changes, (2) the standard deviations of the series, (3) the deviations from the trend, (4) the squared residuals from the ARIMA or ARCH or GARCH processes, and (5) the moving sample standard deviation of the growth rate of the real exchange rate. Since the effects of exchange rate volatility on exports have been found to be empirically and theoretically ambiguous (Bredin, *et al.* 2003), β_4 could be either positive or negative.

South Africa is assumed to be a small developing open-economy, rendering it a price-taker with respect to imports, and therefore supports the use of single-equation techniques for estimating the aggregate import demand function. We further assume that only normal goods are imported, and that as a developing country, real foreign exchange reserves comprises an important variable in the function. Further, the apartheid structure and international sanctions have both had a significant impact on aggregate import demand and are therefore included in the model.

The long-run aggregate import demand function for South Africa (in natural logs) is thus specified as:

$$\ln M_t = \alpha_0 + \alpha_1 \ln Y_t + \alpha_2 \ln RP_t + \alpha_3 \ln FR_t + \alpha_4 \ln VOL_t + \alpha_5 D_t + \varepsilon_t \quad (2)$$

where M_t is real import volume of South Africa in period t, Y_i is the real income of South Africa in period t, RP_t is the relative price of imports in period t, FR_t is the real foreign exchange reserves of South Africa in period t, VOL_t is a measure of exchange rate volatility between the South African rand and the local currency of a partner country in period t, D_t is a dummy variable representing the apartheid era (1980-1994) in South Africa, and ε_t is a white-noise disturbance term bounded with classical statistical properties.

The first explanatory variable, Y_t, in the specified model measures real domestic economic activity, or the real GDP of South Africa. Economic theory suggests that income in the importing country is a major determinant of a country's imports and has a positive impact. Thus, *a priori*, it is expected that $\alpha_1 > 0$. The second explanatory variable, the relative price of imports, is calculated as the ratio of import price to domestic price. Economic theory posits that an increase in the relative price of imports discourages imports and therefore α_2 is expected to be negative. The third explanatory variable is a measure of availability of foreign exchange, which can be used to represent the ability to import. Following Hoque and Yusop (2010), we have included the real foreign exchange reserve variable to capture the impact of export earnings on import demand, as export earning is one of the major sources of foreign reserves. This variable does not appear in the traditional import demand function. However, it is an important determinant of imports for developing countries. Since higher real foreign reserves tend to encourage imports, we would expect that $\alpha_3 > 0$. The expected signs of α_1, α_2, and α_3 are borne out in empirical results by Hoque and Yusup (2010), Akinlo (2008), Agbola and Damoense (2005), Narayan and Narayan (2005), Razafimahefa and Hamori (2005), Dutta and Ahmed (2004), Tsionas and Christopoulos (2004), Tang (2004, 2002a), Matsubayashi and Hamori (2003), Gumede (2000), Senhadji (1998), and Mwega (1993), among others. Since the effects of exchange rate volatility on imports have been found to be empirically and theoretically ambiguous, α_4 could be either positive or negative.

Equations (1) and (2) shows the long-run relationships among the dependent and independent variables in our models. They can be represented within a panel setting by incorporating a subscript "i" depicting each of the European Union countries in the sample. This can be represented as follows:

$$\ln X_{i,t} = \beta_{0i} + \beta_{1i} \ln Y_{i,t} + \beta_{2i} \ln RP_{i,t} + \beta_{3i} \ln RER_{i,t} + \beta_{4i} \ln VOL_{i,t} + \varepsilon_{i,t} \tag{3}$$

$$\ln M_{i,t} = \alpha_{0i} + \alpha_{1i} \ln Y_{i,t} + \alpha_{2i} \ln RP_{i,t} + \alpha_{3i} \ln FR_{i,t} + \alpha_{4i} \ln VOL_{i,t} + \varepsilon_{i,t} \tag{4}$$

Panel unit root tests

Before proceeding to cointegration techniques, we need to verify that all of the variables are integrated to the same order. In doing so, we have used the panel unit roots test proposed by Breitung (2005) and Breitung and Pesaran (2008).

Panel Cointegration Tests

We investigate the existence of cointegrating relationship using the standard panel tests for no cointegration proposed by Pedroni (1999, 2004). These tests allow for heterogeneity in the intercepts and slopes of the cointegrating equation. Pedroni's tests provide seven test statistics: Within dimension (panel tests): (1) Panel ν -statistic; (2) Panel Phillips–Perron type ρ-statistics; (3) Panel Phillips–Perron type t-statistic; and (4) Panel augmented Dickey–Fuller (ADF) type t-statistic. Between dimension (group tests): (5) Group Phillips–Perron type ρ-statistics; (6) Group Phillips–Perron type t-statistic; and (7) Group ADF type t-statistic. These statistics are based on averages of the individual autoregressive coefficients associated with the unit root tests of the residuals for each country in the panel. All seven tests are distributed asymptotically as standard normal. Following Pedroni (1999, 2004), the heterogeneous panel and heterogeneous group mean panel of rho (ρ), parametric (ADF), and nonparametric (PP) statistics are calculated as follows:

Panel ν - statistic:

$$Z_\nu = \left(\sum_{i=1}^{N} \sum_{t=1}^{T} \hat{L}_{11i}^{-2} \; \hat{e}_{it-1}^{2} \right)^{-1} \tag{5a}$$

Panel ρ - statistic:

$$Z_\rho = \left(\sum_{i=1}^{N} \sum_{t=1}^{T} \hat{L}_{11i}^{-2} \; \hat{e}_{it-1}^{2} \right)^{-1} \sum_{i=1}^{N} \sum_{t=1}^{T} \hat{L}_{11i}^{-2} \; (\hat{e}_{it-1} \; \Delta\hat{e}_{it} - \hat{\lambda}_i) \tag{5b}$$

Panel *ADF* - statistic:

$$Z_t = \left(\tilde{s}_{NT}^{*2} \sum_{i=1}^{N} \sum_{t=1}^{T} \hat{L}_{11i}^{-2} \; \hat{e}*_{it-1}^{2} \right)^{-\frac{1}{2}} \sum_{i=1}^{N} \sum_{t=1}^{T} \hat{L}_{11i}^{-2} \; \hat{e}_{it-1}^{*} \; \Delta\hat{e}_{it}^{*} \tag{5c}$$

Panel *PP* - statistic:

$$Z_{pp} = \left(\tilde{\sigma}_{NT}^{2} \sum_{i=1}^{N} \sum_{t=1}^{T} \hat{L}_{11i}^{-2} \; \hat{e}_{it-1}^{2} \right)^{-\frac{1}{2}} \sum_{i=1}^{N} \sum_{t=1}^{T} \hat{L}_{11i}^{-2} \; (\hat{e}_{it-1} \; \Delta\hat{e}_{it} - \hat{\lambda}_i) \tag{5d}$$

Group ρ - statistic:

$$\widetilde{Z}_{\rho} = \sum_{i=1}^{N}\left(\sum_{i=1}^{T}\hat{e}_{it-1}^{2}\right)^{-1}\sum_{i=1}^{T}(\hat{e}_{it-1}\ \Delta\hat{e}_{it} - \hat{\lambda}_{i}) \tag{5e}$$

Group *ADF* - statistic:

$$\widetilde{Z}_{t} = \sum_{i=1}^{N}\left(\sum_{i=1}^{T}\bar{s}_{i}^{*2}\ \hat{e}_{it-1}^{*2}\right)^{-\frac{1}{2}}\sum_{i=1}^{T}\hat{e}_{it-1}^{*}\ \Delta\hat{e}_{it}^{*} \tag{5f}$$

Panel *PP* - statistic:

$$\widetilde{Z}_{pp} = \sum_{i=1}^{N}\left(\sum_{i=1}^{T}\hat{\sigma}_{i}^{2}\ \hat{e}_{it-1}^{2}\right)^{-\frac{1}{2}}\sum_{i=1}^{T}(\hat{e}_{it-1}\ \Delta\hat{e}_{it-1} - \hat{\lambda}_{i}) \tag{5g}$$

where

$$\hat{\lambda}_{i} = \frac{1}{T}\sum_{j=1}^{k_i}\left(1-\frac{j}{k_i+1}\right)\sum_{t=j+1}^{T}\hat{\mu}_{it}\hat{\mu}_{i,t-j}\ ;\ \hat{s}_{i}^{2} = \frac{1}{T}\sum_{t=1}^{T}\mu_{it}^{2}\ ;\ \hat{\sigma}_{i}^{2} = \hat{s}_{i}^{2} + 2\hat{\lambda}_{i}\ ;\ \widetilde{\sigma}_{NT}^{2} = \frac{1}{N}\sum_{i=1}^{N}\hat{L}_{11i}^{-2}\hat{\sigma}_{i}^{2}\ ;\ \hat{s}_{i}^{*2} = \frac{1}{T}\sum_{t=1}^{T}\hat{\mu}_{it}^{*2}\ ;$$

$$\widetilde{s}_{NT}^{*2} = \frac{1}{T}\sum_{i=1}^{N}\hat{s}_{i}^{*2}\ ;\ \text{and}\ \hat{L}_{11i}^{2} = \frac{1}{T}\sum_{t=1}^{T}\hat{\eta}_{it}^{2} + \frac{2}{T}\sum_{j=1}^{k_i}\left(1-\frac{j}{k_i+1}\right)\sum_{t=j+1}^{T}\hat{\eta}_{it}\hat{\eta}_{i,t-j}\ .$$

The error terms $\hat{\mu}_{i,t}$, $\hat{\mu}_{i,t}^{*}$, and $\hat{\eta}_{i,t}$ are respectively derived from the following auxiliary regressions: $\hat{\varepsilon}_{it} = \hat{\rho}_{i}\hat{\varepsilon}_{i,t-1} + \hat{\mu}_{it}$; $\hat{\varepsilon}_{it} = \hat{\rho}_{i}\hat{\varepsilon}_{i,t-1} + \sum_{j=1}^{k_i}\hat{\rho}_{ik}\Delta\hat{\varepsilon}_{i,t-j} + \hat{\mu}_{it}$; and $\Delta y_{it} = \sum_{m=1}^{M}\hat{\gamma}_{mi}\Delta x_{mit} + \hat{\eta}_{it}$.

Of the seven test statistics, except for the panel ν - statistic, the other six Pedroni test statistics are left-tailed tests. In order to find evidence for long-run relationship between the variables, the null hypothesis of no cointegration for these tests should be rejected. If the null hypothesis cannot be rejected, there is no long-run relationship between the variables.

DATA SOURCES AND VARIABLES

Our export and import data time series span a 30-year period from 1980:Q1 through 2009:Q4, leading to 120 quarterly observations. Of the 27 member countries of the European Union, only 20 countries were used in the study due to the unavailability of all relevant data. Bulgaria, Cyprus, Estonia, Latvia, Lithuania, Malta, and Slovenia were dropped from the sample, and the rest of the following countries were used in the study: Austria, Belgium, Czech Republic, Denmark, Finland, France, Germany, Greece, Hungary, Ireland, Italy, Luxembourg, Netherlands, Poland, Portugal, Romania, Slovakia, Spain, Sweden, and the United Kingdom. In addition, while the majority of the countries had data for the entire study period, the following five countries had data only for the period specified: Czech Republic, 1993:Q1-2009:Q4; Hungary, 1995:Q1-2009:Q4; Ireland, 1997:Q1-2009:Q4; Poland, 1991:Q1-2009:Q4; and Slovakia, 1993:Q1-2009:Q4. Quarterly data on exports and imports are taken from the International Monetary Fund, *Direction of Trade Statistics Database*. Quarterly nominal data on exports were converted into real export series by deflating them using the export price index with 2005 serving as the base (=100). The export price index series are taken from the International Monetary Fund, *International Financial Statistics Database*.

The real foreign income variable for EU countries is the real quarterly GDP. The underlying series is obtained from the Organization for Economic Cooperation and Development's online database. The real income variable for South Africa is the real quarterly GDP of South Africa. The underlying series is obtained from the International Monetary Fund's *International Financial Statistics database*.

The data on nominal imports, the import price index, foreign exchange reserves series, and domestic price index are taken from the International Monetary Fund's *International Financial Statistics database*. Nominal quarterly imports in local currency are deflated by the import price index (2005 = 100) to obtain the real import variable for South Africa. The real GDP variable is computed in millions of 2005 constant *Rand.* The relative price of imports series is constructed as the ratio of the import price index (2005=100) to domestic price index, as measured by the consumer price index (CPI) (2005=100). To obtain the real foreign reserves series, we deflated the nominal foreign exchange reserves series by the CPI (2005=100).

Following Bahmani-Oskooee and Wang (2008, 2009), and Sekkat and Varoudakis (2000), the real exchange rate, RER_t, is constructed as:

$$RER_t = \left(\frac{ER_t \times P_t^f}{P_t^{SA}} \right) \tag{6}$$

where RER_t is the real exchange rate, ER_t is the bilateral nominal exchange rate between South Africa and a given OECD country at time t, P_t^f is the consumer price index (2005=100) of a given OECD country at time t, and P_t^{SA} is the consumer price index (2005=100) of South Africa at time t. The quarterly data on nominal exchange rates are taken from the IMF, *International Financial Statistics database*.

Finally, we use a commonly used measure of exchange rate volatility in this study, though there are several alternative measures are available. It should be noted at this juncture that there is no unique way to measure real exchange rate volatility. Our measure of volatility is constructed following Bredin, Fountas, and Murphy (2003), Weliwita, Ekanayake, and Tsujii (1999), Chowdhury (1993), Lastrapes and Koray (1990), and Koray and Lastrapes (1989). Following these authors the real exchange rate volatility measure is constructed as:

$$VOL_t = \left[\frac{1}{m} \sum_{i=1}^{m} (\ln RER_{t+i-1} - \ln RER_{t+i-2})^2 \right]^{1/2} \tag{7}$$

where VOL_t is the volatility of real exchange rate, RER_t is the real exchange rate and m = 4 is the order of the moving average. According to Koray and Lastrapes (1989), this measure can capture general movements in real exchange rate volatility and exchange rate risk over time.

EMPIRICAL RESULTS

Panel Unit Root Tests

Table 1 shows the summary statistics of the main variables for the full sample. The starting point of our econometric analysis is to check whether the variables included in Equations (3) and (4) contain panel unit roots. In other words, in Equation (3), we need to check whether [X, Y, RP, RER, VOL] contains a unit root and also whether [M, Y, RP, FR, VOL] contains a unit root in Equation (4). While there are several panel unit root tests are available, this study uses the unit root test developed by Breitung (2000, 2005). These results of the exports model and the imports model are presented in Table 2.

Table 1: Basic Summary Statistics

	Exports Model					**Imports Model**				
	X	**Y**	**RP**	**RER**	**VOL**	**M**	**Y**	**RP**	**FR**	**VOL**
Mean	888.09	175.55	1.82	0.24	0.02	1207.25	1232.22	2.24	47.17	0.02
Median	502.39	163.97	1.40	0.23	0.02	923.59	1125.23	1.52	13.10	0.02
Maximum	3297.96	272.25	3.85	0.31	0.07	4606.02	1824.38	6.25	235.97	0.07
Minimum	71.78	105.42	0.83	0.15	0.00	122.18	914.08	0.79	3.84	0.00
Std. Deviation	795.33	50.89	0.93	0.04	0.01	1062.62	259.76	1.54	58.79	0.01
Skewness	0.99	0.37	0.79	-0.06	1.17	1.07	0.96	1.16	1.59	1.17
Kurtosis	3.06	1.86	2.19	2.15	4.51	3.61	2.71	3.03	4.53	4.51
Jarque-Bera	582.72	276.86	468.00	110.53	923.11	748.81	561.93	802.03	1859.86	923.11
Probability	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00
Observations	2,400	2,400	2,400	2,400	2,400	2,400	2,400	2,400	2,400	2,400

Note: This table shows the summary statistics of the main variables for each of the six regions.

Table 2: Breitung Panel Unit Root Test Results

Exports Model (Model 3)		**Imports Model (Model 4)**	
Variable	**Breitung t-statistic**	**Variable**	**Breitung t-statistic**
Level:		**Level:**	
ln(EXP)	-0.42 (0.337)	ln(IMP)	-1.74 (0.337)
ln(Y)	3.03 (0.997)	ln(Y)	-0.51 (0.307)
ln(RER)	-0.89 (0.185)	ln(FR)	1.22 (0.997)
ln(RP)	10.57 (0.998)	ln(RP)	-1.03 (0.152)
ln(VOL)	-1.68 (0.264)	ln(VOL)	-1.33 (0.254)
First Differences:		**First Differences:**	
Δln(EXP)	-5.46 (0.000)***	Δln(IMP)	-14.94 (0.000)***
Δln(Y)	-15.29 (0.000)***	Δln(Y)	-14.20 (0.000)***
Δln(RER)	-2.44 (0.007)***	Δln(FR)	-27.14 (0.000)***
Δln(RP)	-10.87 (0.000)***	Δln(RP)	-13.17 (0.000)***
Δln(VOL)	-12.43 (0.000)***	Δln(VOL)	-31.14 (0.000)***

*Notes: Breitung indicates the Breitung and Pesaran (2008) Breitung t-test for panel unit root tests. This test examines the null hypothesis of unit root (non-stationary). The figures in parentheses are the p-values. *** indicates the significance at the 1 percent level.*

The panel unit root tests indicate all the variables in both the exports model and the imports model are integrated of order one.

Panel Cointegration Tests

With the respective variables integrated of order one, the heterogeneous panel cointegration test advanced by Pedroni (1999, 2004), which allows for cross-section interdependence with different individual effects, is performed and the results are presented in Tables 3. The results for both within and between dimension panel cointegration test statistics are given in the table. All seven test statistics reject the null hypothesis of no cointegration at the 1% significance level for both exports model and imports model, indicating that the five variable are cointegrated in each model.

Table 3: Heterogeneous Panel Cointegration Results

(a) Exports Model (Eq.3)	
Panel Cointegration Statistics (within-dimension)	**Test Statistic**
Panel ν-statistic	10.378 (0.000)***
Panel ρ-statistic	-18.995 (0.000)***
Panel *t*-statistic	-15.731 (0.000)***
Panel *t*-statistic	-16.038 (0.000)***
Panel Cointegration Statistics (within-dimension)	
Group PP type ρ-statistic	-19.998 (0.000)***
Group PP type *t*-statistic	-18.621 (0.000)***
Group ADF type *t*-statistic	-19.016 (0.000)***
(b) Imports Model (Eq.3)	
Panel Cointegration Statistics (within-dimension)	**Test Statistic**
Panel ν-statistic	55.545 (0.000)***
Panel ρ-statistic	-27.950 (0.000)***
Panel *t*-statistic	-18.261 (0.000)***
Panel *t*-statistic	-17.577 (0.000)***
Panel Cointegration Statistics (within-dimension)	
Group PP type ρ-statistic	-29.616 (0.000)***
Group PP type *t*-statistic	-20.891 (0.000)***
Group ADF type *t*-statistic	-20.050 (0.000)***

*Notes: The number of lag length was selected automatically based on SIC with a maximum lag of 15. The figures in the parentheses are p-values. *** indicates the significance at the 1 percent level.*

After having found consistent evidence of cointegration, the Dynamic Least Squares (DOLS) technique for heterogeneous cointegrated panels is estimated. The results of the DOLS are presented in Table 4 for both exports model and for imports model. As the results for the exports model show, all the coefficients have the expected signs and all the coefficients are statistically significant at the 1% level of significance. Exchange rate volatility variables has a positive sign indicating that exchange rate volatility does not have any adverse impact on South Africa's exports to the European Union. Given that the variables are expressed in natural logarithms, the coefficients can be interpreted as elasticity estimates. The results indicate that, in the exports model, a 1% increase in real foreign income increases real exports by 4.62%; a 1% increase in real exchange rate decreases real exports by 0.19%; and a 1% increase in the relative price decreases real exports by 2.28%. Similarly, as the results for the imports model show, all the coefficients have the expected signs and all the coefficients are statistically significant at the 1% level of significance. Exchange rate volatility variables has a positive sign indicating that exchange rate volatility does not have any adverse impact on South Africa's imports from the European Union. However, in both models, the dummy variable representing the apartheid era has a negative sign indicating that trade flows between South Africa and the European Union declined during the apartheid period.

Table 4: Dynamic Least Squares (DOLS) Long-Run Elasticities

Dependent Var:	**ln(Y)**	**ln(RER)**	**ln(RP)**	**ln(VOL)**	**D**
	4.622***	-0.193***	-2.279***	0.044***	-1.440***
Ln(X)	(9.84)	(3.44)	(6.51)	(5.51)	(7.49)
Dependent Var:	**ln(Y)**	**ln(FR)**	**ln(RP)**	**ln(VOL)**	**D**
	3.013***	0.126***	-0.081***	0.131***	-1.205***
Ln(M)	(9.92)	(8.88)	(3.74)	(9.51)	(8.67)

*Notes: The figures in parentheses are absolute values of t-statistics. *** indicates the statistical significance at the 1 percent level.*

Finally, in order to evaluate the impact of exchange rate volatility on trade flows during the short-run, a panel vector error correction model (VECM) proposed by Pesaran et al. (1999) is estimated. The results of these models are presented in Table 5 for exports model and for imports model. These results show evidence that exchange rate volatility has an adverse impact on exports in the short-run.

Table 5: Dynamic Least Squares (DOLS) Short-Run Elasticities

Dep. Var.	**Δln(Y)**	**Δln(RER)**	**Δln(RP)**	**Δln(VOL)**	**D**	**ECM_{t-1}**
	3.107***	-0.605***	-2.010***	-0.014**	0.011	-0.372***
Δln(X)	(9.29)	(8.44)	(4.68)	(1.99)	(1.17)	(9.67)
Dep. Var.	**Δln(Y)**	**Δln(FR)**	**Δln(RP)**	**Δln(VOL)**	**D**	**ECM_{t-1}**
	1.034***	-0.020	-0.636	0.009	0.031	-0.497***
Δln(M)	(8.70)	(1.39)	(1.53)	(0.90)	(2.87)	(9.52)

*Notes: The figures in parentheses are absolute values of t-statistics. ** and *** indicate the statistical significance at the 1 percent and 5 percent level, respectively. All the variables (except the dummy variable) are expressed in natural logarithm and therefore the coefficients can be interpreted as elasticities.*

SUMMARY AND CONCLUSIONS

In this paper we have examined the dynamic relationship between trade flows and exchange rate volatility in South Africa's trade with the European Union in the context of a multivariate error-correction model. Estimates of the long-run export and import demand functions were obtained by employing the panel cointegration using quarterly data for the period 1980:Q1 - 2009:Q4.

The panel unit root tests indicate all the variables in both the exports model and the imports model are integrated of order one. The heterogeneous panel cointegration test advanced by Pedroni (1999, 2004) was used to test for panel cointegration for both exports and imports models. The results for both within and between dimension panel cointegration test statistics reject the null hypothesis of no cointegration at the 1% significance level for both exports model and imports model, indicating that the five variable are cointegrated in each model.

After having found consistent evidence of cointegration, the Dynamic Least Squares (DOLS) technique for heterogeneous cointegrated panels is estimated. As the results for the exports model show, all the coefficients have the expected signs and all the coefficients are statistically significant at the 1% level of significance. Exchange rate volatility variables has a positive sign indicating that exchange rate volatility does not have any adverse impact on South Africa's exports to the European Union. However, the results show evidence that exchange rate volatility has an adverse impact on exports in the short-run. Similarly, as the results for the imports model show, all the coefficients have the expected signs and all the coefficients are statistically significant at the 1% level of significance. Exchange rate volatility variables has a positive sign indicating that exchange rate volatility does not have any adverse impact on South Africa's imports from the European Union. However, in both models, the dummy variable representing the apartheid era has a negative sign indicating that trade flows between South Africa and the European Union declined during the apartheid period.

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BUBBLE IN THE INDIAN REAL ESTATE MARKETS: IDENTIFICATION USING REGIME-SWITCHING METHODOLOGY

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ABSTRACT

India has a growing economy that can support high-income levels and in turn sustain higher real estate prices. The high prices of Indian real estate seem to be in harmony with its fast growing economy. However, there are concerns about speculative bubble behavior in the Indian real estate market. In this paper, we utilize a sophisticated regime-switching speculative bubble model developed by van Norden and Schaller (1993) along with other traditional econometric methods to test for the presence of bubbles in the Indian real estate market. Our results provide evidence that India real estate bubble was not affected by the 2007-2008 global economic slowdown. The Indian Real Estate market grew from the end of 2008 through early 2011.

JEL: C22, C41, L85, R33

KEYWORDS: Indian Real Estate, Rational speculative bubbles, Regime-switching tests, duration dependence tests, REITs

INTRODUCTION

While the US and European housing markets are still struggling to recover from their pre-global final crisis of 2007-2008, the Indian housing market has recovered quickly from its 2007 slide. RESIDEX, an Index created by National Housing Bank (NHB) of India clearly indicates that except for a few small cities, the prices of real estate are significantly above pre-crisis levels. In H1 2011 prices in the four major Indian cities of Delhi, Mumbai, Kolkata, and Chennai rose 26%, 75%, 111%, and 118 %, respectively. The Reserve Bank of India - RBI (Central bank of India) observed the higher real estate prices and became concerned about a potentially damaging real estate bubble (Business Standards, July 2011).

The RBI based its concern on two trends: (1) the growth of non-performing assets in residential mortgages and that commercial real estate was at record high levels (2) teaser rates offered by Indian banks since 2009 may have over-stimulated loan demand. The teaser rates are set at fixed low rates for the first few years of the loan period and then, depending on the contract, the rate can become floating or remain fixed based on the State Bank Advance Rate – Benchmark Lending Rate. The RBI's main concern is that teaser rates may cause borrowers to default on loans when interest costs are increased. Jones Lang LaSalle, the leading Indian real estate consultants group has noted the higher level of risk and the bubble behavior of investors and borrowers.

Indian economists and policy makers believe that the current Indian real estate market demonstrates similarities to the U.S. real estate market shortly before the subprime mortgage crisis. However, policy makers do not consider the situation alarming at this point because of several factors. First, property prices in some Indian cities are much higher than the sustainable level of income but fast growing economy and rising income levels may be able support high real estate prices. Second, the cause of the real estate bubble burst in U.S. and other Western countries was the widespread bundling and sale of real estate mortgages in the form of financial derivatives. Financial derivatives are not available in the Indian economy. Finally, Indian savings rates are significantly higher than U.S. savings rates during the time of rising real estate prices. Higher savings rates allow borrowers to put a greater percentage of money down on a property and provide a greater cushion for borrowers in the event that mortgage payments are adjusted upward by lending institutions.

The Indian economy depends heavily on the real estate sector. Therefore, any destabilization can be disastrous to the Indian economy and would create an immediate and long-term recession that will be very hard to overcome (Vishwakarma and French (2010) and Newell and Kamineni (2007) provides a comprehensive sketch of historical background and significance of real estate sector in India.). According to the "Report on Trend & Progress of Housing in India, 2006" by the National Housing Bank of India: "The real estate sector is one of the fastest growing sectors in India with its size close to US $12 billion and with an annual growth rate of 30% with 5% contribution to GDP".

Bubbles in the real estate markets are characterized by rapid increases in the values of real property because market participants are willing to buy and hold the asset because they believe that they can resell the asset at an even higher price in the future. This is possible because of increasing prices in the case of a positive bubble until unsustainable price levels are reached relative to incomes and other economic fundamentals. A similar result occurs for negative bubbles.

Several studies attempt to detect the presence of bubbles in different real estate markets. Arshanapalli and Nelson (2008) find evidence of bubbles in the U.S. housing market during the 2000 to 2007 period. Bordo and Jeanne (2002) and Helbling (2003) find evidence of housing bubbles in 14 OECD countries over the period of 1973-2000. Abraham and Hendershott (1996), Case and Shiller (2003) and Cecchetti (2005) present evidence for the presence of a speculative bubble in the U.S. housing market. Roche (2001) finds some evidence of a speculative bubble in the Irish housing market in Dublin. Clayton (1997) studies the Canadian real estate market and reports evidence of bubble behavior. Levin and Wright (1997) divided housing prices into two components for UK housing market, one driven by economic fundamentals and the other due to speculation in the housing market. They find strong evidence of bubble behavior in the UK housing market. Hendershott (2000) and Bjorklund and Sodeberg (1999) find evidence of bubbles in the commercial real estate market in Sydney and Sweden, respectively. Qin (2005) tests for bubbles in Seoul and Hong Kong markets and finds some evidence of bubbles lasting a few years in both markets. Paskelian, Hassan and Huff (2011) test for the presence of rational speculative bubbles in the U.S. REITs market and find evidence of bubble-behavior using regime-switching methodology. Also, Paskelian and Vishwakarma (2011) test for the presence of bubbles in the Chinese real estate market and find evidence of a growing real estate bubble. Finally Joshi (2006), attempts to isolate the factors affecting the real estate bubble in the Indian property market using VAR framework. He finds that interest rate levels and credit growth play an important role in influencing housing prices as well as stabilizing other sectors in the Indian economy.

In this paper, we employ various econometric methods to test for the presence of rational speculative bubbles in the Indian REITs market. Our study focuses on the methodology of van Norden (1996), and van Norden and Schaller (1999). The regime-switching model proposed by van Norden and Schaller tests for periodically collapsing speculative bubbles by estimating time-varying probability of the collapse of positive and negative speculative bubbles. This paper is the first paper that examines data for the presence of rational speculative bubbles in Indian real estate markets using the regime-switching model. We find a relatively strong presence of bubbles in the Indian REITs market. In particular, our findings suggest that bubble behavior in the Indian REITs market intensifies at the end of 2007 and continues to increase in intensity through the end of 2010.

In the succeeding sections of this paper, we present a brief review of literature dedicated to the behavior and existence of real estate bubbles. We describe the data and the methodology used in the paper, including the advantages and disadvantages of the econometric techniques employed in the analysis. The results section follows with explanations of the analysis and test result interpretation. Finally, the last section is a summary of conclusions.

LITERATURE REVIEW

The econometric literature describes several methods to test for the presence of bubbles in asset prices. The three major techniques for testing bubbles are variance bound tests, stationarity and cointegration tests, and regime switching models. Hart and Kreps (1986) were first to test for bubbles in stock markets using excess volatility and variance-bound tests. Shiller (1981) and Cochrane (1992) used similar procedures to test the movement of stock prices. These tests compared actual data with fundamentals to find evidence of a speculative bubble. However, finding the appropriate fundamental is a major challenge. Stationarity and cointegration tests are proposed by Diba and Grossman (1988a and 1988b), and Hamilton and Whiteman (1985), to establish the presence of speculative bubbles. According to this methodology, asset prices that are cointegrated with their dividends are evidence against the presence of bubbles. However, these methods tend to reject the presence of bubbles too often (Evans, 1991) because other factors tend to cause a lack of cointegration.

Regime-switching models were introduced by Blanchard and Watson (1983). These models test for the presence of bubbles as changes in regimes occur, and then analyze price process properties in out-of-the bubble regimes. van Norden and Schaller (1993) and Schaller and van Norden (1997) refined this model by formulating a periodically collapsing, positive and negative speculative bubble model that has a time varying probability of collapse (see more detail in methodology section).

Real estate research has incorporated the above-mentioned techniques with some improvisation i.e. Brooks, Katsaris, McGough and Tsolacos (2001) for UK market, Jirasakuldech, Campbell, and Knight (2006) for US market, Payne and Waters (2005), Payne and Waters (2007), Waters and Payne (2007), and Paskelian et al. (2011).

Brooks, Katsaris, McGough and Tsolacos (2001) applied excess volatility and variance-bound methods to test for the presence of speculative bubbles in the real estate market of the UK. They designated the dividend growth rate taken from Gordon's model in order to test for the presence of a speculative bubble. They found evidence of a UK real estate bubble due to the presence of low volatility in fundamental measures rather than real estate prices.

Jirasakuldech et al. (2006) looked for the presence of rational speculative bubbles in the US securitized real estate market from January 1973 to December 2003 using four methods. They used various factors for calculation of fundamentals e.g., the consumer price index (CPI), industrial production, the federal funds rate and the default risk premium. First, they applied unit root test following Diba and Grossman (1988a, 1988b) but found no evidence of a speculative bubble. Second, they used Blackburn and Sola (1996) methodology to conduct cointegration tests. Again they found no evidence of a bubble. Their third test was Johansen's cointegartion test with Junttila's (2003) methodology that resulted in a finding of no bubble evidence. Lastly, they applied the McQueen and Thorley (1994) model for duration dependence test and found no evidence of a bubble.

Payne and Waters (2005, 2007) argue that the Diba and Grossman (1988a) approach will not detect periodically collapsing bubbles. Payne and Waters (2005) followed the methodology of Evans (1991) finding some evidence of negative periodically collapsing bubbles in mortgage and hybrid REITs markets. However, Payne and Waters (2007) find mixed evidence of bubble in equity REIT markets. Conversely, Waters and Payne (2007) test for both positive and negative periodically collapsing bubbles using the residual-augmented Dickey-Fuller model of Taylor and Peel (1998) and the momentum threshold autoregressive (MTAR) model of Enders and Siklos (2001); which resulted in a finding of some evidence of negative periodically collapsing bubbles in the mortgage real estate market.

An exhaustive study conducted by Paskelian et al., (2011) on equity, mortgage and hybrid real estate markets in the US from 1972-2009 employed the unit-root test, variance ratio test, duration dependence test and regime switching regression models of van Norden and Schaller (1993). Study results revealed that only the regime-switching model showed weak evidence of a periodically collapsing bubble in the mortgage and hybrid REITs sectors during the period 2000-2005. However, traditional methods employed in the same study failed to detect the existence of any speculative bubble.

This paper adds to the current literature by providing insights into the presence of periodically collapsing speculative bubbles in the Indian securitized real estate market. The study employs a regime-switching bubble methodology not previously utilized in existing Indian securitized real estate bubble literature.

DATA AND METHODOLOGY

This paper uses daily and monthly data from the CNX realty index from January 2007 to July 2011 as a proxy for the Indian real estate sector. The daily CNX realty index is used for the regime-switching methodology; while the monthly CNX realty price index is employed for all other estimations. The CNX realty index is a sector index of Indian real estate maintained by the National Stock Exchange (NSE) of India. It is comprised of ten listed companies from the real estate sector with considerable track record. Monthly Indian rental property series data and consumer price index data was obtained from the Labour Bureau of the Government of India.

Gurkaynak (2005) provides a comprehensive reference concerning the econometric tests for speculative bubbles. Jirasakuldech et al. (2006) analyze the U.S. Equity REITs market by testing for the presence of rational speculative bubbles using the unit root, variance ratio and duration dependence methodologies. Following Jirasakuldech et al. (2006), we test for stationary properties of the Indian CNX Realty price index utilizing a fundamental variable measured by the Indian rental index series that tests for the existence of rational speculative bubbles. The Indian rental index is used as the proxy for the fundamental process. If the Indian CNX Realty price index contains bubbles, then the explosive nature of speculative bubbles will make it less likely for the series data to achieve a stationary process by repeatedly differencing the series. Therefore, if the Indian CNX Realty prices are stationary in the first differences, it is more likely that the non-stationary is caused by the nature of market fundamentals, rather than explosive bubbles. We test for the stationarity of the Indian CNX Realty price index returns by using the Augmented Dickey Fuller (ADF) (1979) and the Phillips Perron (PP) (1988) unit root tests.

The variance ratio test has been widely used to test for weak-form efficiency of financial markets since Lo and MacKinlay (1988). A comprehensive survey of recent methodological developments is given in Charles and Darne (2009). We use the following procedure to calculate the variance of an asset's return over the holding period k, denoted as Vk. We define the variance ratio V(k) as the ratio of the variance of the k-period return to that of one-period return times k.

Next, we follow the popular McQueen and Thorley's (1994) duration dependence methodology to test for rational speculative bubbles in the Indian real estate index. The continuously compounded monthly real return as the first difference of the natural log normalized using the first difference of the natural log of the Indian monthly consumer price index (CPI) is used for the duration dependence test.

The duration dependence procedure tests for the presence of rational speculative bubbles by looking at the length sequence of positive returns. To sustain a bubble the probability of negative return decreases with the increase of length of sequence of positive returns. Technically, this means the presence of bubbles creates negative duration dependence and decreasing hazard rate. Therefore, to test for rational speculative bubbles, we need to examine the hazard rate h_i for runs of positive and negative returns. If condition of $h_{i+1} < h_i$ exists in runs of positive returns then it means bubble exists. Where $h_i = Prob\ (\epsilon_t < 0 | \epsilon_{t-1} > 0, \epsilon_{t-2} > 0 \ldots \ldots, \epsilon_{t-i} > 0, \epsilon_{t-i-1} > 0)$ rate (McQueen and Thorley (1994); Jirasakuldech, Campbell and Knight (2006)).

Finally, in order to circumvent the shortcomings of conventional econometric tests for the existence of speculative bubble, and continuing with the line of research started by Payne and Waters (2005, 2007), we test for the presence of periodically partially collapsing speculative bubbles based on the regime-switching methodology of van Norden and Schaller (1993). van Norden (1996) and van Norden and Schaller (1993 and 1999) show that periodically collapsing bubbles incorporate regime switching processes in asset returns which may cause the bubbles to either survive or collapse. Schaller and van Norden (1997) show that the probability of the collapse depends on the size of the bubble, thus switches in regime can be predicted using a measure of the size of the bubble in the previous period.

The paragraph below is a brief overview of the van Norden and Schaller (1993) model and estimation procedure. The time varying probability of collapse is measured by the model which incorporates periodicallity, partial collapsibility as it pertains to the positive and negative speculative bubble model and is stated as follows:

$$E_t(b_{t+1}) = \begin{cases} \frac{(1+i)b_t}{q(B_t)} - \frac{1-q(B_t)}{q(B_t)}\, u(B_t)P_t \ with\ Probability\ of\ q(B_t) \\ \\ u(B_t)\ P_t\ with\ probability\ of\ 1 - q(B_t) \end{cases} \tag{1}$$

$E_t(b_{t+1})$ is the expected size of bubble and $u(B_t)$ is the relative size of bubble. Since this model allows negative bubbles the probability of bubbles ($q(B_t)$) survival is modeled as a negative function of the size of bubble i.e. $q(B_t)$ as a function of $\frac{\partial q(B_t)}{\partial |B_t|} < 0$. Gross returns in surviving and collapsing state can be written as follows:

$$E_t(r_{t+1} | W_{t+1} = S) = \left[M(1 - B_t) + \frac{MB_t}{q\ (B_t)} - \frac{1-q(B_t)}{q\ (B_t)}\, u(B_t)\right] with\ probability\ q\ (B_t) \tag{2}$$

$$E_t(r_{t+1} | W_{t+1} = C) = [M(1 - B_t) + u(B_t)] with\ probability\ 1 - q(B_t) \tag{3}$$

Where r_{t+1} is the return of period t+1 conditioned on the survival state S and collapsing state C; W_t is an unobserved indicator that determines the current state at time t using probit model
$P(W_{t+1} = S) = q(B_t) = \Omega(\beta_{q,0} + \beta_{q,b}|B_t|)$, and M is the gross fundamental return on asset. After linearizing, the estimable linear regime-switching model becomes:

$$r^S_{t+1} = \beta_{S,0} + \beta_{S,b}B_t + u^S_{t+1}$$
$$r^C_{t+1} = \beta_{C,0} + \beta_{C,b}B_t + u^C_{t+1}$$
$$P(W_{t+1} = S) = q(B_t) = \Omega(\beta_{q,0} + \beta_{q,b}|B_t|) \tag{4}$$

Where u^S_{t+1} and u^C_{t+1} are the unexpected returns for period t+1 in the surviving and collapsing regimes respectively and are assumed to have zero mean and constant variance i.i.d. normal random variables. Four restrictions are imposed in case of periodically collapsing speculative bubbles. The first restriction $\beta_{S,0} = \beta_{C,0} = \beta_0 = \beta_{S,b} = \beta_{C,b} = \beta_{q,b} = 0$ implies that the mean across the two regimes is different. The second restriction is $\beta_{C,0} < 0$ which implies the expected return should be negative if the collapsing regime is observed. The third restriction is $\beta_{S,b} > \beta_{C,b}$ which implies the bubble yields higher (lower) returns if a positive (negative) bubble is observed in the surviving regime than if it is observed in the collapsing regime. The fourth restriction is $\beta_{q,b} < 0$ which implies the probability of the bubble continuing to exist decreases with the size of increasing bubble.

This study also tests the above model in three different specifications: Volatility of regimes, mixture of normals, and mean reversion specifications. Volatility of regimes specifications is achieved by applying ARCH (Auto regressive conditional heteroskedasticity) on the estimation by imposing condition of $\beta_{S,0} = \beta_{C,0} = \beta_0 = \beta_{S,b} = \beta_{C,b} = \beta_{q,b} = 0$ and $\sigma_s \neq \sigma_c$. Mixture of normals is achieved by $\beta_{S,b} = \beta_{C,b} = \beta_{q,b} = 0$. This specification tests the leverage effect of markets. Mean reversion can be achieved by $\beta_{S,0} = \beta_{C,0} = \beta_0 = \beta_{S,b} = \beta_{C,b} = \beta_1 \; and \; \beta_{q,b} = 0$. This specification can capture the linearly predictability of returns with different mean across regimes.

For the relative bubble size, we use van Norden and Schaller (1993) approximation: $B_t = \frac{b_t}{p_t} = 1 - \frac{\rho d_t}{p_t}$ with ρ being the sample mean of the price over rent ratio.

EMPIRICAL RESULTS

Table 1 provides descriptive statistics of daily real returns of the Indian realty price index. During the period 2007 to July 2011, the Indian realty price index provided an average daily real return of 0.052%, with a standard deviation of 0.083%. The corresponding monthly average return was 1.56%. The return of the Indian CNX realty price index shows peaks of maximum and minimum returns with significant negative skewness and excess kurtosis which indicates the potential for the existence of a rational speculative bubble. Further indication of the existence a of speculative bubble can be obtained from significant six month lags (Q6) and twelve month lags (Q12) autocorrelation results in Table 1.

Table 1: Summary Statistics for the Indian CNX Realty Price Index Returns

Variable	**CNX Realty Index Returns**
Mean (%)	0.0519
Median (%)	0.0832
Minimum (%)	-12.202
Maximum (%)	17.745
Standard Deviation (%)	1.934
Skewness	-0.3833*** (0.000)
Kurtosis	8.575*** (0.000)
Jarque-Bera	42.317*** (0.000)
Q(6)	9.268*** (0.0027)
Q(12)	18.241** (0.0147)

*The returns are continuously compounded. The mean, the median, the standard deviation, the minimum and the maximum are expressed in percentage terms. Q9(6) and Q(12) are the Ljung-Box portmanteau test statistics for 6 and 12 autocorrelations. ***, ** and * indicate significance at 1, 5 and 10 percent levels respectively.*

After preliminary investigation of normality and serial autocorrelation, the paper examines the stationarity of the Indian realty price returns by using the Augmented Dickey Fuller (ADF) (1979) and the Phillips Perron (PP) (1988) unit root tests. The ADF and PP unit root tests are applied to the Indian realty price index after normalizing the monthly price levels by the monthly Indian consumer price index to obtain real price levels. Table 2 reports the ADF and PP unit root tests for the Indian realty price index (Panel A) and changes in the Indian realty price index levels (Panel B) with trend as well as without trend. Results from panel A indicate show non-stationarity, however, panel B is stationary. This implies there is no speculative bubble in the Indian real estate market. Had there been some kind of speculative bubble, differencing the series cannot make it stationary (Campbell and Shiller (1988) and Diba and Grossman (1988a and 1988b)). However, the application of stationarity test methodology is not sufficient to reach meaningful conclusions about the existence or lack of existence of speculative bubbles.

Table 2: The ADF and PP Unit Root Tests on Monthly Real Returns of The Indian CNX Realty Price Index

	ADF without Trend	**ADF with Trend**	**PP without Trend**	**PP with Trend**
Panel A: Real Price Index Level				
CNX Realty Index Returns	4.284	0.9582	4.845	1.052
Panel B: Change In Real Price Index Level				
CNX Realty Index Returns	-5.325***	-6.843***	-11.585***	-17.922***

The table shows the estimates of the Augmented Dickey Fuller (ADF) regression: $\Delta Y_t = a_0 + \gamma Y_{t-1} + a_2 t + \sum_{i=1}^{p} \beta_i \Delta Y_{t-1} + \varepsilon_t$ *, where*$\Delta Y_t = Y_{t-i} + Y_{t-i-1}$ *and t is the time period. If* $\gamma_0 = 0$ *, then the* Y_t *series has a unit root, indicating that he series is non-stationary. The Phillips-Perron regression is* $Y_t = \widetilde{a_0} + \widetilde{a_1} Y_{t-1} + \widetilde{a_1}\left(t - \frac{T}{2}\right) + \varepsilon_t$ *, where T is the number of observations. If* $a_1 = 1$ *then* Y_t*has unit root. ***, ** and * indicate significance at 1, 5 and 10 percent levels respectively.*

The variance ratio test (which is a test for random walk hypothesis) is an indirect way of testing for speculative bubbles. Table 3 reports the variance ratio test results for the Indian realty index and corresponding Z-statistics for various lags. The real returns are greater than one for all lags and seem to be increasing with lags, indicating some form of positive autocorrelation or mean reversion which is an indication of rejection of the applicable random walk hypothesis. Rejection of the random walk hypothesis further confirms the absence of bubbles in the Indian real estate market.

Table 3: The Variance Ratio Test On Monthly Real Returns Of The Indian CNX Realty Index Returns.

	2-Month	**Z-Statistic**	**4-Month**	**Z-Statistic**	**8-Month**	**Z-Statistic**	**16-Month**	**Z-Statistic**	**32-Month**	**Z-Statistic**
CNX Realty Index Returns	1.124*	(2.963)	1.151*	(2.931)	1.365	(1.501)	1.368**	(6.105)	1.368**	(3.158)

The table shows the estimates of the variance ratio test given by the following model: $VR(q) = \frac{\bar{\sigma}_c^2(q)}{\bar{\sigma}_a^2}$ *, where* $\bar{\sigma}_a^2$ *is the estimated variance of the monthly differences* $X_t - X_{t-1}$ *, and* $\bar{\sigma}_c^2(q)$ *is the unbiased estimation of 1/q times the variance of* $X_t - X_{t-q}$ *. Under the random walk null hypothesis, the variance ratio is 1 and the test statistic z(q) follows a standard normal distribution asymptotically. ***, ** and * indicate significance at 1, 5 and 10 percent levels respectively.*

Table 4, reports the results of duration dependence tests to detect the possibility of rational speculative bubbles. For this purpose, hazard function (h_i) statistics for the actual number of positive and negative runs for monthly real returns for the Indian realty price index are calculated. Following McQueen and Thorley's (1994) duration dependence methodology, the study uses continuously compounded monthly real returns as the first difference of the natural log normalized using the first difference of the natural log of the monthly Indian consumer price index (CPI) as the data to be used for the duration dependence test.

Table 4: The Duration Dependence Test Results for The Indian CNX Realty Index Returns

Positive Runs				Negative Runs			
Variable	**Run Length**	**Actual Run**	**Sample Hazard Rate**	**Variable**	**Run Length**	**Actual Run**	**Sample Hazard Rate**
	1	12	0.3521		1	5	0.2674
	2	8	0.2588		2	2	0.3684
	3	3	0.2515		3	2	0.3847
	4	4	0.2141		4	1	0.1576
	5	3	0.4216		5	2	0.3654
	6	1	0.4285		6	0	0.000
	7	1	0.3333		7	0	0.0000
	8	0	0.0000		8	0	0.0000
	9	0	0.0000		9	0	0.0000
	10	0	0.0000		10	0	0.0000
Total Runs		32		Total Runs		12	
Log-Logistic Test				**Log-Logistic Test**			
α		-0.5217		α			-0.4147
β		-0.1284		β			0.3847
LRT of H_1: $\beta = 0$		0.5714		LRT of H_1: $\beta = 0$			0.6581
(p-value)		(0.0521)		(p-value)			(0.0524)

*The hazard function (h_i), defined as $h_i = Prob(I = i|I \geq i)$, represents the probability that a specific run ends at length i, provided that it lasts until length i. The log likelihood expression of the hazard function is defined as: $h_i L(\theta|S_T) = \sum_{i=1}^{\alpha} N_i Lnh_i + M_i Ln(1-h_i) + Q_i Ln(1-h_i)$, the estimated hazard rate for length i is derived by maximizing the log likelihood function with respect to h_i.***, ** and * indicate significance at 1, 5 and 10 percent levels respectively.*

From Table 4, the longest positive run lasts 12 months for the Indian realty price index. The negative runs are typically shorter in length. The longest negative run lasts 5 months for the Indian realty price index. The sample hazard rates reported in Table 4 determine the probability that a specific run ends at length i, given that the run has lasted until i. For the Indian realty price index, there are 32 positive runs and 12 negative runs for a total of 44 runs of real returns. The hazard rate associated with a positive run length of 7 months for the Indian realty index is 0.3333. Therefore, there is a probability of 33.33% that a bubble will burst at the 8 month mark. Our results show that there is a moderately increasing pattern in the hazard rate of the Indian realty price index, but there is no such evidence for negative runs. These findings are supportive evidence for the presence of rational expectations bubbles. The significantly positive beta found in the Indian realty price index indicates that as the sequence length of positive returns increases, the probability that the positive run will end increases. Therefore, the presence of bubble behavior using the duration dependence test is confirmed.

The literature review section and methodology section review of the regime-switching methodology of van Norden and Schaller (1993) revealed many advantages over unit root tests, variance ratio tests, and duration dependence tests. Thus, the paper incorporates a regime-switching estimation model for the Indian real estate market. The estimated parameters of the speculative regime-switching model are presented in Table 5 along with the likelihood ratio test statistic for the restrictions implied by the volatility regime, mixture of normals, and mean reversion models of stock returns. The volatility regime model tests for equal variance in two regimes. Failure to reject means no evidence of speculative bubbles. Table 5 shows that the volatility regime specification has a value of 14.252 which is significant at 1%. Therefore, the model rejects the null hypothesis which results in a conclusion that there is some evidence of speculative bubbles in the Indian real estate market. The mixture of normals model specification tests the deviation from the fundamentals which are not simply related to the leverage effect while still having predictive power for the distribution of returns. Again, in this case we find an estimated value of 12.361 which is significant at 1 %. This result supports the conclusion that bubble-like behavior exists in the Indian real estate market. Finally, the mean reversion model tests for a possible relationship between the return process and the B_t. Rejection of the null hypothesis implies a different return process which is indicative of the presence of a speculative bubble. The mean reversion statistic has a value of 14.586 which is significant at 1%, thus this model also demonstrates evidence of the existence of speculative bubbles in the Indian real estate market.

The coefficient restrictions of the model are highly significant and correctly predicted for the Indian real estate market. The point estimate of $\beta_{S,0}$ is 1.085 and is significant at 1% implying a monthly rate of return of 8.5%. Conversely, the point estimate of $\beta_{C,0}$ is 0.978 and is significant at 1% implying a monthly rate of return of -2.2% in the collapsing regime. The difference in the monthly rate of return between the two regimes indicates the presence of speculative bubbles in the Indian real estate market. Therefore, based on highly significant estimates for three stylized specifications, significant coefficient restrictions, and large-sized differences in the respective monthly rates of return of the surviving and collapsing regimes, it can be concluded that strong evidence confirms the existence of speculative bubble behavior phases in the Indian real estate market.

Table 5 Results of the Regime-Switching Speculative Bubble model for the Indian CNX realty Index Returns

Variable	Equity REITs
$\beta_{S,0}$	1.085***
$\beta_{S,b}$	0.005
$\beta_{C,0}$	0.987***
$\beta_{C,b}$	-0.0364
$\beta_{q,0}$	3.158**
$\beta_{q,b}$	-4.028**
σ_S	0.105***
σ_C	0.154***
Log-Likelihood	584.638
AIC	-5.158
SC	-8.529
HQ	-4.325
$\beta_{N,0} \neq \beta_{S,0} \neq \beta_{C,0}$	5.185**
$\beta_{C,b} < 0$	1.521
$\beta_{S,b} > \beta_{C,b}$	5.236**
$\beta_{S,Y} > 0$	7.415**
Volatility Regimes	14.252***
Mixture of Normals	12.361***
Mean Reversion	14.586***

The table shows the estimates of the following model: $r^S_{t+1} = \beta_{S,0} + \beta_{S,b}B_t + u^S_{t+1}$ and $r^C_{t+1} = \beta_{C,0} + \beta_{C,b}B_t + u^C_{t+1}$; with $P(W_{t+1} = S) = q(B_t) = \Omega(\beta_{q,0} + \beta_{q,b}|B_t|)$. r_{t+1} *denotes the return of period t+1 conditioning on being in the surviving (S) or collapsing (C) state and t,* Ω *is the standard normal cumulative density function,* $\beta_{S,0}, \beta_{S,b}, \beta_{C,0}, \beta_{C,b}$ *are the coefficients to be estimated,* u^S_{t+1} *and* u^C_{t+1} *are the error terms with mean zero and variance* σ^C_u *and* σ^S_u *. ***, ** and * indicate significance at 1, 5 and 10 percent levels respectively.*

Our results provide evidence of the presence of bubbles in the Indian real estate market. The regime-switching model coefficients are highly significant and are indicative of speculative bubble behavior in the Indian real estate market. The test results rule out regime shifts based on volatility, leverage effects or linear predictability of returns.

CONCLUSION

The Indian real estate sector suffered a slight slowdown due to the global financial crisis of 2007-2008. However, and in contrast to its western counterparts, 2011 Indian real estate prices are already above their pre-crisis benchmark. Some economists and the central bank of India are viewing the combination of extraordinarily high prices in the real estate sector and low house financing rates as indicative of a real estate bubble in India. Other economists and property dealers argue that the fast growing economy of India has rapidly growing income levels and relatively high savings rates which serve to distinguish India from other countries suffering from the effects of the recent financial crisis. Thus, they argue that real property price increases in India are perfectly normal and not the result of the existence of speculative bubbles. The unique characteristics of the Indian economy and the real estate sector provide an opportunity to study the Indian real estate sector and test for the existence of speculative bubbles.

Our study is one of the first attempts to analyze and test for the existence of speculative bubbles in the Indian real estate sector. The paper tests for the presence of rational speculative bubbles in the Indian real estate market during the period 2007-2011 by studying the CNX Realty Index maintained by the National Stock Exchange (NSE) of India. Several econometric bubble identification techniques are used including the unit root tests, the variance ratio test, the duration dependence test and a regime-switching test. The regime-switching test is based on van Norden-Schaller (1993) methodology.

The conventional techniques employed provide no conclusive evidence of the existence of a speculative bubble in the Indian real estate market. However, the regime-switching test provides conclusive evidence of such a bubble. The volatility regime specification has a value of 14.252 with 1% significance indicating there are two distinct regimes in the Indian REITs sector which differ in more than their variances; thus indicating the presence of two distinct and opposite behaviors in the Indian REITs sector. The mixture of normals statistic has a value of 12.361 with 1% significance which indicates that the Indian REITs prices have significant deviations from their fundamentals values during the period of 2007-2011. This provides evidence of bubble-like behavior in the Indian REITs sector. Finally, the mean-reversion statistic has a value of 14.586 with 1% significance which indicates the presence of two distinct regimes with different constant and autoregressive terms and different volatility. The difference in the monthly rate of return between the surviving and collapsing regimes also indicates the presence of speculative bubbles in the Indian real estate market.

The results from our study have significant policy implications for the Indian government as well as for practitioners. For the Indian government, it is necessary to devise a solution to softly diffuse the speculative bubble without creating an anti-investing sentiment in the markets or alienating the real property owners. Practitioners should not rely on conventional tools for detecting bubbles. Global portfolio holders should incorporate the existence of the Indian real estate bubble into their investment strategy. The Indian government most probably will take some actions to stop bubble growth or burst which will affect any portfolio including REITs shares. Hence, an appropriate hedging strategy should be devised well in advance.

We have provided a thorough analysis of the Indian real estate sector by analyzing the CNX realty index. However, we do not pretend to have written a flawless paper. The paper has some limitations such as unavailability of suitable proxy for Indian real estate market; none of realty indices goes back beyond year 2007. It would have been optimal to include other proxies of real estate market thereby providing a more complete coverage of the Indian real estate sector. We believe that we were consistent in our work and accurate, in which the results are robust in all material respects. To check the robustness of our results, we used several different econometric specifications. An extension of our study can be done using hand collected data from major cities of India to test the behavior of real estate market in different cities of India. In doing so, our conclusions can be stronger and the results more robust.

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INTERNATIONAL VOLATILITY TRANSMISSION OF REIT RETURNS

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ABSTRACT

This study examines whether volatility of REIT returns can transmit across national borders. Two competing hypotheses are proposed. The first is the Transportable Risk Hypothesis which suggests geographic risk can be transmitted overseas if the general equity and real estate securities markets are integrated internationally. The second is the Non-Transportable Risk Hypothesis which argues that geographic risk factors are country-specific and therefore not transmittable across national borders. Using GARCH and EGARCH econometric models, international spillovers of volatility of REIT returns are found among United States, United Kingdom, and Japan. The finding has major implications for formulating international portfolio strategies as it improves forecasting ability. The finding also implies that better international portfolio diversification can be achieved with real estate securities from countries that have a lower degree of integration between the real estate sector and the general stock market.

JEL: C51**;** G11; G15

KEYWORDS: REIT volatility, multivariate GARCH, volatility spillovers, international portfolio diversification

INTRODUCTION

The progressive integration of international capital markets and the globalization of investor portfolios have provided the impetus to understand the dynamics of stock prices across national borders. Academic studies have found substantial evidence of transmission of stock returns volatility across international equity markets (e.g. King and Wadhwani (1990), Hamao, Masulis, and Ng (1990), Lin, Engle, and Ito (1994), Karolyi (1995), Bekaert and Harvey (1997)). The phenomenon of volatility transmission has also been found in currency markets (e.g. Melvin and Melvin (2003), Huang and Yang (2002), Kearney and Patton (2000)) and in futures markets (e.g. Gannon and Choi (1998), Franses et al. (1997), Najand et al. (1992)). Stevenson (2002) further documents the spillover of returns volatility from equity REITs to other classes of REITs in the United States. However, only few studies have investigated the spillover of volatility across international real estate securities. This study therefore examines the international transmission of REIT returns volatility using REITs of Japan, United Kingdom, and United States.

By applying Generalized Autoregressive Conditional Heteroscedasticity (GARCH) and Exponential GARCH (EGARCH) models, we find that there are significant international spillovers of REIT return volatility among the three countries. The geographic risk factors of a country affect volatility of REIT returns of other countries significantly. We also find that the volatility spillovers are symmetric. The negative news in one market does not increase volatility more than positive news in another market. Overall, our findings suggest that investors could benefit from international diversification by investing in real estate securities from countries that have a lower integration between their property sector and the general stock market.

The reminder of the paper is organized as follows. Section 2 summarizes relative literature about volatility spillovers across national markets and develops research hypotheses. Section 3 describes data selection, research methodology, and empirical models. Section 4 presents analysis and interpretations of the empirical findings. Section 5 concludes the paper.

LITERATURE AND HYPOTHESES

Financial markets around the world have become increasingly integrated over the past decade. Literature has documented that asset prices exhibit substantial co-movements internationally since. One strand of literature focuses on whether there is a market contagion across national markets. For example, King and Wadhwani (1990) analyze cross-market correlations between U.S., U.K., and Japan stock markets and find contagion effects are increased significantly after market crash in 1987. By using intra-daily data of New Your (S&P500 index) and Japan (Nikkie 225 Index), Lin, Engle and Ito (1994) suggest that information revealed in U.S. or Japan market could have a significant impact on the returns of the other market. Hamao, Masulis and Ng (1990) measure the interdependence of return and volatilities across international stock markets and find spillover effects in the returns and volatilities from the U.S. and the U.K. stock markets to the Japanese market are very strong, but the volatility spillovers from Japan to both U.S. and U.K. are not significant.In recent years, several studies also examine volatility transmission in multiple global markets. Ng (2000) examine the volatility spillovers from Japan and the US to Pacific–Basin equity markets and find both Japan and U.S. have significant impacts on market volatility in the Pacific–Basin region, although the effect of U.S. market is greater. Bekaert and Harvey (1997) analyze volatility in 20 emerging markets and find that increased market integration often increases the correlation between local market returns and the world market returns. Baele (2005) investigates volatility transmission from the aggregate European (EU) and U.S. market to 13 European equity markets and find high EU and U.S. Stock spillover intensity increased significantly over the 1980s and 1990s, and significant contagion effects from the U.S. to a number of local European markets in times of high equity market volatility. Diebold and Yilmaz (2009) examine return and volatility spillovers for 19 global markets and find volatility spillovers display no trend but clear bursts.

Another strand of literature focuses on whether asymmetries exist in the volatility transmission process. Bae and Karolyi (1994) examine Nikkie 225 index and S&P500 index return volatility and find asymmetric effect of good and bad news in a market. That is, bad news increasing volatility in one market has greater impact on volatility in the other market than good news. Koutmos and Booth (1995) also find similar results when they examine return and volatility spillover between U.S., U.K. and Japan markets. With real estate security data, Michayluk et al. (2006) document that asymmetric volatility transmission exists between U.K. and the U.S. real estate markets.

Although the volatility spillover in equity markets has been studied intensively, few papers have studied spillover of volatility across international real estate securities. Therefore, in this study, we examine whether real estate securities are globally integrated using REITs of Japan, United Kingdom, and United States. There are two competing hypotheses on this issue. We call the first hypothesis the Transportable Risk Hypothesis which states that geographic risk is transportable across national borders. This hypothesis is supported by many studies (e.g. Liu, et al. (1990), Mei and Lee (1994), Li and Wang (1995), Ling and Naranjo (1999)) that show REITs are integrated with common equities. If REIT and stock markets are truly integrated and that stock returns volatility can transmit internationally, then it is logical to assume that REIT returns volatility should also transmit internationally. The other hypothesis is the Non-Transportable Risk Hypothesis which states property markets are less integrated globally than general equity markets. Several studies (e.g. Asabere, Kleiman and McGowan (1991), Hudson-Wilson and Stimpson (1996)) evidence that the U.S. property markets are less globally integrated than the U.S. equity markets due to low positive correlations between U.S. real estate investment and international real estate equities. Gordon and Canter (1999), in explaining why there are considerable diversification benefits from an international real estate securities portfolio, suggest that it is because the revenues of most real estate securities remained closely tied to individual countries. Thus, cross-border influences are still of relatively low significance. As such, their explanation suggests that real estate securities should have low international transmission of returns volatility.

DATA AND METHODLOGY

We collect daily returns of the three REIT indices of United States, United Kingdom and Japan between Jan 1, 1999 and 2006 from the National Association of Real Estate Investment Trusts of the United States (NAREIT, www.reit.com). To be included in an index, the firm must be a closed-end company listed on an official stock exchange. In addition, the firm must meet specific geographic and financial standards. These standards in general request that the majority of earnings or bulk of total assets is the result of relevant real estate activity. Relevant real estate activities include the ownership, trading and development of income producing real estate. The majority of the earnings must also be derived from domestic operations. Such a requirement ensures no significant cross-correlations of the cash flows of the REITS of different countries. The company must also meet a minimum requirement regarding market capitalization. To investigate REIT returns volatility transmission among the three markets, we use both the GARCH and EGARCH models. The GARCH specification was developed by Bollerslev (1987) from the basic Autoregressive Conditional Heteroscedasticity (ARCH) procedures of Engle (1982). Both procedures have been found to perform remarkably well in modeling financial time series. These models allow for a time-varying conditional variance and that the conditional variance is modeled as a function of its past values as well as independent and/or exogenous variables. Specifically, the following GARCH (1, 1) specification is used.

$$R_{i,t} = \beta_{i,0} + \beta_{i,1} R_{i,t-1} + \varepsilon_{i,t} \tag{1}$$

$$\sigma^2_{i,t} = \alpha_{i,0} + \sum_{j=1}^{3} \alpha_{i,j} \varepsilon^2_{j,t-1} + \gamma_i \sigma^2_{i,t-1} \quad for\ i, j = 1,2,3; \tag{2}$$

where $R_{i,t}$ is the daily REIT return series of country i, i=1,2,3 (i.e. 1 = United States, 2 = United Kingdom, and 3 = Japan) at time t, and $R_{i,t-1}$ is the daily REIT return series of country i, i=1,2,3 (i.e. 1 = United States, 2 = United Kingdom, and 3 = Japan) at time t-1. The time subscripts correspond directly to trading time but not necessarily to calendar time.

In the conditional variance equation, the variance of REIT returns depends on its own lagged value as well as the lagged squared residuals (innovations) of all the three countries. The lagged squared residuals in the equation are used for detecting volatility transmission across international boundaries. The EGARCH specification was developed by Nelson (1991). An advantage of EGARCH is that it is ideally suited to test the possibility of asymmetries in the volatility transmission mechanism because it allows own market and cross market innovations to exert an asymmetric impact on the volatility in a given market. In other words, news generated in one market is evaluated in terms of both size (i.e. the quantity) and sign (i.e. the quality) by other markets. Nelson (1991) finds that, for the US stock market, negative innovations increase volatility more than positive ones. Cheung and Ng (1992), Koutmos (1992), and Poon and Taylor (1992) all report evidence of the asymmetric impact of news shocks on volatility. The following EGARCH specification is also used in the study.

$$R_{i,t} = \beta_{i,0} + \beta_{i,1} R_{i,t-1} + \varepsilon_{i,t} \tag{3}$$

$$\sigma_{i,t}^2 = \exp\{\alpha_{i,0} + \sum_{j=1}^{3}\alpha_{i,j} f_j(z_{j,t-1}) + \gamma_i In(\sigma_{i,t-1}^2)\}, \ \ for\ i,j = 1,2,3; \tag{4}$$

$$f_j(z_{j,t-1}) = \{|z_{j,t-1}| - E(|z_{j,t-1}| + \delta_j z_{j,t-1}\}, for\ \ j = 1,2,3; \tag{5}$$

$$\sigma_{i,j,t} = \rho_{i,j}\sigma_{i,t}\sigma_{j,t,} \text{ for } i,j = 1,2,3, \text{ and } i \neq j. \tag{6}$$

Equation (4) stipulates that the conditional variance of a country's REIT returns can be affected by volatility spillovers from the other two countries. Volatility spillovers across markets are measured by α_{ij} for i,j = 1,2,3 and i≠j. A significant α_{ij} coupled with a negative δ_j implies that negative innovations in market j have a higher impact on the volatility of market i than positive innovations, i.e. the volatility spillover mechanism is asymmetric. The γ_i in equation 4 measures the persistence of volatility.

EMPIRICAL RESULTS

The descriptive statistics in Table 1 shows that the REIT returns of UK and Japan are positively skewed while the REIT returns of US is negatively skewed. The kurtosis values of all the three returns series are much larger than three. This indicates that all the REIT returns series are leptokurtic and have fat tails relative to the normal distribution. All the three REIT returns series exhibit significant departure from normality, as indicated by the Jarque-Bera statistics, which reject the null hypothesis of normal distribution at the 1% significance level.

Table 1: Summary Statistics of REIT Index Returns

Statistics	US	UK	JP
Mean(μ)	0.0741	0.0558	0.0442
Median	0.0736	0.0438	0.0122
SD(σ)	0.7642	0.9167	2.1658
Skewness	-0.0090	0.0824	0.4284
Kurtosis	5.9775	6.1257	4.3038
Min.	-3.4527	-4.2778	-5.9254
Max.	4.7621	5.4135	10.2171
Jarque-Bera	370.5084	409.4431	101.7230
Probability	0.0000	0.0000	0.0000
ADF test at the Level I(0)[a]	-14.7216***	-13.1416***	-14.4086***

ljung-box q test results

Q-Statistics	US	UK	JP
A: Ljung-Box Q test for Autocorrelation of Raw Returns			
Lag(4)	25.310***	21.773***	5.745
Lag(8)	35.100***	23.430***	6.916
Lag(12)	35.734***	24.142***	10.524
Lag(16)	38.673***	27.686***	14.145
Lag(20)	41.950***	33.865***	17.925
Lag(24)	49.542***	36.380***	19.868
B: Ljung-Box Q test for Autocorrelation of Squared Raw Returns			
Lag(4)	338.41***	42.437***	1.640
Lag(8)	430.49***	47.550***	15.717**
Lag(12)	457.93***	49.338***	30.097***
Lag(16)	461.96***	66.208***	35.768***
Lag(20)	463.39***	76.864***	41.270***
Lag(24)	464.59***	78.868***	43.389***

[a] *The ADF test is Augmented Dickey-Fuller Unit Root Test. The ADF test is a test of stationarity. The critical values for ADF test are –2.5685, -2.8649, and –3.4396 for significance level of 10%,5%, and 1% respectively. ***, **, and * indicate significance at the 1, 5, 10 percent levels respectively.*

In addition to the above test, the Augmented Dickey-Fuller (ADF) test was conducted to check for unit root (stationarity) in order to determine whether the REIT returns need to be transformed before model estimation.

The ADF statistics strongly indicated that all the REIT returns are stationary. We reject the null hypothesis of unit root at the 1% level. The Ljung-Box Q test statistics indicate that the squared raw REIT returns have substantially higher autocorrelations than the raw returns. This indicates the presence of strong conditional heteroscedasticity in REIT returns. In short, statistical properties of the data strongly support the usage of GARCH and EGARCH models in this study.

Table 2 reports results of the GARCH (1, 1) model which allows for volatility transmission. The US and Japan stock markets open and close sequentially, as do the Japan and UK markets. There is no overlap in the daily open-to-close returns between US and Japan. Between UK and US, there is approximately a two-hour overlap. To simplify the analysis we assume that all three markets open and close sequentially. Non-overlapping trading implies that the estimation of the variances in each market is conditional on its own past information as well as information generated by the last two markets to close. In the estimation equations, $R_{i,t}$ is the REIT index return at time t for market i, (i =1, 2, 3, where, 1=US, 2=UK, and 3=Japan). The time subscripts in equation 1-2 correspond directly to trading time but not necessarily to calendar time. For example, if i =Japan, the time subscripts in equation 1-2 correspond directly to trading time , and the time subscripts also correspond directly to calendar time. If i =Japan, the information set for traders in Japan at the opening of the market in a given day includes past Japan innovations as well as innovations from US and UK in the previous day. Here, the trading time and the calendar time are consistent. But if i =US, the information set for traders in US at the opening of the market in a given day includes past US innovations as well as innovations from UK and Japan during the same day. In terms of trading time all this is past information (information set at t-1). But in terms of calendar time, US, UK and Japan are in the same day.

Table 2: Multivariate GARCH Model with Volatility Spillovers

From New York($\alpha_{3,1}$) & London ($\alpha_{3,2}$) to Tokyo		From Tokyo($\alpha_{2,3}$) & New York($\alpha_{2,1}$) to London		From London($\alpha_{1,2}$) & Tokyo($\alpha_{1,3}$) to New York	
$\beta_{3,0}$	0.0181 (0.0694)	$\beta_{2,0}$	0.0651 (0.0253)***	$\beta_{1,0}$	0.0802 (0.0208)***
$\beta_{3,1}$	0.0669 (0.0336)**	$\beta_{2,1}$	0.1230 (0.0360)***	$\beta_{1,1}$	0.0959 (0.0350)***
$\alpha_{3,0}$	0.1793 (0.0789)**	$\alpha_{2,0}$	0.0375 (0.0169)**	$\alpha_{1,0}$	0.0605 (0.0198)***
$\alpha_{3,1}$	-0.0446 (0.0234)**	$\alpha_{2,1}$	0.0265 (0.0154)*	$\alpha_{1,1}$	0.1700 (0.0361)***
$\alpha_{3,2}$	0.0045 (0.0232)	$\alpha_{2,2}$	0.1005 (0.0209)***	$\alpha_{1,2}$	0.0229 (0.0108)**
$\alpha_{3,3}$	0.0210 (0.0097)**	$\alpha_{2,3}$	0.0046 (0.0019)***	$\alpha_{1,3}$	0.0011 (0.0018)
γ_3	0.9440 (0.0224)***	γ_2	0.8122 (0.0371)***	γ_1	0.6698 (0.0704)***

$R_{i,t} = \beta_{i,0} + \beta_{i,1} R_{i,t-1} + \varepsilon_{i,t}$

$\sigma_{i,t}^2 = \alpha_{i,0} + \sum_{j=1}^{3} \alpha_{i,j} \varepsilon_{j,t-1}^2 + \gamma_i \sigma_{i,t-1}^2$, for i,j=1,2,3

*Numbers in parentheses are standard error. ***, **, and *indicate significance at the 1, 5, 10 percent levels respectively.*

Results in Table 2 show that there are significant volatility spillovers from US to Japan (α_{31} =-0.0446) and UK (α_{21} = 0.0265) from Japan to UK (α_{23} = 0.0046) and from UK to US (α_{12} = 0.0229). That is, the findings support the Transportable Risk Hypothesis. Geographic risk factors of a country can affect volatility of REIT returns of other countries. Given that the REITs included in the NAREIT international indices are required to derive most of their earnings from domestic operations, it is therefore unlikely that the international transmission of REIT returns volatility is caused by correlated cash flows among the REITs. Hence, it appears that the reason for the volatility transmission is due to the integration of international REIT markets which makes possible the international transportation of geographic risk factors after they are securitized. Such an observation has major implications for investors who seek portfolio risk diversification by investing in international real estate securities. It appears that investors would get better diversification benefits by investing in countries that have a lower degree of integration between the general equity and real estate securities markets. Gordon and Canter (1999) find that some countries have a low correlation between the property securities and the broader equity market.

The persistence of volatility implied by equation 2 is measured by γ_1, γ_2 and γ_3. They are all significantly less than one, a result that is necessary for the unconditional variance to be finite. Persistence is strongest in Japan and least in US. This can be interpreted by using half-life concept, which measures the time it takes a shock to reduce its impact by one half. For UK the half-life is 3.33 days, for US half-life is 1.73 days, and for Japan half-life is 12.03 days. (Half-life for market i equals $\ln(0.5)/\ln \gamma_i$.)

Table 3: Diagnostic Statistics for the Standardized Residuals for the GARCH (1, 1) Spillover Model

Panel A: Diagnostic Tests			
	US	**UK**	**JP**
Skewness	-0.2337	0.0397	0.3733
Kurtosis	4.4499	5.0964	4.0013
Jarque-Bera	96.88	183.57	65.06
Probability	0.0000***	0.0000***	0.0000***
ARCH test p-value	0.6145	0.2939	0.2421
Panel B: Autocorrelation Q-statistics for Standardized Residuals			
Lag			
4	2.9736	4.2433	1.5175
8	6.5966	6.3539	3.0201
12	9.8177	6.3698	4.9647
16	12.430	9.2318	8.2920
20	16.494	15.012	11.704
24	25.599	17.727	13.076
Panel C: Autocorrelation Q-statistics for Standardized Residuals Squared			
Lag			
4	3.5972	1.1198	1.1621
8	4.7640	2.0885	6.4615
12	8.2988	4.9456	17.917
16	9.6303	11.010	19.335
20	16.238	21.494	21.800
24	20.774	27.291	27.623

****, **, and *indicate significance at the 1, 5, 10 percent levels respectively.*

Tablet 3 provides results of several diagnostic tests of the standardized residuals obtained from the multivariate GARCH model. The standardized residuals of UK and Japan are positively skewed, but the standardized residuals of US are negatively skewed. The kurtosis values are significantly larger than three for all countries, which indicate that the standardized residuals are leptokurtic and have fat tails relative to the normal distribution. Consequently, the Jarque-Bera normality tests reject the null hypothesis of normally distributed standardized residuals in all three countries. On the other hand, the ARCH-LM tests do not indicate the presence of a significant ARCH effect in all the three REIT returns series. This result suggests that the standardized residuals have constant variances and do not exhibit serial correlation. Panel B and panel C display the Ljung-Box Q statistics for the standardized residuals and squared standardized residuals at the 4^{th}, 8^{th}, 12^{th}, 16^{th}, 20^{th}, 24^{th} day lags. All the Q-statistics are insignificant. There is no autocorrelation in the standardized residuals and squared standardized residuals. The GARCH model therefore appears to provide a good parameterization of the three REIT returns series.

Table 4: Multivariate EGARCH Model with Volatility Spillovers

From New York($\alpha_{3,1}$) & London ($\alpha_{3,2}$) to Tokyo		From Tokyo($\alpha_{2,3}$) & New York($\alpha_{2,1}$) to London		From London($\alpha_{1,2}$) & Tokyo($\alpha_{1,3}$) to New York	
$\beta_{3,0}$	0.0432 (0.0681)	$\beta_{2,0}$	0.0614 (0.0266)**	$\beta_{1,0}$	0.0672 (0.0199)***
$\beta_{3,3}$	0.0533 (0.0312)*	$\beta_{2,2}$	0.1091 (0.0372)***	$\beta_{1,1}$	0.0848 (0.0370)**
$\alpha_{3,0}$	0.9966 (0.6169)*	$\alpha_{2,0}$	-0.2513 (0.0561)***	$\alpha_{1,0}$	-0.3774 (0.0669)***
$\alpha_{3,1}$	-0.0324 (0.0369)	$\alpha_{2,1}$	0.0238 (0.0148)*	$\alpha_{1,1}$	0.3055 (0.0584)***
$\alpha_{3,2}$	0.0082 (0.0387)	$\alpha_{2,2}$	0.2401 (0.0607)***	$\alpha_{1,2}$	0.0302 (0.0119)***
$\alpha_{3,3}$	-0.0088 (-0.1095)	$\alpha_{2,3}$	0.0055 (0.0020)**	$\alpha_{1,3}$	0.0027 (0.0040)
δ_3	0.0896 (0.0505)	δ_2	-0.0432 (0.0434)	δ_1	-0.0528 (0.0459)
γ_3	0.3621 (0.3930)	γ_2	0.8922 (0.0357)***	γ_1	0.8656 (0.0373)***

$R_{i,t} = \beta_{i,0} + \beta_{i,1} R_{i,t-1} + \varepsilon_{i,t}$

$\sigma_{i,t}^2 = \exp\left\{\alpha_{i,0} + \sum_{j=1}^{3} \alpha_{i,j} f_j(z_{j,t-1}) + \gamma_i In(\sigma_{i,t-1}^2)\right\}$, for $i,j = 1,2,3$;

$f_j(z_{j,t-1}) = \{|z_{j,t-1}| - E(|z_{j,t-1}| + \delta_j z_{j,t-1}\}, for\ j = 1,2,3;$

$\sigma_{i,j,t} = \rho_{i,j}\sigma_{i,t}\sigma_{j,t,}\ for\ i,j = 1,2,3, and\ i \neq j$

****, **, and *indicate significance at the 1, 5, 10 percent levels respectively.*

In Table 4, we report results of the EGARCH specification. Similar to the results of the GARCH model, we find strong evidence of international transmission of volatility of REIT returns. Specifically, there are significant volatility spillovers from US to UK (α_{21} = 0.0238), from Japan to UK (α_{23} = 0.0055), and from UK to US (α_{12} = 0.0302). The results are similar to those reported in Table 2 that there are significant volatility spillovers across the Atlantic, however, the volatility spillover from the Pacific only affects UK and does not reach US. The results again support the Transportable Risk Hypothesis. Moreover, the volatility spillovers are symmetric since the coefficients measuring asymmetry δ_3, δ_2 and δ_1 are not significant for all three markets. That is, negative news (innovations) in one market does not increase volatility more than positive news in another market. Asymmetry in volatility transmission is modeled by equation (5) and can be examined using its derivatives:

$$\begin{aligned} \partial f_j(z_{j,t}) / \partial z_{j,t} &= 1+\delta_{j,} \text{ for } z_j > 0 \\ &= -1+\delta_{j,}, \text{ for } z_j < 0 \end{aligned} \qquad (7)$$

The size effect is measured by $|z_{j,t-1}| - E(|z_{j,t-1}|)$, and the corresponding sign effect is given by $\delta_j z_{j,t}$. If δ_j is negative, a negative $z_{j,t}$ tend to reinforce the size effect. If δ_j is positive, a negative $z_{j,t}$ tend to mitigate the size effect. The relative magnitude of asymmetry may be quantified by comparing the right-hand side of equation (7) when Z_j <0 and Z_j >0. It is found that conditional volatility of the REIT returns of UK and US also respond symmetrically to their own past innovations $\alpha_{1,1}$ and $\alpha_{2,2}$. The persistence of volatility implied by equation 4 is measured by γ_1, γ_2 and γ_3. They are all significantly less than one, a result that is necessary for the unconditional variance to be finite. Persistence is strongest in UK and least in Japan.

Table 5: Diagnostic statistics for the standardized residuals for the EGARCH (1, 1) Spillover Mode

Panel A: Diagnostic Tests			
	US	**UK**	**JP**
Skewness	-0.1985	0.0868	0.3611
Kurtosis	4.5422	5.0826	4.0505
Jarque-Bera	105.89	182.16	67.78
Probability	0.0000***	0.0000***	0.0000***
ARCH test p-value	0.4376	0.1494	0.0072***
Panel B: Autocorrelation Q-statistics for Standardized Residuals			
Lag			
4	3.1307	4.1611	2.1729
8	6.7404	6.1032	3.7364
12	9.3862	6.1143	7.2943
16	12.137	9.2288	11.122
20	16.276	14.991	14.268
24	26.792	17.703	15.796
Panel C: Autocorrelation Q-statistics for Standardized Residuals Squared			
Lag			
4	3.6778	1.8471	1.5068
8	4.9819	2.5334	18.481**
12	8.7433	5.1594	31.008**
16	9.8977	15.011	36.045***
20	18.690	25.335	39.728***
24	23.624	30.978	42.665**

****, **, and * indicate significance at the 1, 5, 10 percent levels respectively.*

Table 5 provides several diagnostic tests results of the standardized residuals obtained from the multivariate EGARCH model. Compared with multivariate GARCH model estimation, the results are similar except for Japan. The ARCH-LM test statistic is significant for Japan, indicating the presence of a significant ARCH effect for Japan. Some Q-statistics for squared standardized residuals are also significant for Japan, which suggest the presence of heteroscedasticity. Thus, we reject the null hypothesis of no autocorrelation in the squared standardized residuals for Japan. That is, the GARCH specification presents a better parameterization of REIT returns series for Japan than the EGARCH specification.

CONCLUSION

The purpose of this study is to examine whether volatility of REIT returns can transmit across national markets. More specifically, we examine the international transmission of REIT returns volatility among United States, United Kingdom, and Japan by using daily returns of the three REIT indices from NAREIT. Given the statistical properties of the time series data of the three REIT returns, GARCH (1, 1) and EGARCH (1, 1) models are used in the study to test volatility spillovers.

Our findings are summarized as follows. First, we find that the EGARCH model works better for US and UK REITs whereas the GARCH model is better for Japanese REITs. Second, the empirical results show that there are significant volatility spillovers from US to Japan and UK, from Japan to UK, and from UK to US. It indicates that significant international spillovers of REIT returns volatility exist among the three countries. This finding supports Transportable Risk Hypothesis, which indicates that geographic risk factors of a country can affect volatility of REIT returns of other countries. Third, the result of the

EGARCH model shows that the REIT returns volatility transmission is not asymmetric. That is, negative news does not affect volatility more than positive news.

Our findings are consistent with the implication that the real estate sector and the general equity market are integrated in the three countries. An important implication for investors is that they would get more portfolio risk diversification benefits by investing in real estate securities from countries that have a lower integration between their property sector and the general stock market. The volatility transmission literature has yet to develop a comprehensive theory to explain the observed results of empirical studies. Nevertheless, the results of our study show that a general understanding of the relationship between international real estate securities markets movements has a practical use for formulating international portfolio strategies as it improves forecasting ability.

Even though our results show significant volatility spillovers among those markets, there are some limitations in our study. First, our results are based on the evidence of three major REIT markets from developed countries. Second, readers need to be aware of that the volatility effect we present in this paper could be time-varying and such relation could be changed due to regime switch or market shock. Future studies might consider including other REITs, particularly REITs from emerging markets to verify whether our findings hold.

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CAPITAL STRUCTURE TIMING IN MARKETS WITH DIFFERENT CHARACTERISTICS

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ABSTRACT

Considerable empirical evidence suggests that firm's time equity issues to market movements and that this behavior impacts capital structures. Based on a survey of investigations of this phenomenon, this study observes capital structures in different financial markets and identifies different situations related to the effect of timing on leverage. This study also explains optimal leverage with a simplified dynamic adjusted model. Firms facing financial constraints in debt financing may increase equity issues resulting in considerable leverage variance. On the other hand, firms with fewer financial constraints can time the market when issuing equity. This study takes regional samples from the United Kingdom and Japan, to summarize circumstances involving partial financial constraints and no financial constraints. The market timing effects tests in the United Kingdom are insignificant but the results for Japan are significant. This phenomenon improves understanding of the market timing model under different circumstances.

JEL: G30; G31

KEYWORDS: Market Timing, Capital Structure.

INTRODUCTION

International capital structure is seldom examined, perhaps because of data limitations and insufficient methods for comparing different markets. This study investigates differences in financial market leverage and the impact of market timing on capital structure in markets with different characteristics. Pecking order and tradeoff theories are applied at static points and may lead to misleading leverage information. Consequently, some studies investigate capital structure also across different time periods. Opler and Titman (2001) stated that financial decisions include an optimal target debt ratio. Hovakimian and Titman (2001) used two stage regressions to conclude that firms adjust to an optimal target debt ratio that may change over time and be related to profitability and stock price. Baker and Wurgler (2002) used weighted market-to-book ratios as a proxy for the past impact of equity issuance on capital structure and declared that the market-to-book effect exerts a persistent and long lasting influence on capital structure.

The market timing effect on capital structures under different market characteristics reveals countries with similar characteristics but different financing patterns. This phenomenon can supply data for investigating whether weighted market-to-book ratios represent a good proxy for deviation between current and target debt ratios (Baker and Wurgler, 2002). The United Kingdon (UK) and Japan are chosen for examination in this study. Both countries belong to the G-7 and have similar market capitalization percentages, at around 80% (Rajan and Zingales, 1995). Firms from these countries have more financing via banks, meaning external financing is not fully reflected in their capital structure, or perhaps leverage can be adjusted to reach the target. This study proposes a simplified dynamic adjustment model, similar to that of Banrjee et al. (2000) to explain how leverage effects vary among markets.

The remainder of this paper is organizes as follows. The following section reviews the relevant literature. Next, the theory is presented. A discussion of the data and methodology and presentation of test results follows. The paper closes with some concluding comments.

LITERATURE REVIEW AND BACKGROUND

Modigliani and Miller (1958) showed that in an idealized world without taxes, firm value is independent of the debt-equity mix. In short capital structure is irrelevant to firm value. Other researchers, such as Hamada (1969) and Stiglitz (1974), support the perspective of Modigliani and Miller. However, these conclusions do not match observations of the real world, in which capital structure matters and banks are extremely unwilling to finance projects entirely using debt capital. The main theories explaining capital structure are the pecking order and tradeoff theories. Extensive empirical tests have been completed in relation to these theories, but their robustness and the situations in which they can be applied remain unclear. Smith and Watts (1992) obtained contradictory evidence from testing implication of the pecking order theory, finding that high-growth firms with high financial needs also have high debt ratio.

The most widespread theories explaining capital structure are pecking order theory and tradeoff theory. According to pecking order theory, information asymmetry makes equity financing more expensive than debt financing (Myers and Majluf, 1958).Opler and Titman (2001) identified optimal target debt ratios in market timing. Hovakimian and Titman (2001) used two stage regressions to conclude that firms adjust to an optimal target debt ratio that changes over time and is related to profitability and stock price. Korajczyk and Levy (2002) examined the effect of macroeconomic conditions on debt or equity choice. Furthermore, Baker and Wurgler (2002) used weighted market-to-book ratios as a proxy for the past impact of equity issuance on capital structure and identified the effect of market-to-book on capital structure as persistent and long-lasting. Frank and Goyal (2003) tested the pecking order theory for different firm size. The pecking order and tradeoff theories are only applied at static points and may result in misleading leverage information. Consequently, some recent studies investigate capital structure not only for certain static points, but also across different periods using panel data.

Urbonavičius *et al.* (2006) concentrated on measuring of company's marketing orientation and its relationship with manager-related factors. Specifically, Urbonavičius *et al.* analyzed such factors as personality traits; manager work needs motivation, leadership style, conflict-solving style and source of power. Analysis is performed using interdisciplinary methodology that permits broad discussion of firm market orientation and factors that influence it. The findings suggest that companies with similarly high market orientation are also similar in management-related factors, namely these factors exist in a favorable combination that closely matches high market orientation. Empirical data is obtained by surveying the management of Lithuanian furniture companies, and reflects the specific circumstances of transitional economies.

Henderson *et al.* (2006) presented a sample of firms raising approximately $25.3 trillion of new capital, including $4.9 trillion from overseas during 1990-2001. International debt issuances are more common than equity issuances, accounting for 87% of all securities issued internationally, and approximately 20% of all public debt issuances. In contrast, international equity issues account for approximately 9% of all international security issues, and 12% of all equity issues during the sample period. Market timing considerations appear very important in security issuance decisions. International firms are more likely to issue equity before periods of low market returns. Most cross-border equity is issued in the U.S. and the U.K., and these issues tend to occur in 'hot' markets and before several periods of relatively low market returns. Finally, firms issue more debt when interest rates are lower, and before they increase.

Firms may borrow money or issue equity to raise funds. Debt and equity are the two key items of external financing in the capital structure. In an extreme example, if a firm faces large financial constraints restricting its ability to borrow, that firm may issue as much equity as it can, even to the point of exceeding its immediate future needs. External financing is consistent with tradeoff theory, but not with pecking order theory that debt is always preferred to equity. Conversely, a firm facing no financial constraints can optimize its capital structure. A firm may issue equity only when the market timing is favorable. That is, firms raise capital through markets depending on the relative cost of debt and equity, consistent with the tradeoff theory. Debt is much more expensive than equity for firms with constrained access to debt financing. These firms issue equity more frequently, potentially causing larger variance of leverage ratio. Firms with fewer financial constraints can obtain debt financing more easily, and may have higher debt ratio and can time the market (until the costs of issuing equity become advantageous to raise equity on the most advantageous terms). Therefore, capital markets strongly affect capital structure.

According to the tradeoff theory, the cost of debt financing versus equity financing is the most important determinant of capital structure. This study examines the financial constraints that can influence the costs of equity and debt. This study defines a firm as facing financial constraints when the sum of the market value of equity and the book value of debt, divided by total assets, exceeds one. A ratio exceeding one means the firm depends heavily on financial markets, and raises as much finance as possible. Two main sources determine the ratio, equity and debt. Firms facing large constraints in financial markets face higher borrowing costs regardless of their internal financial condition, and may issue as much equity as they can, even exceeding their likely near future needs. However, while such firms are heavily dependent on financial markets, they do not necessary have high leverage ratio. Consider the following example: a newly established enterprise may find itself unable to borrow sufficiently to meet its needs, and thus may issue equity. For such growth firms, equity represents an important means of financing. On the other hand, firms facing fewer conflicts in financial markets can time the market in issuing equity. Such firms may extensively use debt leverage, because of their ability to acquire funds on financial markets.

THEORY

Before testing the experimental model of market timing, we discuss the theory. Baker and Wurgler (2002) declared that firms issue equity in a way that times the market. Firm leverage ratio thus does not fully reflect the ideal debt ratio. Firms can gradually optimize their leverage. Although the model does not assume optimal leverage, it declares that the gap between real and target debt ratios is largely explainable by the weighted market-to-book, *wmb,* ratio. The model provides extensive empirical testing and comparison of the coefficients of the following three equations. The model is as follows:

$$\left(\frac{D}{A}\right)_{t+1} = a_1 + b_1\left(\frac{M}{B}\right)_{efwa,t} + c_1\left(\frac{M}{B}\right)_t + d_1\left(\frac{PPE}{A}\right)_t + e_1\left(\frac{EBITDA}{A}\right)_t + f_1\log(S)_t + u_{1,t+1} \quad (1)$$

$$\left(\frac{D}{A}\right)_{t+\tau} = a_2 + b_2\left(\frac{M}{B}\right)_{efwa,t} + c_2\left(\frac{M}{B}\right)_t + d_2\left(\frac{PPE}{A}\right)_t + e_2\left(\frac{EBITDA}{A}\right)_t + f_2\log(S)_t + u_{2,t+1} \quad (2)$$

$$\left(\frac{D}{A}\right)_{t+\tau} = a_3 + b_3\left(\frac{M}{B}\right)_{efwa,t} + c_3\left(\frac{M}{B}\right)_{t+\tau-1} + d_3\left(\frac{PPE}{A}\right)_{t+\tau-1} + e_3\left(\frac{EBITDA}{A}\right)_{t+\tau-1} + f_3\log(S)_{t+\tau-1} + u_{3,t+\tau} \quad (3)$$

The first equation investigates the *wmb* effect for the current state. We let *wmb*substitute for $\left(\frac{M}{B}\right)_{efwa,t}$.

$\left(\frac{M}{B}\right)_{efwa,t-1} = \sum_{s=0}^{t-1}\frac{e_s + d_s}{\sum_{r=0}^{t-1} e_r + d_r} \bullet \left(\frac{M}{B}\right)_s$). Weighted market-to-book helps explain the unrebalanced target debt ratio. The second equation replaces leverage at time t with (t+ τ) leverage. The third equation uses (t+ τ -1) as an explaining variable to control the future condition. If leverage is adjusted to the target level, past *wmb* no longer impacts leverage, and thus the b2 and b3 coefficients should be close to zero. If this is not the case, *wmb* still has power to explain future leverage, and the b2 and b3 reflect the influence strength.
Capital structure theory comprises two main strands, based on static and dynamic models. In static

models, the pecking order and tradeoff theory are well known. Empirical models also exist for identifying the determinants of an optimal debt ratio. The dynamic models assume an optimal debt ratio (Fischer, Heinkel, and Zechner, 1989), but this may be achieved within j periods. (Hovakimian, Opler, Titman, 2001 and Banrjee *et el.*, 2000). This study discusses the market timing model of Baker and Wurgler (2002) from the perspective of dynamic optimal capital structure.

This study uses the following algorithm. Similar to the dynamic adjustment model, the debt ratio can be optimized over j periods. Market-to-book ratios can largely explain the difference between optimal and true value (Baker and Wurgler, 2002). Hence, this study obtains the optimal debt leverage at time t.

$$\left(\frac{D}{A}\right)_{t+1} - \left(\frac{D}{A}\right)_{t} = d_{t+1} - d_t + v_{t+1} = c_t \times (M/B)_t + v_{t+1}, \qquad d_t = \left(\frac{D}{A}\right)_t, \quad v_t \sim IID\ (0, \sigma_v^2)$$

$$\left(\frac{\hat{D}}{A}\right)_{t+1} = X_t\beta + \sum_{n=0}^{j-1}(d_{t+1-n} - d_{t-n}) + u_{t+1} = X_t\beta + p_1 \sum_{n=0}^{j-1} W_{t-n} \times (M/B)_{t-n} + u_{t+1} = p_1\ wmb_t + X_t\beta + u_{t+1} \tag{4}$$

Equation (4) resembles Eqn. (1) of Baker and Wurgler (2002), but differs in that the model emphasizes that the *wmb* effect decreases over time without the assumption of optimal leverage. If the optimal leverage is achieved at time t, and can be estimated using factors X_{t-1}. Optimal leverage estimation is reduced to Eq. (5). This equation closely resembles Eq. 3.

$$\left(\frac{\hat{D}}{A}\right)_{t+\tau+1} = X_{t+\tau}\beta + \sum_{n=0}^{j-1}(d_{t+1-n} - d_{t-n}) + u_{t+\tau+1} = X_{t+\tau}\beta + p_3 \sum_{n=0}^{j-1} W_{t-n} \times (M/B)_{t-n} + u_{t+\tau+1} = p_3\ wmb_t + X_{t+\tau}\beta + u_{t+\tau+1} \tag{5}$$

If leverage is optimized at time t, and leverage holds at the steady state, then Eq. 6 is obtained, which resembles Eq. (2).

$$\left(\frac{\hat{D}}{A}\right)_{t+\tau+1} = \left(\frac{\hat{D}}{A}\right)_{t+1} = X_t\beta + \sum_{n=0}^{j-1}(d_{t+1-n} - d_{t-n}) + u_{t+1} = X_t\beta + p_2 \sum_{n=0}^{j-1} W_{t-n} \times (M/B)_{t-n} + u_{t+1} = p_2\ wmb_t + X_t\beta + u_{t+1} \tag{6}$$

Based on the dynamic adjustment model and empirical evidence of Baker *et al.* (2002), this study obtains Eq. (4) (6) (5), which resemble Eqs. (1) (2) (3). However, the two equation sets have different basic assumptions and testing purposes. The *wmb* coefficient ratios of (6) to (4), p2/p1, and (5) to (4), p3/p1 resemble the implication of *b2/b1*, *b3/b1*. Given an optimal capital structure, the coefficient of *wmb* indicates the leverage adjustment. The influence of *wmb* decreases over time, as do p2/p1 and p3/p1. The change in the ratio reveals whether *wmb* exerts a persisting effect.

This study compares the adjustments for j periods to weighted time periods. In Baker (2002), periods start from either the start of the data or the IPO time, meaning j starts somewhere between time 0 and t. For financially unconstrained firms, *wmb,* is an important factor in adjusting debt ratio. If firms rapidly adjust their debt ratio, leverage is optimized within i periods; i < j, and *wmb* number reveals (j-i) surplus weighted market-to-book ratios. Coefficients of *wmb* are smaller than the values at optimal adjustment. The *wmb* variable disturbs the debt leverage explanation because of excessive numbers of explanatory variables. In this condition where leverage is assumed to be optimized, p2/p1 and p3/p1 are ambiguous.

If the adjustment is still ongoing, leverage remains sub-optimal at time t. That is, the firm takes a long time to optimize leverage; i periods, i > j, and the optimal leverage requires future explanatory variables. There are (i-j) weighted market-to-book ratios not in the *wmb* number. However, further explanation is required for the optimal leverage at time t in Eq. (4). Equation (6) uses the same explanatory variables as Eq. (4), but explains leverage at time $t + \tau$. Coefficients are insignificant in *wmb* because of insufficient explanatory variables. Comparisons of the *wmb* effect between Eq. (4) to (6) are dubious owing to the assumption of optimal leverage and a long adjustment period.

If τ > (i-j), the leverage is optimized, although *wmb* lacks an adequate adjustment period, improving the explanatory power in Eqn. (5) over time. The same phenomenon may also explain why the *wmb* factor increases over time in Eq. (3) of Baker *et al.* (2002). This phenomenon also re-emphasizes the importance of the market timing test proposed by Baker *et al.* (2002). The test is dynamic, and leverage is not optimized as expressed in the assumption of this study, given optimal leverage and an adjustment time of j, the *wmb* coefficient is meaningful and the comparison of , p2/p1 and p3/p1 is identical to the market timing model of Baker and Wurgler (2002). This study thus proposes that the dynamic adjustment assumption model is simply a special case of the general market timing model (Baker and Wurgler, 2002).

DATA AND METHODOLOGY

Firm behavior varies by market. This study focuses on financial market characteristics that impact capital structure. To facilitate comparison, this study provides an averaged aggregate firm sample representing the overall market. Concentrating on financial constraints, this study roughly divides firms into financially constrained and unconstrained categories using aggregate financial market data. Financially constrained firms are defined as those with a ratio of aggregate market value of equity plus book value of debt, divided by total assets, exceeding one, similar to Tobin's Q.

The evidence suggests that capital structure is affected by different firm characteristics, methods of external financing and different time periods. This study investigates leverage in different types of financial markets using aggregate data. To compare capital structure among different markets, this study collects data from a worldwide database, DataStream. National data for the UK and Japan are collected to provide a sample from outside the US. This study collects data for all firms listed on DataStream from 1985 to 2001. It investigates firm leverage and tests the effect of market timing on capital structure for markets with different characteristics. Similar to Rajan and Zingales (1995), this study begins from a partial balance sheet of averaged annual firm data. Following the approach of Baker and Wurgler (2002), which used U.S. COMPUSTAT data, this study uses DataStream data which closely approximates COMPUSTAT's accounting definitions, thereby allowing a meaningful comparison.

Book equity, *BE,* is measured as Total Asset - Total Liabilities - Preference capital + Total Deferred Taxes + Convertible debt. Book Leverage is the percentage of Book Debt to Total Asset. We drop firms with Book Leverage above one from the sample because of the extreme value. Market Leverage, is the percentage of Book Debt. In Table 1, Japan's Book Leverage% is higher than that of UK, which is consistent with the G7 Balance Sheets result (Rajan and Zingales, 1995). The Market to Book ratio, *MB*, is defined as Total Assets minus Book Equity plus Market equity all divided by Total Assets. A higher MB ratio represents a higher growth firm.

Table 1 lists averaged balance sheet items for UK firms from 1985 to 2001. The table averages firms meeting certain requirements involving fixed assets (DataStream Item 339), total assets (Item 392), total capital employed (Item 322),total current liabilities (Item 389), convertible loans (Item 320), total deferred taxes (Item 312), preferred stock (Item 306), common equity (Item 305), total stockholder equity (Item 307), and ordinary dividend-net (Item 187). Item 628 is substituted for item 187 when the latter is unavailable. The table reports balance sheet data for UK firms from 1985 to 2001. The figure in each cell is the individual item divided by total assets and averaged across firms reported on DataStream during the year. The table lists data for odd numbered years to reveal trends.

Table 1: Balance Sheet Format Report for Kingdom Firms

Year	**1985**	**1987**	**1989**	**1991**	**1993**	**1995**	**1997**	**1999**	**2001**
Number of Observations	263	369	282	364	608	718	915	1,057	1,303
Assets									
current assets									
+Fixed assets(#339)	0.4567	0.4489	0.4738	0.4125	0.4617	0.4593	0.4599	0.4818	0.3729
Total assets(#392)	1	1	1	1	1	1	1	1	1
Liabilities:									
+Total capital employed(#322)	0.4759	0.4529	0.4691	0.5011	0.5030	0.4949	0.4871	0.4770	0.4907
+Total current liabilities(#389)	0.2468	0.2395	0.2264	0.2175	0.2057	0.2167	0.2253	0.2367	0.1924
=liabilities-total	0.7228	0.6924	0.6995	0.7186	0.7088	0.7117	0.7124	0.7137	0.6831
*convertible debt, convertible loans(#320)	0.0036	0.0045	0.0029	0.0028	0.0037	0.0041	0.0027	0.0039	0.0031
*Total Deferred Taxes(#312)	0.0117	0.0076	0.0077	0.0109	0.0067	0.0056	0.0050	0.0079	0.0065
+Preferred stock(#306)	0.0028	0.0047	0.0084	0.0076	0.0057	0.0051	0.0058	0.0063	0.0030
+Common equity(#305)	0.2743	0.3027	0.2959	0.2736	0.2853	0.2830	0.2817	0.2798	0.3137
=Total Stockholders' equity(#307=#305+#30x)	0.2771	0.3075	0.3044	0.2813	0.2911	0.2882	0.2875	0.2862	0.3168
=Total liabilities& Stockholders' equity	40,829	418,816	525,890	968,576	1,064,347	1,076,081	955,145	950,015	1,285,411
Ordinary dividend-net(#187, or #628)	0.0141	0.0178	0.0186	0.0191	0.0199	0.0231	0.0289	0.0254	0.0163

This table shows balance sheet data for UK Firms

Table 2 lists averaged Balance Sheet items for Japanese firms from 1985 to 2001. The table lists firms meeting certain requirements involving fixed assets (Item 339),total assets (Item 392),total capital employed (Item 322),total current liabilities (Item 389),convertible loans (Item 320),total deferred taxes(Item 312),preferred stock(Item 306),common equity (Item 305),total stockholder equity (Item 307), and ordinary dividend-net (Item 187,if unavailable, item 628 is selected instead). The table lists a balance sheet format report for Japanese firms from 1985 to 2001. The figure in each cell is the individual item divided by total assets and averaged across all firms included in the DataStream for that year. Table 2 only lists data for odd number years to make it easier to reveal trends more clearly.

Table 2: Balance Sheet Format Report for Japanese firms

Year	**1985**	**1987**	**1989**	**1991**	**1993**	**1995**	**1997**	**1999**	**2001**
Number of Observations	723	818	1236	1302	1326	1372	1548	1710	1859
Assets									
current assets									
+Fixed assets(#339)	0.2508	0.2620	0.2486	0.2598	0.2918	0.3004	0.3019	0.3188	0.3177
Total assets(#392)	1	1	1	1	1	1	1	1	1
Liabilities:									
+Total capital employed(#322)	0.3917	0.4308	0.4249	0.4268	0.4433	0.4625	0.4525	0.4766	0.4632
+Total current liabilities(#389)	0.4271	0.3756	0.3700	0.3681	0.3488	0.3282	0.3337	0.3055	0.2680
=liabilities-total	0.8188	0.8065	0.7950	0.7949	0.7922	0.7908	0.7863	0.7821	0.7313
*convertible debt, convertible loans(#320)	0.0087	0.0141	0.0177	0.0151	0.0155	0.0177	0.0157	0.0117	0.0051
*Total Deferred Taxes(#312)	0.0002	0.0002	0.0001	0.0003	0.0003	0.0013	0.0012	0.0003	0.0001
+Preferred stock(#306)	0	0	0	0	0	0	0	0	0
+Common equity(#305)	0.1811	0.1934	0.2049	0.2050	0.2077	0.2091	0.2136	0.2178	0.2686
=Total Stockholders' equity(#307=#305+#30x)	0.1811	0.1934	0.2049	0.2050	0.2077	0.2091	0.2136	0.2178	0.2686
=Total liabilities& Stockholders' equity	1	1	1	1	1	1	1	1	1
* Ordinary dividend-net (#187, or #628)	0.0046	0.0045	#N/A	#N/A	0.00390	#N/A	0.0035	0.0034	0.0035

This table shows balance sheet data for UK Firms

Comparison of different financial market characteristics, this study gathered international firm data from

1985 to 2001. As Tables 1 and 2 show, aggregated leverage during the year is defined as the average of total liabilities divided by total liabilities and shareholder equity. The leverage for Japanese firms is near 0.8, and decreases slightly from 1985 to 2001. Furthermore, the average ratio of current liabilities to total liabilities and shareholder equity ranges between 0.26 and 0.42 for Japanese firms. UK firms displayed lower aggregated average leverage than Japanese firms, at around 0.7, and current liabilities to total liabilities and shareholder equity for UK firms ranged between 0.19 and 0.24. Japanese firms thus have current and total leverage ratios nearly 10 % higher than UK firms.

The ratio of stockholder equity to total liabilities and shareholder equity for Japanese firms lies between 0.18 and 0.27. The same ratio for UK firms is between 0.28 and 0.33. Although liability leverage is lower for UK firms, the equity ratio is 5% to 10% higher than in Japan. The ratio of fixed assets to total assets for UK firm's ranges from 0.37 to 0.46 and ranges from 0.24 to 0.31 for Japanese firms. This indicates that the current ratio to total assets is higher for Japanese firms. On average, Japanese firms appear to adopt more relaxed asset investment policies than UK firms. Systematic differences also exist in averaged financial ratios between these two markets. Using the definition of financially constrained and financially unconstrained firms used in this study, (ME+BD)/TA ratio, market value of equity and book value of debt over total assets are added, and then averaged by the total number of firms in the market during the year. This yields a time series of aggregated average (ME+BD)/TA ratio, as listed in Table 2.

Table 3 lists the average (ME+BD)/TA ratio of the UK and Japan. The table compares data for the years (1985-2001). The figure in panel A lists the results for the sample of UK firms. Meanwhile, panel B lists the computational results for Japanese firms. Numerous factors may influence the target ratio, including firm size, industry category, etc. However, this study averages whole market data that can reduce fluctuations arising from other factors. This approach is mainly focused on investigating the evolution of the target ratio and comparing it with the financial characteristics of different markets. The ratios in the table are averaged by total firm number in the same year for both countries.

Table 3: The Ratio of Market Value of Equity Plus Book Value of Debt Divided by Total Assets

Panel A UK Firms			**Panel B: Japanese Firms**		
Year	**Number**	**(ME+BD)/TA**	**Year**	**Number**	**(ME+BD)/TA**
1985	263	0.949	1985	719	0.7454
1986	346	1.2654	1986	765	1.0113
1987	369	1.2702	1987	813	1.1356
1988	347	1.2043	1988	1166	1.3742
1989	282	1.2394	1989	1229	1.5086
1990	209	1.0553	1990	1272	0.9385
1991	364	1.0358	1991	1296	0.8765
1992	577	1.2399	1992	1310	0.7328
1993	608	1.3916	1993	1321	0.7872
1994	663	1.2977	1994	1344	0.8563
1995	718	1.3676	1995	1361	0.8540
1996	810	1.4579	1996	1400	0.8345
1997	915	1.5870	1997	1535	0.7506
1998	1005	2.5033	1998	1612	0.7246
1999	1057	2.0673	1999	1688	1.0928
2000	1196	1.6973	2000	1726	0.8879
2001	1303	1.4048	2001	1848	0.6521

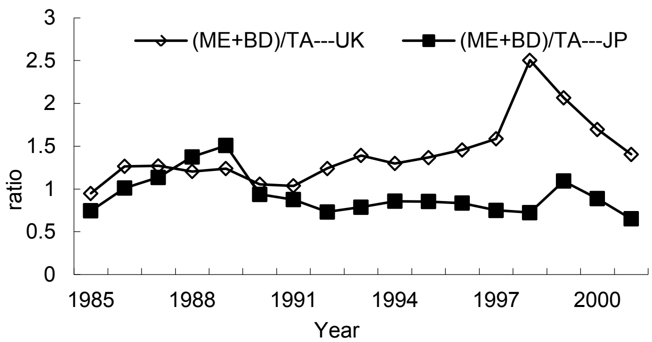
This table shows the average (ME+BD)/TA ratio of the UK and Japan for the years (1985-2001).

The average ratio for the UK firms exceeds one for every year, and sometimes even exceeds two. In contrast, the ratio for Japanese firms is usually below one, and only exceeds one in four years. From 1990, the (ME+BD)/TA ratio of the UK systematically exceeds one, and also exceeds that of Japan, as shown in

Figure 1. Related research by Rajan and Zingales (1995) demonstrated that bank-oriented firms such as those from Japan have better access to finance than market-oriented firms like those from the UK.

Figure 1 reports UK's and Japan's average (ME+BD)/TA ratio for the years 1985-2001. The figure shows the number and (ME+BD)/TA ratio for UK and Japan. We can not explain why the (ME+BD)/TA ratio of Japan is exceptionally high, even higher than UK in 1988 and 1989.

Figure 1: (ME+BD)/TA between UK and JP



This figure reports UK's and Japan's average (ME+BD)/TA ratio for the years 1985-2001.

The (ME+BD)/TA ratio is defined as the sum of the market value of equity and book value of debt divided by total assets. A ratio exceeding one means a firm is using finance that exceeds its real investment. Such a level also implies that the firm lacks sufficient financial capital reserves to exploit investment opportunities. Such firms rely on financial markets. Firms with (ME+BD)/TA ratios below one may be under-investing. Firms which have sufficient free cash can take investment opportunities.

According to the previously used (ME+BD)/TA ratio, UK firms resemble financially constrained firms, displaying high fixed asset ratio, high ratio of equity to total liabilities and shareholder equity, but low debt leverage. Meanwhile, Japanese firms typically display a lack of financial constraints, including low ratio of fixed to total assets, relatively low equity leverage, and high leverage ratio, but these characteristics necessarily indicate financially unconstrained firms. The criteria for a firm being financially constrained requires explicit definition, which this study defines as an (ME+BD)/TA ratio exceeding one. Since the (ME+BD)/TA ratio changes over time the analysis is time-dependent.

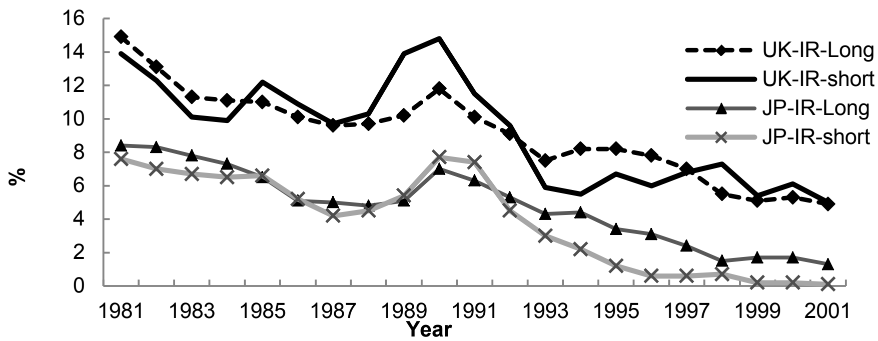
Change in debt to Assets Δ D/A is the difference of book value of debt between the t and t-1 periods divided by total assets for the period t is the change in book debt relative to total assets between t and t-1, which can explain fluctuation of debt. Δ D/A in Japan is relatively small and the standard deviation is about 10% for both small and large firms. Δ D/A in the UK has relatively large standard deviation, but the standard deviation is steady in large firms ranging from 20 to 40 percent. Small firm size may cause larger variation in book debt. Δ E/A (the difference of book equity between t and t-1 periods divided by total assets in period t). denotes the change of book equity, and is similar for firms from both countries. However, the standard deviation in the UK Significantly exceeds that in Japan. This finding is consistent with the above argument that firms facing financial constraints lack sufficient cash, causing large variation in equity and debt.

$$\Delta D/A_t + \Delta E/A_t + \Delta RE/A_t = \frac{D_t - D_{t-1}}{A_t} + \frac{E_t - E_{t-1}}{A_t} + \frac{RE_t - RE_{t-1}}{A_t} = \frac{A_t - A_{t-1}}{A_t}$$

The sum of $\Delta D/A_t + \Delta E/A_t + \Delta RE/A_t$ equals the change of assets at time t. Asset change in Japan is minimal compared with the UK, and is even negative during some periods during the 1990's because of the economic recession.

Figure 2 shows long-term and short-term interest rates from 1980 to 2001. Short term interest rates in Japan were near zero during the late 1990s. Leverage ratios steadily decreased during this period, partly explaining why the Δ D/A is small and changes only slightly during the 1990s. When market equity exceeds book equity, Market Leverage exceeds Book Leverage. Market leverage reduces with increasing market prices. The figure shows that interest rates change between the short and long-term.

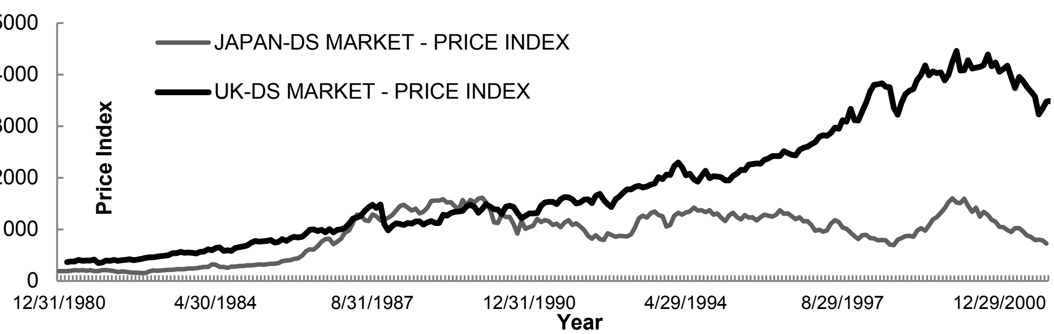
Figure 2: Long-term and Short-term Interest Rate



This figure shows long and short term interest rates from 1980 to 2001 in Japan and the UK

Market Price Indexes of Japan and UK, priced in Dollars were collected from DataStream, Item (TOTMJP$) and Item (TOTMKUK). Figure 3, shows the aggregated market price index. The UK market grew from the 1980s, and then declined during the late 1990s, possibly because of the bursting of the Internet Bubble. Market Leverage thus was much lower than Book Leverage during the 1990s. However, the situation in Japan was quite different, with a short period of growth in market price indexes during the late 1980s, followed by a slight decrease during the 1990s. Market price indexes suddenly climbed at end of the 1990s, before crashing in the early 2000s. Thus in Japan, just as in the UK, Market leverage was much lower than book leverage before 1990. Subsequently, the two leverage ratios drew closer together as a result of the low market price index.

Figure 3: Market Price Index



This figure shows the aggregated market price index for the UK and Japan.

RESULTS

This study tested the market timing effect using UK and Japanese firms. The results are shown in Tables 4, 5, 6 and 7. The market timing test sample of Baker *et el.* (2002) covered 20 years of UK data, but the timing effect was investigated only for ten years. The international firm sample used in this study runs only from 1985 to 2001. Thus *wmb* persistent effects are tested for only five years.

Table 4 lists the book leverage for UK firms. The first figure in each cell is the regression coefficient. Meanwhile, the second figure is the t-statistic. ***, ** and * indicate significance at the 1, 5 and 10 percent levels, respectively. Furthermore, for the Book Leverage regression, the b2/b1 ratio and b3/b1 ratio are increasing with time, partly consistent with *wmb* persistently explaining power on leverage. Several reasons must exist for the clear difference between Book Leverage and Market Leverage. Each cell shows the t-statistic.

Table 4: UK Leverage of Book Leverage

Year	b1	t(b1)	b2	t(b2)	b3	t(b3)	c3	t(c3)	b2/b1	b3/b1
t+1	0.2585	0.5807	0.2584	0.5807	0.2585	0.5807			1	1
t+2	0.3091	0.6666	0.0310	0.0949	-0.3157	-1.1295	0.4350	1.1640	0.1023	-1.0215
t+3	0.4529	0.9149	0.2072	0.5255	-0.1054	-0.3232	0.5846	1.4276	0.4577	-0.1227
t+4	0.4939	0.9323	0.5783	1.8943 *	-0.1274	-0.5439	0.7333	1.5011	1.1708	-0.2579
t+5	0.6782	1.1894	0.6041	1.8265 *	-0.0844	-0.2939	0.9985	2.0510 **	0.8908	-0.1243
t+6	0.6782	1.4926	0.5818	0.7714	-0.5540	-2.3120 **	1.2011	3.2335 ***	0.6177	-0.5882
t+7	1.0411	1.6271	0.2404	0.4368	-0.3328	-1.1967	1.3722	2.3621 **	0.2309	-0.3196
t+8	1.0861	1.5229	-0.2110	-0.5608	-0.6834	-2.4406 ***	1.0912	3.8267 ***	-0.1945	-0.6292
t+9	1.4132	1.8088 *	0.1549	0.5982	-0.2344	-0.5842	1.1142	2.8413 ***	0.1096	-0.1658
t+10	0.8666	1.0054	-0.1980	-0.4633	-0.4186	-1.4938	1.1602	3.1703 ***	-0.2291	-0.4830

*This table shows book leverage for UK firms. The first figure in each cell is the regression coefficient. Each cell shows the t-statistic. ***, ** and * indicate significance at the 1, 5 and 10 percent levels respectively.*

In Table 5, the sample of UK firms displays a significant time effect in *wmb.* The coefficients of b1, b2 and b3 are negative and significant in the market leverage regressions. Furthermore, the b2/b1 ratio decreases over time, indicating that the *wmb* effect exits and reduces over time. The b3/b1 ratio is near one, and increases slightly over time, demonstrating the importance of *wmb* in explaining leverage. The result is consistent with previous empirical evidence that firms time the market and adjust their leverage over periods of many years, and *wmb* can describe the discrepancies with target leverage. In contrast, the empirical evidence from the UK is quite different.

Table 5: UK Leverage of Market Leverage

Year	b1	t(b1)	b2	t(b2)	b3	t(b3)	c3	t(c3)	b2/b1	b3/b1
t+1	-0.9281	-2.4420 ***	-0.9281	-2.4421 ***	-0.9281	-2.4420 ***			1	1
t+2	-1.3783	-3.8543 ***	-1.8166	-5.1565 ***	-1.8277	-7.1723 ***	-7.3545	-13.3103 ***	1.3179	1.3259
t+3	-1.3118	-3.1648 ***	-2.1287	-6.4983 ***	-2.2872	-9.0572 ***	-7.9072	-9.6788 ***	1.6227	1.7435
t+4	-1.1608	-2.8156 ***	-2.0332	-9.0292 ***	-1.7325	-4.7255 ***	-8.9283	-7.8419 ***	1.7515	1.4924
t+5	-1.1456	-2.6809 ***	-2.3423	-8.2389 ***	-1.8520	-4.7343 ***	-9.2079	-10.385 ***	2.045	1.6166
t+6	-1.1067	-2.2625 **	-2.0164	-4.0156 ***	-1.8926	-4.0626 ***	-9.4914	-8.0524 ***	1.8219	1.7101
t+7	-1.2084	-2.5486 ***	-2.5764	-4.7435 ***	-2.1484	-5.2253 ***	-9.2644	-6.7935 ***	2.1319	1.7778
t+8	-1.2576	-2.1712 **	-2.3725	-2.6621 ***	-2.1363	-3.6683 ***	-9.2526	-6.3724 ***	1.8865	1.6987
t+9	-1.234	-1.7293 *	-2.0744	-1.9299 *	-1.6856	-2.7958 ***	-9.4138	-6.8834 ***	1.6809	1.3659
t+10	-1.656	-3.2305 ***	-1.6496	-1.678 **	-1.5588	-3.9408 ***	-8.6069	-7.5298 ***	0.9961	0.9473

*This table shows book leverage of market leverage for UK firms. The first figure in each cell is the regression coefficient. Each cell shows the t-statistic. ***, ** and * indicate significance at the 1, 5 and 10 percent levels respectively.*

First, the difference between Book Leverage and Market Leverage caused significant difference in their regressions. Figure 3 reveals that the UK market index rapidly climbed from the beginning of the 1990s, increasing variation between market leverage and book leverage for both small and large firms. However, these two forms of leverage drew close together for Japanese firms.

Second, an important question is why *wmb* significantly influences market leverage but not book leverage. We offer the following explanation. The *wmb* factor is closely related to market leverage and thus can effectively explain unadjusted leverage. Because *wmb* is the weighted average of market equity to book equity, and because market leverage is the ratio of book debt divided by total assets minus book equity plus market equity. Assuming firm retained earnings are zero, the first difference of Market Leverage to market-to-book ratio are clearly correlated. The *wmb* is the weighted average of several market-to-book ratios, and a complex relation appears to exist between them. However, further investigation is necessary to identify empirical evidence of this relation.

Table 6 lists the Japanese leverage of book leverage. The first figure in each cell is the regression coefficient, while the second figure is the t-statistic. ***, ** and * indicate significance at the 1, 5 and 10 percent levels respectively. Meanwhile, in the regression of book leverage, the b2/b1and b3/b1 ratios increase with time which is partly consistent with the persistent ability of *wmb* to explain leverage. The clear difference between these two kinds leverage may have several causes.

Table 6: Japan Firms Leverage of Book Leverage

Year	b1	t(b1)	b2	t(b2)	b3	t(b3)	c3	t(c3)	b2/b1	b3/b1
t+1	-2.7981	-6.167 ***	-2.7980	-6.167 ***	-2.7980	-6.167 ***			1	1
t+2	-2.9212	-6.121 ***	-2.6295	-5.709 ***	-3.080	-7.947 ***	4.1679	4.7973 ***	0.9001	1.0545
t+3	-2.9249	-5.78 ***	-2.4206	-5.157 ***	-3.3463	-9.546 ***	4.7506	5.8081 ***	0.8275	1.1440
t+4	-2.976	-5.433 ***	-2.2688	-4.713 ***	-3.5209	-10.57 ***	5.3281	7.3501 ***	0.7623	1.1831
t+5	-2.831	-4.803 ***	-2.0487	-3.976 ***	-3.3841	-10.75 ***	5.5506	7.1362 ***	0.7236	1.1953
t+6	-2.553	-4.262 ***	-1.8060	-3.874 ***	-3.3845	-9.214 ***	5.8514	6.4504 ***	0.7073	1.3257
t+7	-2.2801	-3.885 ***	-1.5088	-3.427 ***	-0.2575	-6.996 ***	6.3417	6.2706 ***	0.6617	1.4286
t+8	-1.9818	-3.662 ***	-1.0804	-2.254 ***	-3.258	-7.105 ***	6.4769	6.4544 ***	0.5452	1.6439
t+9	-1.7026	-3.356 ***	-0.8864	-1.559 ***	-3.1593	-8.026 ***	6.9193	6.9424 ***	0.8206	1.8556
t+10	-1.7288	-3.062 ***	-1.05832	-1.474 ***	-3.0509	-7.727 ***	6.2860	4.8067 ***	0.6121	1.7647

*This table shows leverage of book leverage for Japanese firms. The first figure in each cell is the regression coefficient. Each cell shows the t-statistic. ***, ** and * indicate significance at the 1, 5 and 10 percent levels respectively.*

Table 7 lists the JP leverage of Market leverage. The first figure in each cell is the regression coefficient. The second figure in each cell is the t-statistic. ***, ** and * indicate significance at the 1, 5 and 10 percent levels respectively. While in book leverage regression, the b2/b1 ratio and b3/b1 ratio are increasing with time, which is partially consistent with the persistent power of *wmb* to explain leverage. The clear difference between these two may have several causes.

Figure 4 shows the market-to-book ratio for UK and Japanese firms from 1985-2001. The figure shows the UK MB-ratio exceeds the Japanese MB-ratio between 1985 and 2001. Figure 4 shows that the trend of market-to-book ratio in the UK is stronger than in Japan. Figure 3 reveals that Market Price Index and average market-to-book ratio exhibit similar trends in both the UK and Japan.

Firms facing financial constraints often rely on raising equity. Hence, the leverage of such firms varies considerably, and is difficult to capture using the market timing testing equation. This phenomenon may contradict the argument regarding how *wmb* explains capital structure (Baker *et al.*, 2002). However, it is important to exclude other influences on leverage degree in further studies.

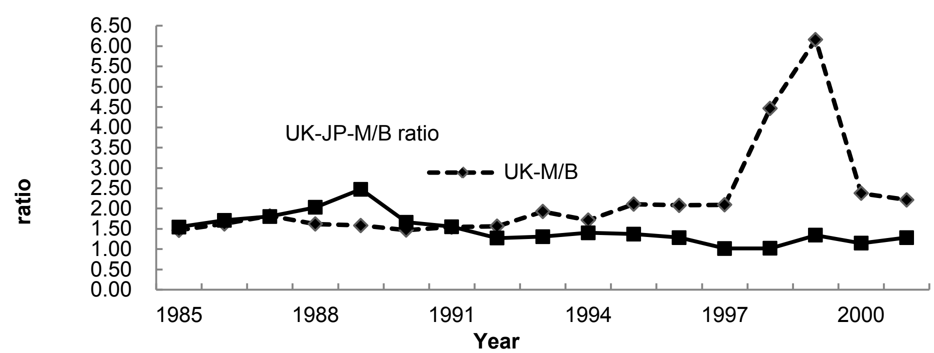
Table 7: Japan Leverage of Market Leverage

Market Leverage%														
Year	**b1**	**t(b1)**		**b2**	**t(b2)**		**b3**	**t(b3)**		**c3**	**t(c3)**		**b2/b1**	**b3/b1**
t+1	-3.0378	-5.7617	***	-3.0378	-5.7617	***	-3.0378	-5.76	***				1	1
t+2	-3.0169	-5.258	***	-2.8108	-4.55	***	-3.1801	-6.598	***	-12.6818	-15.8283	***	0.9316	1.0540
t+3	-3.1077	-5.049	***	-2.8863	-4.602	***	-3.5306	-8.037	***	-12.6722	-16.2739	***	0.9287	1.1361
t+4	-3.1204	-4.801	***	-2.8985	-4.521	***	-3.6779	-8.285	***	-12.8633	-15.2717	***	0.928	1.1789
t+5	-2.9997	-4.2954	***	-2.6884	-3696	***	-3.5847	-6.423	***	-13.2616	-14.5537	***	0.8962	1.1950
t+6	-2.7194	-3.762	***	-2.5278	-3.163	***	-3.6088	-6.232	***	-13.5667	-13.976	***	0.9295	1.3270
t+7	-2.3966	-3.378	***	-2.3735	-2.593	***	-3.4764	-5.074	***	-14.135	-13.2897	***	0.9903	1.4505
t+8	-1928	-3.232	***	-2.2182	-2.731	***	-3.5376	-7.778	***	-14.8987	-14.6943	***	1.1505	1.8348
t+9	-1.4789	-3.4133	***	-2.0463	-3.226	***	-3.4154	-15.33	***	-15.2487	-15.4939	***	1.3836	2.3094
t+10	-1.3454	-3.083	***	-2.4085	-2.681	***	-3.5446	-10.43	***	-16.3309	-22.7822	***	1.7902	2.6346

*This table shows leverage of market leverage for Japanese firms. The first figure in each cell is the regression coefficient. The second figure is the t-statistic. ***, ** and * indicate significance at the 1, 5 and 10 percent levels respectively.*

From the perspective of optimal dynamic argumentations, when firm leverage is homogeneously optimized, comparison of *wmb* coefficients between the equations is meaningful. The example involving Japanese firms lacks the timing effect and can be largely explained by the *wmb* factor in. However, if the leverage is not well-explained in Eq. 4 the same factors also cannot explain better in Eq. 6. Any comparison then becomes too weak. Only when b1 is significant does comparison with b2 and , b2/b1become meaningful. However, b3 in Eq. 3 does not represent the same case, and the *wmb* effect intensifies with time. Once leverage has been optimized, although the real adjustments are not entirely in *wmb*, it retains significant explanatory power.

Figure 4: United Kingdom and Japan Market to Book Ratio



This figure shows the United Kingdom and Japan Market to Book ratio.

CONCLUSIONS

This study used international data to investigate the impact of different financial market characteristics on firm leverage. The Tobin Q regression methodology was used in this study. This study divided the firm sample using the (ME+BD)/TA definition, and classified firms in the UK as financially constrained, and those in Japan as financially unconstrained. This investigation also investigated the characteristics of these two types of firms. Firms in the UK behave similarly to financial constrained firms, with high fixed asset ratio, high equity ratio and large variation of leverage, even excluding size factor. Firms in Japan exhibit low fixed assets and low rate of change of debt ratio, but high leverage.

This study concludes that more evidence using internal samples is required before concluding the findings apply elsewhere. Second this study finds that *wmb* is related to market leverage both by difference and by displaying similar trends over many years. The study shows that *wmb* can be influenced by unexplained market leverage, but not book leverage. On the other hand, book leverage may be more closely related to true debt value, which is little affected by market timing. Third, leverage in Japanese firms is relatively steady, and the timing effects remain significant even after many years. From the perspective of dynamic adjustment leverage, when leverage is optimized the degree of previous adjustments are largely explained in *wmb*.

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THE CONSOLIDATION OF THE GLOBAL BREWING INDUSTRY AND WEALTH EFFECTS FROM MERGERS AND ACQUISITIONS

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ABSTRACT

The brewing industry has recently experienced increased merger activity. This paper analyzes the short-term wealth effects of horizontal mergers and acquisitions on acquirers in the brewing industry. Based on a sample of 69 takeover announcements between 1998 and 2010, significant positive announcement returns were identified. In addition, the study finds significant positive returns for domestic transactions as well as cross-border deals involving targets in emerging markets. Other identified drivers of short-term success include transaction size, acquirer size and the target's public status. Furthermore, significant negative rival effects are identified across leading brewing groups, when missing a potential M&A opportunity.

JEL: G14, G34, Q14

KEYWORDS: Mergers and Acquisitions, Brewing Industry, Announcement Returns, Acquirers, Industry Rivals, Event Study

INTRODUCTION

Over the last decades, the wealth implications of mergers and acquisitions (M&A) have been widely discussed in empirical M&A research.Studies focusing on short-term announcement effects unambiguously conclude that M&A create value for shareholders of target companies (Bradley et al., 1988). However, the situation is not as clear-cut, as a closer look at returns to acquiring companies shows a different pattern: While overall acquirer returns average around zero (Bruner, 2002), industry specific event studies provide mixed findings of negative abnormal acquirer returns, positive abnormal returns or acquirer returns that are not significantly different from zero. Besides measuring the performance of the merging firms, there is also growing interest in the wealth effects of M&A on other firms from the same industry. Existing evidence shows that rival companies gain at the M&A announcement due to positive information signaling effects (See Eckbo (1983), Fee and Thomas (2004), Sharur (2005), Song and Walking (2000)). Towards the turn of the last century, many industries including the brewing industry,

have experienced a sharp increase in M&A activity. Consolidation has and continues to be a major trend in the sector as multi-national breweries seek to expand their activities into new emerging markets. At the same time, declining mature markets (in particular Western Europe) and resulting pressure on profit margins have encouraged brewers to engage in M&A, in order to gain in scale and benefit from synergies. In contrast to many other sectors, the production, distribution and marketing of beer is characterized by a relatively high fixed cost base, resulting in high levels of operational leverage (Earlam et al., 2010) providing larger brewers with material size advantages. Moreover, increased size has enabled brewers to exercise a significant amount of market-power (Schwankl, 2008), as larger brewers are able to negotiate favorable terms with their suppliers and benefit from greater bargaining power for negotiations with retail customers. Hence, it is not surprising that the global beer market today is dominated by large national/multinational brewers rather than local, regional brewers. The four largest brewers Anheuser-Busch Inbev, Heineken, SABMiller and Carlsberg ("the big four") control about 50% of the global beer market. Despite rising market concentration, competition among the large brewing groups has remained fierce (Iwasaki et al., 2008). Going forward, many industry experts predict that the consolidation process will continue and that the "big four" will increase their control to 75% of the global beer market (Jones, 2010). As sector debt levels are expected to decrease further, research analysts are certain that M&A will remain a major theme in the coming years, as brewers will continue their quest for suitable M&A targets (Earlam, et al., 2010). In the light of these specific industry characteristics and the recent developments in the sector, the question arises whether the global synergy and efficiency potential of M&A transactions are reflected by capital markets in the form of abnormal stock price reactions to acquiring and close rival companies. Even though the global beer industry has gone through significant consolidation and seen a lot of M&A in recent years, empirical evidence remains scarce.

Therefore, the aim of this study is to fill this research gap and provide empirical evidence for investors and managers of beer companies. In contrast to previous research, our study determines the short-term performance of brewing companies based on a global dataset, uses a multi-factor model (Fama-French 3 Factor model) to determine statistically reliable indications of short-term performance, and analyses a comprehensive list of deal, acquirer and target characteristics for their impact on the short-term wealth effect to acquiring companies. Moreover, our study specifically analyzes rival effects among the "big four". The objective of the study is twofold: Firstly, we aim to update and extend previously published announcement effects on acquirers in the brewing industry. The main focus lies in analyzing short-term return patterns in order to detect and categorize determinant variables. Secondly, we aim to provide empirical evidence regarding M&A announcement effects among the "big four" as well as the value implications to rivals.The remainder of this paper is structured as follows: Section 2 gives a brief overview of the relevant literature and outlines the derived hypotheses. Section 3 provides details on the applied methodology as well as the sample selection procedure. The following section 4 presents the empirical results and elaborates on the derived hypotheses. Finally, section 5 summarizes the findings and concludes.

LITERATURE REVIEW

While researchers unambiguously conclude that overall the announcements of mergers and acquisitions have positive value effects, Bradley et al. (1988) find that the short-term value creation is mostly attributed to the shareholders of target firms, which benefit from premiums paid by acquirers. In contrast, studies, that analyze acquirer returns, provide evidence of short-term value losses: Examining 4,265 M&A transactions between 1973 and 1998, Andrade et al. (2001) report insignificant negative returns to acquiring companies during a 3-day event window surrounding the announcement date of the transaction. Likewise, Loughran and Vijh (1997) document short-term acquirer returns that are overall negative, or insignificant. After reviewing 44 separate studies, investigating short-term value effects on acquiring companies, Bruner (2002) comes to the conclusion that on average abnormal returns for acquirers are essentially zero. In addition to cross-industrial studies, numerous industry specific analyses, in many cases seem to confirm the general results: Beitel et al. (2004) report negative abnormal acquirer returns for banks, while Akdogu (2009) and Berry (2000) find negative acquirer revaluations for the telecommunications and electric utilities industry, respectively. At the same time, alternative evidence identifies certain industries where acquirers are able to realize significant positive short-term returns. For example Mentz and Schiereck (2006) using a sample of 201 M&A transactions in the automotive supply industry document significant abnormal returns to acquirers of +1.6% during a 10-day event window and argue this finding to be the result of the extraordinary synergy potential in the industry perceived by capital markets. Similarly, Choi and Russel (2004) report positive abnormal returns to acquirers in the construction industry. Obviously, cross-industrial studies cover industry specific divergences, which result from the unique industry logic of value chains.

Empirical research also investigates the impact of M&A on rival firms. Overall, findings show positive as well as negative effects on rivals: On the one hand, rival companies may benefit from the M&A announcement due to a positive signaling effect regarding industry attractiveness and future takeover activity (Eckbo, 1983), (Song & Walkling, 2000). At the same time a merger in the industry decreases the number of competitors and thus increases the likelihood of collusion, which may lead to greater monopoly rents to rival firms (Eckbo, 1983), (Shahrur, 2005). Moreover, Snyder (1996) argues that rival firms may benefit from greater buyer power due to increased competition among suppliers, which may lead to lower input prices. On the other hand, rival firms may be affected by negative competitive effects as a result of more-intense competition in the industry due to a new, more-efficient combined firm (Eckbo, 1983). Overall, the documented positive effects outweigh the negative competitive effects: For example, Eckbo (1985) reports positive announcement effects to rivals for horizontal transactions. Similar results are found by Song and Walking (2000) in a study including horizontal and non-horizontal transactions. More recent studies by Clougherty and Duso (2009), Fee and Thomas (2004) and Shahrur (2005) confirm these results. As mentioned above empirical research on M&A in the brewing industry remains scarce and primarily focuses on the US brewing industry. The specific topics addressed in the studies focus on technological change in the sector (Kerkvliet et al., 1998), its tendencies towards concentration (Lynk, 1985), (Adams, 2006), the determinants and motives for horizontal M&A (Tremblay & Tremblay, 1988), as well as competition in the industry (Horowitz & Horowitz, 1968). More recently, Ebneth and Theuvsen (2007) analyze the short-term value effects of M&A to acquirers using event study methodology. Based on a sample of 29 cross-border transactions involving European acquirers from 2000-2005, they find insignificant positive acquirer returns of 0.9% in the 5-day event window surrounding the announcement date of the transactions.

With this paper, we aim to contribute to and extend existing literature with respect to geographical scope and methodology applied. First, we provide an analysis using a dataset, that in addition to cross-border transactions also includes domestic acquisitions and overall the merger wave of recent years. This database enables us to cover the complete M&A-cycle where usually later transactions significantly differ from the early ones. Second, we use a multi-factor (Fama French 3-factor model) model to determine and measure abnormal performance, which is then tested for significance using parametric and non-parametric statistical methods as well as multivariate regression analyses. Our main research interest is concentrated on the following aspects:

Acquirer Announcement Effects

As pointed out above, previous event studies focusing on single industries predominantly report negative acquirer returns. On the other hand, certain industries have been identified as outlier industries reporting positive abnormal acquirer returns. Given the particularities of the brewing industry and its development over the last few years, we expect M&A in the sector to be a feasible measure to realize synergy- and efficiency gains. Accordingly, we assume the capital markets to reflect the industry-specific synergy potentials, resulting in positive short-term value effects to acquirers. Since our study is based on a larger dataset than used by Ebneth and Theuvsen (2007) and additionally considers global as well as domestic transactions, we assume significant positive abnormal returns to acquiring brewers.

Analysis of Determinants of Acquirer Returns

While various studies focus on the impact of geography on short-term performance, the reported results are mixed: Several studies on cross-border M&A have found significant positive value gains to investors of acquiring firms around the announcement date (see e.g. Zhu and Malhotra (2008), Goergen and Renneboog (2004), Morck and Yeung (1992)). On the other hand, there have also been studies documenting negative or insignificant gains to acquirers involved in cross-border M&A (see e.g. Datta and Puia (1995) and Eckbo and Thorburn (2000)). Negative acquirer returns are also reported for transactions involving targets in emerging market economies (Williams & Liao, 2008). In case of the brewing industry, Ebneth and Theuvsen (2007) find insignificant positive acquirer returns for cross-border transactions. However, their sample is restricted to 29 transactions. Due to the continuous decline in beer volumes in many mature markets, we regard cross-border M&A as a viable strategic option to diversify into international markets and thus expect significantly positive abnormal returns to acquirers. In particular, we assume acquirer returns to be positively impacted if the targets are based in emerging market economies and expect to find significant differences compared to domestic transactions.

The brewing industry has seen a significant increase in transaction volumes in recent years. Given the particular characteristics of the sector, increased company size provides brewers with a material competitive advantage. Analyzing technological change and economic efficiency in the U.S. brewing industry, Kerkvliet, et al. (1998) report substantial increases in economies of scale. Consequently, it can be argued that the acquisition of big targets will significantly contribute to the success of a transaction due to greater potential for economies of scale and revenue and/or cost synergies. On the other hand, the integration of larger targets may be more difficult than for small targets (Hawawini & Swary, 1991). In addition to target size, the size of the acquirer may also influence the success of a transaction. Comprehensive studies by Asquith et al. (1983), Jarrell and Poulsen (1989) and Moeller et al. (2003) find a negative impact of acquirer size on acquirer returns. Moeller et al. (2003) argue that managers of larger companies are more likely to overestimate their own abilities. Due to the fact that larger companies may benefit from bigger cash reserves and might need to face less hurdles in the execution of M&A transactions, the authors argue that managers of larger companies are more likely to engage in M&A transactions that are not always beneficial to the company. Consequently, we expect to find significant differences in abnormal returns between large and small acquirers

Over the last few years, the brewing industry has experienced a sharp increase in M&A activity and seen strong industry consolidation. As a consequence, the competitive landscape has materially changed as the "big four" today control more than 50% of global beer volumes. While the sector is expected to further consolidate in the future, it is becoming increasingly difficult to find suitable targets (Gibbs et al., 2010). We assume the changes in market structures and concentration to have an impact on acquirer returns and expect to find significant differences in abnormal acquirer returns over time.

Several studies analyze the impact of the method of payment on acquirer returns. Myers & Majluf (1984) argue that bidders prefer to pay using stock when they believe that the market overvalues their shares and on the other hand prefer using cash, when they regard their stock as undervalued. Similarly, Martynova and Renneboog (2006) suggest that the means of payment used is an important signal of the quality of the target firm and its potential synergy value. They argue that a cash offer by the bidding company signals a willingness to pay off target shareholders in order to avoid sharing future cash flows and bear the sole risk of the combined firms. On the other hand, an all-equity offer signals the willingness to keep the target shareholders involved in the merged company and share its risk. The theoretical framework behind the mentioned signaling effects is supported by the results of different studies (see e.g. Brown and Ryngaert (1991) and Wansley et al. (1983)) that report a significantly negative market reaction following the announcement of equity offerings, in contrast to positive announcements returns in the case of cash offers. We expect to find similar results for the brewing industry.

The acquisition of privately held companies accounts for the majority of M&A transactions in the brewing industry. The general consensus among researchers is that bids for privately held companies generate higher bidder returns than bids for publicly held companies. Martynova and Renneboog (2006) argue that, in the case of privately held targets, bidders are likely to benefit from price discounts as compensation for buying a comparably illiquid stake. At the same time, they see advantages due to the fact that private companies usually have fewer shareholders, which facilitates negotiations. These theoretical assumptions are confirmed in studies from Moeller et al. (2003) and Faccio et al. (2006) who report substantially higher announcement returns to acquirers for bids on privately held targets as opposed to bids on public firms. Consequently, we expect to find similar results for the brewing industry.

In a study focusing on short-term value effects of acquirers Haleblian and Finkelstein (1999) determine a positive relation between the number of completed transactions of an acquirer and the magnitude of the acquirer's abnormal return. They argue that with each completed transaction, acquirers gain experience in the integration of targets, which can be, leveraged in future transactions. In case of the brewing industry, which is dominated by large brewing groups that frequently engage in M&A activities, we expect to find greater returns for bidders with transaction experience.

<u>Announcement and Rival Effects among the "Big Four"</u>

The global beer market is dominated by the "big four". Given the strong competition among them and due to the increased difficulty to find suitable targets, we expect to find positive acquirer returns if one of the "big four" announces an M&A transaction. At the same time, we expect the remaining three rival companies to be negatively impacted by the announcement as they are put into a disadvantageous competitive position by missing out on a potential M&A opportunity in a fairly concentrated market.

DATA AND METHODOLOGY

The sample of mergers and acquisitions for the event study is drawn from the Securities Data Corporation (SDC)/ Thomson One Banker Deals database and the Merger Market M&A database. It includes all worldwide M&A events announced between January 1st, 1998, and September 1st, 2010. The total number of M&A deals is reduced to yield only those transactions meeting the following criteria:

1. At the time of the transaction, acquirer and target companies both had active operations in the brewing industry.

2. The acquiring company has been publicly listed for at least 250 days prior to the announcement of the transaction.

3. The total transaction value accumulates to at least USD 50 million.

4. The completion of the transaction leads to a change of control in the target; Prior to the announcement of the transaction the bidder holds less than 50% in the target company, following the transaction the bidder owns a controlling stake in the target company.

5. The transaction has been successfully completed.

In addition, the transactions were validated by a press research using the Factiva database as well as company websites in order to ensure that all transactions are horizontal and the announcement dates provided by the databases are correct. Moreover, acquirers with multiple transactions on the same day were removed from the dataset. The described selection criteria result in a final sample of 69 transactions. The frequency distribution of the transactions over time is provided in Table 1. While the number of transactions is spread fairly even over the years, the average transaction size varies strongly from 268 USD mil. to 11.973 USD mil. due to a number of high-profile transactions such as InBev's acquisition of Anheuser-Busch (52 USD bil.), Heineken and Carlsberg's takeover of S&N (USD 15 bil.) and Heineken's recent acquisition of Femsa Cerveza (USD 5.7 bil). In terms of geography, more than 80% of the transactions involve acquirers that are based in Europe.

The relevant daily stock prices, market capitalizations and local market indices for acquirers were downloaded from the Thomson Datastream database. Acquirer returns are calculated using the Datastream Total Return Index, which adjusts the closing share prices for dividend payments as well share issuances or repurchases. Moreover, global value and growth indices for large and small cap companies from data and index provider Russell serve as proxies for the Fama French model.

Table 1: Sample Overview: Descriptive Statistics

			Avg. Trans. Val.	Trans. Val.	Acquirer Region - Number of Deals			
Year	**Deals**	**(%)**	**(USD mil.)**	**(USD mil.)**	**Europe**	**Americas**	**Asia**	**RoW**
2010	1	1.4	5,700	5,700	1			
2009	4	5.8	737	2,946	2		2	
2008	6	8.7	11,973	71,835	6			
2007	5	7.2	405	2,025	1	2	1	1
2006	6	8.7	332	1,991	6			
2005	8	11.6	785	6,281	8			
2004	9	13.0	875	7,873	6	2	1	
2003	6	8.7	582	3,489	6			
2002	6	8.7	1,526	9,157	5	1		
2001	4	5.8	599	2,396	3	1		
2000	6	8.7	514	3,086	5		1	
1999	4	5.8	495	1,981	3	1		
1998	4	5.8	268	1,071	4			
Sum	**69**	**100.0**	**1,736.7**	**119,831.0**	**56**	**7**	**5**	**1**

This table provides the frequency distribution of the M&A transactions in the sample. It includes all successfully completed transactions between 1998 and 2010 where at the time of the transaction, acquirer and target companies both had active operations in the brewing industry, the acquiring company has been publicly listed for at least 250 days prior to transaction announcement, the total transaction value accumulated to at least USD 50 mil. and the bidder through the transaction acquired a controlling stake in the target company. Additionally, total and average transaction values (in USD mil.) and details on acquirer region are provided.

RESEARCH METHODOLOGY

In order to determine and analyze short-term announcement effects our study applies event study methodology. Event studies have a long history that goes back to the 1930s (see MacKinley (1997)). Since then the methodology has become more and more sophisticated and found its application in empirical research on M&A. In particular, it has become a widely accepted tool to analyze the short-term value effects of M&A transactions. In our study, we assess short-term announcement returns using traditional event study methodology as for example described by Brown and Warner (1985) in connection with the "Fama-French-3-Factor-model" (FF3F). The use of the multi-factor FF3F model enables us to more accurately detect and determine abnormal performance than with a single factor market model as used by Brown and Warner (1985). Formula (1) shows how the abnormal returns were derived:

$$E(R_i) = R_f + b_i\left[E(R_m) - R_f\right] + s_i E(SMB) + h_i E(HML) \qquad (1)$$

where $E(R_i)$ is the expected return on asset i, R_f is the return on the risk-free asset, $E(R_m)$ is the expected return on the market portfolio, $E(SMB)$ is the expected return on the mimicking portfolio for the "small minus big" size factor and $E(HML)$ is the expected return on the mimicking portfolio for the "high minus low" book-to-market factor.

Fama and French (1992), in probably one of the most influential papers in the area of asset pricing in the past decade, argue that the single factor Capital Asset Pricing Model of Sharpe (1964) and Lintner (1965) has little ability to explain the cross-sectional variation in equity returns. They find that two other factors related to fundamental variables, namely size and the ratio of book equity to market equity, have strong roles in explaining variation in cross-sectional returns. In our study, the multi-factor FF3F model is used to determine abnormal returns of acquiring and rival companies by regressing a time series of the companies' excess returns (return less risk-free rate) with the time series of market excess returns, the time series of the difference in returns of small and big companies (SMB), and the time series of differences in returns of companies with high and low (HML) market-to-book values (formula 2).

$$R_{i,t} - R_{f,t} = \alpha_i + b_i\left(R_{M,t} - R_{f,t}\right) + s_i SMB_t + h_i HML_t + e_{i,t} \quad (2)$$

where $R_{i,t}$ is the realized return on asset i at time t, $R_{f,t}$ is the realized return on the risk-free asset at time t, $R_{M,t}$ is the realized return on the market portfolio at time t, SMB_t is the realized return on the mimicking portfolio for the size factor at time t and HML_t is the realized return on the mimicking portfolio for the book-to-market factor at time t.

The return of the market portfolio within the model usually refers to a market index that is associated with the particular security. In order to account for regional differences in industry returns and country-specific risk profiles our study determines local indices for each acquirer in the sample. For example, the DAX 30 index is used for German acquirer companies and the FTSE All Shares index is used for UK-based acquirer companies within the sample. Our study uses the 3-month US T-bill rate as a proxy for the risk free rate. The difference in returns of small and big companies as well as the difference in returns of companies with high and low market-to-book ratios is determined using global Frank Russell style portfolios as proposed by Faff (2003). The Russell style portfolios are utilized to create proxies for the Fama and French SMB and HML factors. Specifically, the style indices chosen are: (a) Global Russell large-cap Growth Index, (b) Global Russell large-cap Value Index, (c) Global Russell small-cap Growth Index, (d) Global Russell small-cap Value Index.

The Global Russell large-cap Growth Index (The Global Russell large-cap Value Index) measures the performance of the largest global companies with higher (lower) price-to-book ratios and higher (lower) forecasted growth values. Similarly, the Global Russell small-cap Growth Index (The Global Russell small-cap Value Index), measure the performance of global small-cap companies with higher (lower) price-to-book ratios and higher (lower) forecasted growth.

Having determined all relevant factors we finally estimate the acquirer return model by using a multivariate Ordinary Least Squares (OLS) regression over a 230 day estimation period starting at trading day t=-250 relative to the announcement date of the transaction. Finally, on the basis of these estimated FF3F Model parameters, we calculate the abnormal returns for all acquirer companies using different event windows. To test for statistical significance of acquirers' abnormal returns this study employs three test statistics. First, we apply a simple parametric t-test. Second, we use a cross-sectional test as proposed by Boehmer, Musumeci and Poulsen (1992). The cross-sectional test is commonly used in event study literature as it accounts for a potential event-induced increase in standard deviation. Third, since non-parametric test statistics can be more powerful than parametric t-statistics (see Serra (2002), Barber & Lyon (1996)), we apply the Wilcoxon Signed Rank test to provide for a thorough statistical review.

EMPIRICAL RESULTS

In the following, the empirical results of our analyses are presented. We start off reporting the results for the total acquirer sample. In order determine potential drivers of abnormal performance we then report the results of the univarate and multivariate analyses. Finally, we specifically present announcement and rival effects among the “big four”.

Acquirer Announcement Effects

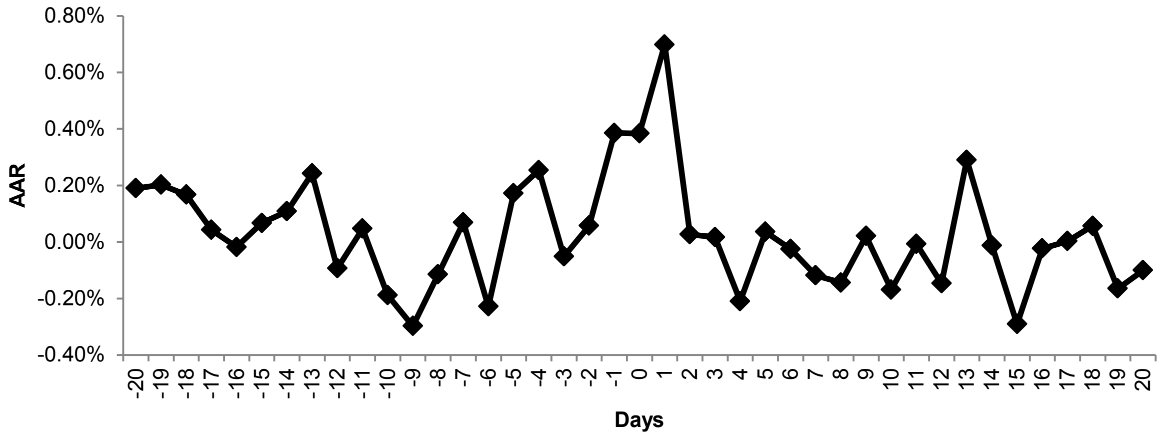
Table 2 reports the short-term announcement effects of M&A transactions on the total sample of acquirers in the brewing industry. The results show that acquirers earn a significant 1.77% in the [-5;5] and 1.47% in the [-1;1] event windows surrounding the announcement date. As Figure 1 shows, the abnormal returns peak following the day of the announcement of the M&A transaction. In case of the [-5;5] event window the results are significant at the 5%-level for the t-statistics as well as the cross-sectional test, and significant at the 10%- level for the Wilcoxon test statistics. In case of the [-1;1] event window the results are significant at the 1%-level for the t-statistics and significant at the 5%-level for the cross-sectional test. These findings are in line with the expected results and confirm the exceptional characteristics of the brewing industry. The short-term value effects show that capital markets in fact value the extraordinary synergy potentials in the brewing sector.

Table 2: Cumulative Average Abnormal Returns to Acquirers

Acquirers (N=69)										
Event-Window	**CAAR**	**t-Test**			**z-Test**			**WCX Test**		
		t-value	**p-value**		**z-value**	**p-value**		**z-value**	**p-value**	
[-20; 0]	1.48%	1.66	0.10	*	2.04	0.05	**	1.38	0.17	
[-10; 0]	0.44%	0.60	0.55		0.67	0.51		0.40	0.69	
[-5; 0]	1.20%	1.96	0.05	**	1.79	0.08	*	1.43	0.15	
[-1; 0]	0.77%	1.52	0.13		1.22	0.22		0.53	0.60	
[0]	0.38%	0.93	0.36		0.70	0.49		0.22	0.82	
[0; +1]	1.08%	2.16	0.03	**	1.84	0.07	*	1.41	0.16	
[0; +5]	0.95%	1.56	0.12		1.43	0.16		1.30	0.19	
[0; +10]	0.52%	0.70	0.49		0.79	0.43		0.83	0.41	
[0; +20]	0.12%	0.12	0.90		0.42	0.68		0.11	0.91	
[-1; +1]	1.47%	2.51	0.01	***	2.23	0.03	**	1.41	0.16	
[-5; +5]	1.77%	2.37	0.02	**	2.26	0.03	**	1.81	0.07	*
[-10; +10]	0.57%	0.57	0.57		0.81	0.42		0.17	0.86	
[-20; +20]	1.13%	0.87	0.39		1.47	0.15		0.42	0.68	

*This table shows the cumulative average abnormal returns (CAAR) to acquiring companies in mergers and acquisitions in the brewing industry. It contains all public acquirers whose trading data was available between 250 before and 20 days after transaction announcement. Statistical significance at the 10%, 5% and 1% level is denoted by *, ** and *** respectively. The statistical significance has been tested using a standard t-test, the cross-sectional test as proposed by Boehmer et al. (1992) (z-Test) and the Wilcoxon Signed Rank Test as described by Barber and Lyon (1996).*

Figure 1: AARs of Acquirers Surrounding Announcement of the Transaction



This figure provides daily abnormal average returns to acquiring companies of mergers and acquisitions in the brewing industry between 20 days before and 20 days after transaction announcement. It contains all public acquirers whose trading data was available between 250 before and 20 days after transaction announcement.

The Robustness of Results with Respect to Deal, Acquirer and Target Characteristics

Geographical Scope: In order to analyze the impact of geographical diversification on acquirer returns we compare domestic transactions with cross-border transactions and emerging market transactions.

Table 3: Abnormal Returns to Acquirers Differentiated by Geographical Scope

	Cross-Border		Domestic		Mean Comparison	
N	49		20			
EVENT WINDOW	CAAR	MEDIAN	CAAR	MEDIAN	ΔCAAR	ΔMEDIAN
[-1; +1]	1.02% *	0.54%	2.56%	-0.81%	-1.54%	1.35%
[-5; +5]	0.86%	0.20%	4.00% **	1.79% **	-3.14% *	-1.59%
[-10; +10]	-0.70%	0.11%	3.69% *	2.05%	-4.39% **	-1.94%
[-20; +20]	-0.36%	-2.41%	4.80% **	4.77% *	-5.16% *	-7.18%

*This table shows the cumulative average abnormal returns (CAAR) to acquiring companies involved in cross-border and domestic mergers and acquisitions in the brewing industry. Statistical significance at the 10%, 5% and 1% level is denoted by *, ** and *** respectively. The statistical significance has been tested using the cross-sectional test as proposed by Boehmer et al. (1992) (z-Test) and the Wilcoxon Signed Rank Test as described by Barber and Lyon (1996).*

Table 4: Abnormal Returns to Acquirers Differentiated by Target Region

	Emerging Market		Domestic		Mean Comparison	
N	27		20			
EVENT WINDOW	CAAR	MEDIAN	CAAR	MEDIAN	ΔCAAR	ΔMEDIAN
[-1; +1]	1.05%	0.46%	2.56%	-0.81%	-1.51%	1.27%
[-5; +5]	1.20% *	1.46%	4.00% **	1.79% **	-2.80%	-0.32% *
[-10; +10]	1.94% **	2.57%	3.69% *	2.05%	-1.75%	0.52%
[-20; +20]	1.52%	0.09%	4.80% **	4.77% *	-3.27%	-4.68% *

*This table shows the cumulative average abnormal returns (CAAR) to acquiring companies involved in emerging market and domestic mergers and acquisitions in the brewing industry. Transactions are classified as emerging market transactions if the acquired target is based in Latin America, Asia (ex Japan) or Eastern Europe. Statistical significance at the 10%, 5% and 1% level is denoted by *, ** and *** respectively. The statistical significance has been tested using the cross-sectional test as proposed by Boehmer et al. (1992) (z-Test) and the Wilcoxon Signed Rank Test as described by Barber and Lyon (1996).*

Tables 3 and 4 present the findings about the impact of geographical diversification on short-term acquirer performance. Overall, acquirers in the brewing industry show a preference for cross-border transactions and in particular emerging market transactions. In our total sample of 69 transactions, only 29% (20 transactions) are domestic/national transactions, while 71% (49 transactions) are cross-border transactions. Approximately 55% (27 transactions) of the cross-border transactions qualify as emerging market transactions and involve targets that are based in Latin America, Asia (ex Japan) or Eastern Europe. On average domestic acquirers in almost every case show positive value effects upon the announcement of the transaction with CAARs and Medians of CARs ranging between 1.79% and 4.80%. Despite the small sample size of only 20 transactions, many of these returns are significant on the 5% and 10% level. On the other hand, acquirers in cross-border transactions show mixed effects with CAARs and Medians of CARs ranging between -2.41% and 1.02% across various event-windows. A comparison of means shows that CAARs for cross-border transactions are significantly lower for the [-5;5], [-10;10] and [-20;20] event windows. Despite the relatively small sample size, these results indicate that capital markets seem to favor domestic over cross-border transactions. While these results stand in contrast to our predictions, they are in line with studies by Datta and Puia (1995) and Eckbo and Thornburn (2000).

Nevertheless, the results will be challenged in the multivariate analysis due to the relative small amount of domestic transactions. Table 4 compares domestic transactions with emerging market transactions. On average acquirers gain in all event windows upon the announcement of emerging market transactions with value gains ranging between 0.09% and 2.57%. Despite the small sample size of 27 transactions, acquirer CAARs gains are significant on the 5% and 10% level for the [-10;10] and [-5;5] event windows respectively. These results stand in contrast to the findings of Williams and Liao (2008), who report negative abnormal returns to acquirers in emerging market transactions in the banking industry. Again, due to the limited amount of transactions, these results will be challenged in the multivariate analysis.

Size of Transaction and Acquirer: In order to test for the incremental effect of transaction size on acquirer returns the sample was divided into two subsamples containing the 30 largest transactions and 30 smallest transactions by deal volume. The results are summarized in Table 5. On average, acquirers in small transactions yield positive CAARs between 0.94% and 2.47% across all event windows. In case of the [-5;5] event window, acquirers yield a positive 1.52% which is significant at the 5% level. On the other hand, acquirers in large transactions experience mixed value effects across various event windows with insignificant CAARs ranging from -2.67% to 1.14%. A comparison of means shows a significant underperformance of acquirer returns in case of large transactions for the [-10;10] and [-20;20] event windows. While these results stand in contrast to the expected results, they clearly serve as an indication and will be tested in the multivariate regression model.

Table 5: Abnormal Returns to Acquirers Differentiated by Transaction Size

	Top 30		Bottom 30		Mean Comparison	
N	30		30			
Event Window	CAAR	MEDIAN	CAAR	MEDIAN	ΔCAAR	ΔMEDIAN
[-1; +1]	1.14%	0.59%	0.94%	-0.66%	0.19%	1.25%
[-5; +5]	0.88%	-1.24%	1.52% **	1.29% **	-0.64%	-2.53%
[-10; +10]	-2.23%	-1.78% *	1.33%	0.94%	-3.56% *	-2.71% **
[-20; +20]	-2.67%	-4.41% *	2.47% *	0.23%	-5.14% *	-4.64% ***

*This table shows the cumulative average abnormal returns (CAAR) to acquiring companies for the top 30 and bottom 30 transactions by transaction volume in the sample. Statistical significance at the 10%, 5% and 1% level is denoted by *, ** and *** respectively. The statistical significance has been tested using the cross-sectional test as proposed by Boehmer et al. (1992) (z-Test) and the Wilcoxon Signed Rank Test as described by Barber and Lyon (1996).*

In order to analyze the impact of acquirer size, the transactions in the sample were sorted by relative deal size (Transaction Volume/ Acquirer's Market Capitalization) and divided into two subsamples containing the 30 largest and 30 smallest transactions. Table 6 presents the findings. We find insignificant positive CAARs to acquirers in the case of relatively large targets ranging from 0.70% to 0.95%. In the case of small transactions from the acquirer's perspective, the CAARs show greater variance ranging from 0.61% to 3.33% and are significant at the 10% level for the [-1;1] and [-5;5] event windows. While the differences are not significant, the results provide an indication of higher CAARs for small acquirers or the acquisition of relatively large targets as suggested by Asquith et al. (1983), Moeller et al. (2003) and Jarrell and Poulsen (1989), they will be challenged in the multivariate analysis.

Table 6: Abnormal Returns to Acquirers Differentiated by Relative Transaction Size

n	Top 30 30		Bottom 30 30		Mean Comparison	
EVENT WINDOW	**CAAR**	**MEDIAN**	**CAAR**	**MEDIAN**	**ΔCAAR**	**ΔMEDIAN**
[-1; +1]	2.17% *	0.30% *	0.95%	0.41%	1.22%	-0.11%
[-5; +5]	3.33% *	2.05%	0.70%	0.61%	2.63%	1.43% **
[-10; +10]	0.61%	0.18%	0.73%	1.17%	-0.12%	-0.98%
[-20; +20]	1.41%	-3.50%	0.80%	0.23%	0.61%	-3.73%

*This table shows the cumulative average abnormal returns (CAAR) to acquiring companies for the top 30 and bottom 30 transactions by relative transaction size in the sample. Statistical significance at the 10%, 5% and 1% level is denoted by *, ** and *** respectively. The statistical significance has been tested using the cross-sectional test as proposed by Boehmer et al. (1992) (z-Test) and the Wilcoxon Signed Rank Test as described by Barber and Lyon (1996).*

Time Period: In section 2 we argued that the change in market concentration and in particular the increased difficulty in finding suitable targets may have an impact on acquirer returns. Table 7 presents our findings comparing transactions between 1998 and 2003 and 2004 and 2010. While transactions announced between 1998 and 2003 yield insignificant positive and negative value effects, transactions between 2004 and 2010 on average yield positive value effects across all event windows. Moreover, with the exception of event window [-10;10] all of the CAARs reported are significant on the 5% or 10% level. A mean comparison, though not statistically significant, reveals higher returns to acquirers between 2004 and 2010 across all event windows. Despite the lack of statistical significance, these results serve as an indication and will be tested in the regression analysis.

Table 7: Abnormal Returns to Acquirers Differentiated by Transaction Date

n	1998-2003 30		2004-2010 39		Mean Comparison	
EVENT WINDOW	**CAAR**	**MEDIAN**	**CAAR**	**MEDIAN**	**ΔCAAR**	**ΔMEDIAN**
[-1; +1]	1.21%	-0.34%	1.66% **	0.46%	-0.45%	-0.80%
[-5; +5]	1.64%	0.66%	1.87% *	1.19% *	-0.23%	-0.53%
[-10; +10]	-0.62%	-0.17%	1.49%	1.29%	-2.12%	-1.45%
[-20; +20]	-0.38%	-3.35%	2.30% *	0.09%	-2.68%	-3.44%

*This table shows the cumulative average abnormal returns (CAAR) to acquiring companies for mergers and acquisitions in the brewing industry between 1998-2003 and 2004-2010. Statistical significance at the 10%, 5% and 1% level is denoted by *, ** and *** respectively. The statistical significance has been tested using the cross-sectional test as proposed by Boehmer et al. (1992) (z-Test) and the Wilcoxon Signed Rank Test as described by Barber and Lyon (1996).*

Public Status of Target: Table 8 presents the results of acquirers' abnormal returns comparing transactions involving public and private targets. The acquisition of public targets yields mixed results with acquirers' CAARs ranging between -1.01% and 1.31% all of which are statistically insignificant. On the other hand, acquirers of private targets experience positive value effects with significant CAARs ranging between 1.58% and 2.49%. Moreover, the CAAR of 2.49% for the [-5;5] event window is statistically significant on the 1% level. While the mean comparison shows that across all analyzed event-windows the acquisition of private targets leads to higher acquirer returns, the differences are not statistically significant. Nonetheless, these results are in line with studies by Moeller et al. (2003) and Faccio et al. (2006).

Table 8: Abnormal Returns to Acquirers Differentiated by Legal Status of Target

	Public		Private		Mean Comparison	
n	29		40			
EVENT WINDOW	CAAR	MEDIAN	CAAR	MEDIAN	ΔCAAR	ΔMEDIAN
[-1; +1]	1.31%	0.06%	1.58% *	0.45%	-0.28%	-0.39%
[-5; +5]	0.77%	-1.56%	2.49% ***	1.61% *	-1.72%	-3.17%
[-10; +10]	-1.01%	0.15%	1.72%	0.94%	-2.73%	-0.79%
[-20; +20]	0.02%	-0.95%	1.94% *	-0.65%	-1.92%	-0.30%

*This table shows the cumulative average abnormal returns (CAAR) to acquiring companies for mergers and acquisitions in the brewing industry for publicly-listed and private target companies. Statistical significance at the 10%, 5% and 1% level is denoted by *, ** and *** respectively. The statistical significance has been tested using the cross-sectional test as proposed by Boehmer et al. (1992) (z-Test) and the Wilcoxon Signed Rank Test as described by Barber and Lyon (1996).*

Type of Consideration: Table 9 compares the results of acquirer's abnormal returns of cash only transactions with transactions that use share-based or hybrid forms of consideration. Overall, brewers show a clear preference for cash only transactions. On average, acquirers paying solely with cash experience a positive CAAR of 1.37% for the [-1;1] event window, which is significant on the 5% level. The limited amount of share deals does not allow for a viable comparison and will hence be addressed in the multivariate analysis.

Table 9: Abnormal Returns to Acquirers Differentiated by Consideration Type

	Cash Only		Share Deals		Mean Comparison	
n	52		11			
EVENT WINDOW	CAAR	MEDIAN	CAAR	MEDIAN	ΔCAAR	ΔMEDIAN
[-1; +1]	1.37% **	0.21%	2.09%	0.54%	-0.72%	-0.33%
[-5; +5]	1.08%	0.88%	4.88%	2.23%	-3.80% *	-1.35%
[-10; +10]	0.32%	0.47%	2.31%	2.56%	-1.99%	-2.09%
[-20; +20]	1.03%	-0.83%	3.71%	-0.95%	-2.68%	0.12%

*This table shows the cumulative average abnormal returns (CAAR) to acquiring companies for mergers and acquisitions in the brewing industry for cash-only and share-based transactions. Statistical significance at the 10%, 5% and 1% level is denoted by *, ** and *** respectively. The statistical significance has been tested using the cross-sectional test as proposed by Boehmer et al. (1992) (z-Test) and the Wilcoxon Signed Rank Test as described by Barber and Lyon (1996).*

Transaction Experience: Tables 10a and 10b present the results of acquirers' abnormal returns based on transaction experience. Overall, 13 of the total of 69 transactions involve acquirers that have only engaged in one transaction in the sample period (Single-Bidder) while 56 transactions involve acquirers that have engaged in at least one transaction in the sample period (Multi-Bidder). 36 transactions involve bidders that have engaged in more than five transactions in the sample period (Bidder-Champion). While we find no significant returns for Multi-Bidder transactions, we find a positive CAAR of 1.25% for Bidder-Champion transactions in the [-5;5] event window, which is significant at the 5% level. Due to the limited amount of Single-Bidder transactions, we shall provide a viable comparison in the regression analysis.

Table 10a: Abnormal Returns to Acquirers Differentiated by Transaction Experience

	Single-bidder			Multi-bidder		Mean Comparison		
n	13			56				
EVENT WINDOW	CAAR		MEDIAN	CAAR	MEDIAN	ΔCAAR		ΔMEDIAN
[-1; +1]	3.38%	**	1.88%	1.02%	0.21%	2.36%		1.67%
[-5; +5]	4.34%	*	2.27%	1.17%	0.88%	3.17%	*	1.39%
[-10; +10]	1.53%		-3.53%	0.35%	1.17%	1.18%		-4.70%
[-20; +20]	2.22%		-1.23%	0.88%	-0.60%	1.35%		-0.63%

*This table shows the cumulative average abnormal returns (CAAR) acquiring companies for mergers and acquisitions in the brewing industry comparing single-bidders i.e. acquirers with only one announced transaction in the sample with multi-bidders i.e. acquirers with more than one announced transactions in the sample. Statistical significance at the 10%, 5% and 1% level is denoted by *, ** and *** respectively. The statistical significance has been tested using the cross-sectional test as proposed by Boehmer et al. (1992) (z-Test) and the Wilcoxon Signed Rank Test as described by Barber and Lyon (1996).*

Table 10b: Abnormal Returns to Acquirers Differentiated by Transaction Experience

	Single-bidder			Bidder Champions			Mean Comparison		
n	13			36					
EVENT WINDOW	CAAR		MEDIAN	CAAR		MEDIAN	ΔCAAR		ΔMEDIAN
[-1; +1]	3.38%	**	1.88%	1.25%	**	0.50%	2.14%		1.38%
[-5; +5]	4.34%	*	2.27%	0.79%		0.23%	3.55%	*	2.04%
[-10; +10]	1.53%		-3.53%	-0.81%		0.04%	2.34%		-3.57%
[-20; +20]	2.22%		-1.23%	-1.63%		-3.89%	3.86%		2.66%

*This table shows the cumulative average abnormal returns (CAAR) acquiring companies for mergers and acquisitions in the brewing industry comparing single-bidders i.e. acquirers with only one announced transaction in the sample with bidder-champions i.e. acquirers with more than five announced transactions in the sample. Statistical significance at the 10%, 5% and 1% level is denoted by *, ** and *** respectively. The statistical significance has been tested using the cross-sectional test as proposed by Boehmer et al. (1992) (z-Test) and the Wilcoxon Signed Rank Test as described by Barber and Lyon (1996).*

Multivariate Analysis

In order to provide a complete picture of the influential factors and to gain further insights into potential dependencies, a cross-sectional regression is performed on the cumulative abnormal returns to acquirers as presented in formula 3. In total, 11 variables are included in the regression model to represent the parameters, which have been individually analyzed in the univariate subsample analysis. In the following, the respective parameter values will be specified in detail.

$$CAR = \alpha_0 + \gamma_1 * Cross\ Border + \gamma_2 * Target\ LA + \gamma_3 * Target\ EE + \gamma_4 * Target\ Asia\ + \gamma_5 * Transaction\ Value + \gamma_6 * Rel.Transaction\ Value + \gamma_7 * Date\ A + \gamma_8 * Share\ Comp + \gamma_9 * Public\ Deal + \gamma_{10} * Multi\ Bidder + \gamma_{11} * Bidder\ Champion$$
(3)

Geographical Scope: The results presented in the univariate analysis provide a first indication that domestic transactions might have a positive impact on short-term acquirer performance when compared to cross-border transactions. At the same time, the results suggested a positive wealth effect if targets were based in emerging market economies (Latin America, Eastern Europe and Asia (ex Japan)). Both effects are included in the regression model using the dummy variables “Cross-Border”, “Latin-America”, “Eastern-Europe” and “Asia”.

Size of Transaction and Acquirer: The results presented in the univariate subsample showed significant positive returns for the bottom 30 transactions by transaction value as well as the top 30 transactions by relative transaction value, indicating a preference for small transactions and small-sized acquirers. In order to verify these results, transaction value and relative transaction value are included as variables in the regression model.

Time Period: As pointed out above the structures of the global beer market have materially changed in recent years. Consequently, we expected these changes to have an impact on acquirer performance. In the univariate subsample analysis we found significant positive returns for transactions between 2004 and 2010, we did not find any significant abnormal performance between 1998 and 2003. In order to test for a supposed relation, we include a dummy variable for time period 1998 – 2003.

Public Status of Target: The univariate results showed highly significant positive returns for the acquisition of private targets. On the other hand, no significant abnormal performance was found for public targets. We test these results by including a dummy variable for public targets in the regression model.

Type of Consideration: The results provided in the univariate analysis showed significant positive returns for cash transactions for the [-1;1] event windows surrounding the announcement date. On the other hand, transactions with share-based consideration showed even greater abnormal returns, albeit being statistically insignificant. In order to test for a supposed relation, we include a dummy variable considering share-based consideration.

Transaction Experience: The subsample analysis showed significant positive abnormal returns for bidder champions i.e. acquirers with at least five announced transactions in the sample. On the other hand, single bidders i.e. acquirers with only one announced transaction in the sample on average experienced even higher abnormal returns. In order to test for a potential relationship between transaction experience and acquirer return, we include two dummy variables reflecting Multi-Bidder and Bidder-Champion transactions.Table 12 presents the results of the complete regression models on the CAARs for the [-1;1], [-5;5] and [-10;10] event windows. The [-5;5] and [-10;10] models are significant on the 5% and 10% level respectively. Explanatory power is remarkably high with adjusted R-squared ranging between 12% and 14%. Autocorrelation issues can be ruled out due to high Durbin-Watson-Statistics in both cases. The overall abnormal short-term performance as represented by the constant yields a positive 4.7% for the [-1;1] and 6.4% for the [-5;5] event windows and are both highly significant at the 1% level.

Overall, these findings correspond to the positive announcement effects determined in the univariate analysis, clearly confirming our expectations.In addition, the regression models complement the univariate subsample analysis enabling the detection of a number of different value drivers of short-term performance. First of all, transaction value and acquirer size are determined to have a significant positive impact on short-term performance. These results stand in contrast to the findings of the univariate subsample analysis. The regression model confirms both target and acquirer size to be positively related to acquirer performance. With regard to target size, this relation can be confirmed for two of the regression models. On the other hand, the negative impact of relative deal size (positive impact of acquirer size) is confirmed across all three regression models. Overall, these findings provide clear evidence on the importance of size and its advantages in the brewing sector.

While not significant for all models, the multivariate analysis confirms the negative impact of cross-border transactions on acquirer returns. With regard to emerging market transactions, regression coefficients across all models are in most cases higher than those for all cross-border transactions. In the case of Latin America, the regression model for the [-1;1] event window even yields a positive regression coefficient of 3.7% which is significant at the 5% level. The subsample analysis on the impact of time on acquirer returns suggested an increase in abnormal acquirer returns over time. The multivariate regression does not show any significant impact of the announcement date on acquirer returns.

With regard to the public status of the target, the subsample analysis suggested acquirer returns to be positively impacted if the acquired target was not publically listed. The regression models confirm these results with negative coefficients across all three event windows. Moreover, a negative relation is determined for the [-5;5] event window which is significant at the 5% level.

Despite the small set of non-cash transactions, the subsample analysis suggested acquirer returns to be positively affected if the transaction includes share-based compensation. The regression models confirm these results with positive coefficients throughout all three models. Nonetheless, only one of the coefficients is significantly positive and hence we do not believe there is enough evidence to confirm a potential dependency. With regard to transaction experience, the univariate results suggested higher announcement returns to Single-Bidders when compared to Multi-Bidders. At the same time, we found significant positive returns to Bidder-Champions. The multivariate analysis does not provide any additional insight as to a potential dependency between acquirer return and transaction experience. Given the lack of additional evidence, we cannot confirm an impact of transaction experience on acquirer performance.

Table 11: Multivariate Regression Analysis

	[-1; +1]			[-5; +5]			[-10; +10]		
Variable	**Coefficient**		**t-value**	**Coefficient**		**t-value**	**Coefficient**		**t-value**
(Constant)	0.047	***	2.740	0.064	***	2.994	0.047		1.623
Cross_Border	-0.014		-0.744	-0.021		-0.955	-0.050	*	-1.657
Targ_LA	0.037	**	1.982	0.025		1.104	0.044		1.408
Targ_EE	-0.019		-1.036	-0.008		-0.378	-0.014		-0.482
Targ_Asia	-0.005		-0.259	-0.010		-0.432	0.033		1.037
Transaction_Value	0.000	*	1.915	0.000	**	2.074	0.000		0.829
Rel_Value	-0.018	*	-1.660	-0.029	**	-2.153	-0.040	**	-2.174
Date_A	0.002		0.127	-0.003		-0.178	-0.012		-0.602
Share_Cash	0.015		0.873	0.059	***	2.863	0.044		1.609
Public_Target	-0.016		-1.147	-0.036	**	-2.167	-0.036		-1.603
Multi_Bidder	-0.029		-1.592	-0.022		-0.985	0.021		0.697
Bidder_Champion	0.009		0.550	-0.006		-0.278	-0.015		-0.529
R-squared	0.221			0.278			0.267		
Adjusted R-squared	0.071			0.139			0.126		
Durbin-Watson	1.990			2.379			2.530		
F-statistic	1.469			1.997	**		1.890	*	
p (F-stat)	0.169			0.046			0.060		

*CAARs were derived for a sample of 69 transactions in the brewing industry between 1998 and 2010. For a detailed description of the variables and the underlying equation, see section 4.1. The Durbin-Watson statistics were estimated to test for autocorrelation of the residuals. Proximity of the value to "2" is regarded as an indication of no autocorrelation between residuals. Statistical significance at the 10%, 5% and 1% level is denoted by *, ** and *** respectively.*

Announcement and Rival Effects among the "Big Four"

While the analyses presented above covered all publicly listed acquirers, in the following we give particular attention to announcement effects of the "big four" and rival effects among them. Table 12 presents the short-term announcement effects of the "big four". Overall, we find positive announcement returns of 1.25% for the [-1;1] event window. For the t and z statistics, these results are significant on the 5% level. For the non-parametric Wilcoxon test, we do not find any significant return, though it should be noted that the p-value of 0.11 is very close to the threshold for significance at the 10% level. These results are in line with our predictions and support the argument that capital markets positively value the announcement of a M&A transaction in a strongly consolidated market, where suitable targets are becoming increasingly hard to find. However, a closer look at the other event windows suggests that the positive CAARs decrease the bigger the event window is defined eventually turning negative for the [-10;10] and [-20;20] event windows. While these results are not statistically significant, they suggest an interesting trend. Nonetheless, given the significant positive returns for the [-1;1] event window we conclude that the "big four" experience positive short-term value effects following the announcement of a transaction.

Table 12: Cumulative Average Abnormal Returns to the "Big Four"

Acquirers (N=36)										
Event-Window	**CAAR**	**t-Test**			**z-Test**			**WCX Test**		
		t-value	**p-value**		**z-value**	**p-value**		**z-value**	**p-value**	
[-1; +1]	1.25%	2.04	0.04	**	1.99	0.05	**	-1.59	0.11	
[-5; +5]	0.79%	0.96	0.34		1.12	0.27		-0.77	0.44	
[-10; +10]	-0.81%	-0.72	0.48		-0.55	0.58		-0.55	0.58	
[-20; +20]	-1.63%	-1.05	0.30		-0.85	0.40		-1.54	0.12	

*This table shows the cumulative average abnormal returns (CAAR) to the "big four" following the announcement of mergers and acquisitions in the brewing industry. It contains all transactions by the "big four" for which trading data was available between 250 before and 20 days after transaction announcement. Statistical significance at the 10%, 5% and 1% level is denoted by *, ** and *** respectively. The statistical significance has been tested using a standard t-test, the cross-sectional test as proposed by Boehmer et al. (1992) (z-Test) and the Wilcoxon Signed Rank Test as described by Barber and Lyon (1996).*

Table 13 presents the rival returns to the "big four" i.e. the returns to the remaining 3 companies if one of the "big four" announces a transaction (e.g. the returns to Heineken, Carlsberg and SABMiller if Anheuser Busch Inbev announces a transaction). Using this approach we analyze 98 rival events for our sample. While we recognize a similar pattern of decreasing abnormal returns in case of larger event windows, we only find significantly negative returns for the [-10;10] event window. The returns are significant at the 10% level using a standard t-statistic and the non-parametric Wilcoxon signed rank test. The findings suggest that the negative competitive effects of missing out on a potential M&A opportunity or strengthened competition from a newly combined firm, outweigh potential positive signaling effects. These results stand in contrast to existing literature (see e.g. Eckbo (1985), Song and Walking (2000), Clougherty and Duso (2009)) and confirm the exceptional characteristics of the brewing sector.

Table 13: Cumulative Average Abnormal Returns to Rivals of the "Big Four"

Rivals (N=98)										
Event-Window	**CAAR**	**t-Test**			**z-Test**			**WCX Test**		
		t-value	**p-value**		**z-value**	**p-value**		**z-value**	**p-value**	
[-1; +1]	0.33%	1.02	0.31		1.09	0.28		-0.86	0.39	
[-5; +5]	0.25%	0.52	0.60		1.04	0.30		-0.26	0.79	
[-10; +10]	-1.32%	-1.92	0.06	*	-1.26	0.21		-1.69	0.09	*
[-20; +20]	-1.53%	-1.49	0.14		-0.64	0.52		-0.91	0.36	

*This table shows the cumulative average abnormal returns (CAAR) to the "big four" following the announcement of mergers and acquisitions by a rival. It contains all events for which trading data was available between 250 before and 20 days after transaction announcement. Statistical significance at the 10%, 5% and 1% level is denoted by *, ** and *** respectively. The statistical significance has been tested using a standard t-test, the cross-sectional test as proposed by Boehmer et al. (1992) (z-Test) and the Wilcoxon Signed Rank Test as described by Barber and Lyon (1996).*

Overall, our findings seem very consistent with previous industry research (see e.g. Kerkvliet et al. (1998), Earlam et al. (2010), Schwankl (2008)) and confirm the importance of scale and synergies in the brewing sector. In a consolidated market, where it is becoming increasingly hard to find suitable M&A targets, capital markets seem to value the successful quest for a consolidation opportunity, while punishing rivals that miss out. Even though our study focuses solely on the brewing industry, the documented results may be indicative for other industries as well. In particular, industries with other consumer products and similar oligopolistic market structures (e.g. tobacco and breakfast cereals) may yield similar results.

CONCLUSION

The objective of this study was to analyze M&A announcement effects of acquirers and rivals in the brewing industry. For this purpose, a sample of 69 horizontal M&A transactions involving brewing companies between 1998 and 2010 was identified and examined using a combination of two approaches: the traditional event study methodology and the Fama-French-3-Factor model. Our results provide new insights into the perceived short-term success of M&A transactions in the brewing industry and its corresponding evaluation through capital markets.

Firstly, our results indicate that acquirers in the brewing industry experience significant positive short-term value effects following the announcement of an M&A transaction. This positive finding is an outstanding attribute of the sector and stands in contrast to cross-industry studies by Andrade et al. (2001), Bruner (2002) and Loughran and Vijh (1997) and older industry specific research by Ebneth and Theuvsen (2007) all of which provide evidence of significant negative abnormal returns or at most insignificant positive returns. Therefore, it appears that capital markets specifically in recent years value the above-average synergy potential of the sector.

Secondly, our results provide evidence for a number of value drivers of abnormal acquirer performance. Consistent with the findings of Datta and Puia (1995) we find a positive relation between domestic transactions and short-term acquirer performance. However, at the same time and in contrast to the findings of Williams and Liao (2008), we also find a positive impact of cross-border transactions involving targets in emerging markets, in particular Latin America, suggesting that capital markets value the international diversification strategies deployed by brewers. Moreover, in contrast to findings of Asquith et al. (1983), Jarrell and Poulsen (1989) and Moeller et al. (2003) we report a positive impact of acquirer size and transaction value, emphasizing the size advantages in the sector. On the other hand, our results provide evidence with regard to a negative relationship between acquirer returns and the target's public status. These results are consistent with Moeller et al. (2003) and Faccio et al. (2006), suggesting that brewers may have to pay significant premiums for targets that are publicly listed.

Thirdly, our results indicate that the "big four" experience significant positive value effects upon announcing an M&A transaction suggesting that capital markets value the successful search for a suitable target in a strongly concentrated market. Close rivals, missing out on an M&A opportunity suffer from significant short-term value losses.While our study addresses a number of important questions with regard to capital market effects of M&A in the brewing industry and the impact of determinant variables, the presented results also leave open questions and give rise to new research issues. As our study is limited to short-term capital market effects, the question arises whether acquiring brewers are able to sustain the positive announcement effects. Thus, future research could investigate the long-term implications of M&A in the sector. In addition to investigating capital market implications, future research could also analyze the impact of M&A on the operating performance of the involved companies. Given the expected continuation of the consolidation process, we believe that M&A in the brewing industry provide an interesting avenue for future research.

Furthermore, the results underline the importance of industry-specific M&A analyses, as sector related potentials to generate value differ among industries, resulting in possibly biased finding of cross-industry examinations.

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BIOGRAPHY

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WHEN DO COSTA RICA NATIONAL BANKS RESPOND TO RESERVE REQUIREMENT CHANGES?

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ABSTRACT

The process of changing reserve requirements in Costa Rica is a three step process. First the central bank makes the decision to change reserve requirements. Several days to several weeks later, the change is announced in the official newspaper. The actual reserve requirement change takes place from several weeks to several months later. Previous studies have limited their analysis to an examination of the decision and the announcement dates. The research shows that Costa Rica national banks do not respond to reserve requirement change announcements or reserve requirement change decisions. In this paper we examine the extent to which Costa Rica national banks respond to reserve requirement changes on the effective day of the reserve requirement change. We find evidence that Costa Rica national banks change their interest rate spreads on the effective day.

JEL: E42, E58

KEYWORDS: Reserve Requirements, Banking, Costa Rica, Interest Rates

INTRODUCTION

Reserve requirement changes have been extensively examined in U.S. markets. Far less is known about responses to reserve requirement changes in other countries. This paper examines Costa Rica national bank responses to reserve requirement changes. Costa Rica provides a unique setting for examining responses to reserve requirement changes for four reasons. First while geographically small, Costa Rica is home to three of the ten largest Central American banks. Costa Rica has about 28 percent of Central American Bank Assets (Ballestero and Martinez, 2007). Second, Costa Rica banks accept deposits and make loans both in U.S. dollars and colon, the local currency. Banks hold reserves against deposits in both currencies, but reserve requirements have sometimes been different for deposits in the two currencies. Moreover, reserve requirement changes do not always coincide. Third, the Costa Rica Central Bank made ten reserve requirement change announcements between 1996 and 2010. These announcements made 18 changes of colon denominated deposit reserve requirements and ten changes of dollar denominated reserve requirements. By comparison, the most recent reserve requirement change in the United States occurred in 1992. Finally, Costa Rica has both government sponsored banks and privately owned banks providing a rich environment to study responses of various institution classes to regulation changes.

The process of implementing a reserve requirement change in Costa Rica starts with the decision by the Central Bank to change rates. Several days (sometimes several weeks) later the decision becomes public. The public announcement appears in the official Costa Rica Newspaper, *La Gaceta*. A lag of several days to several months occurs between the announcement date and the effective date of the change. This sequence of events is depicted in Figure 1.

The motivation for this study emanates from earlier work which finds that Costa Rica national banks do not respond to reserve requirement change announcements (Stewart, Jalbert and Jalbert, 2008, 2010). One plausible explanation for the lack of response is that reserve requirement change information is leaked to national banks prior to the announcement. Earlier studies have found no information leakage effects (Stewart, Jalbert and Jalbert, 2010). So the combined evidence shows no response to reserve requirement changes on the decision or announcement dates. Thus an open question remains regarding if, when and how Costa Rica national banks respond to reserve requirement changes.

Figure 1: Costa Rica Reserve Requirement Change Process

```
|-------------------------------------------|------------------------------------------------|
Decision                        Announcement                              Effective
  Date                              Date                                     Date
```

This figure shows the sequence of events that occur for a reserve requirement change in Costa Rica. The decision date is the day that the Costa Rica bank makes a decision to change reserve requirements. The announcement date is the day that the Central bank decision is made public. The effective date is the day that the reserve requirement change becomes policy.

In this study we examine the response of Costa Rica national banks to reserve requirement changes on the effective date of the reserve requirement change. Evidence to suggest that banks respond predictably on the effective date suggests a predictable opportunity for market participants who observe the announcement. Market participants that observe a reserve requirement increase would want to borrow in advance of a corresponding interest rate increase. Market participants that observe a reserve requirement decrease would want to lend before the impending interest rate decline. Such a pattern might also explain the findings of Jalbert, Stewart and Jalbert (2006), that uncovered interest rate arbitrage opportunities are present in the Costa Rica deposit account market.

The remainder of this paper is organized as follows. In the second section, we discuss the relevant literature. A discussion of the Costa Rica banking system follows. Next, we discuss the data and methodology used in the paper. In the ensuing section the test results are presented. The paper closes with some concluding comments and suggestions for future research.

LITERATURE REVIEW AND BACKGROUND

Two basic theories relating to the role of reserve requirements appear in the literature. The first argues that reserve requirements are a tool of monetary policy. This theory suggests that central banks reduce reserve requirements in order to free bank reserves. By doing so, they increase the money supply and provide economic stimulus. Hein and Stewart (2002), and Stewart and Hein (2002), find support for the contention that reserve requirement changes have not been recently used as a monetary tool in the United States. They find the Federal Reserve commonly offsets reserve requirement changes with open market operations. The increase in money supply caused by a reduction in reserve requirements is at least partially offset by Federal Reserve sales of Treasury securities.

A second theory posits that reserve requirements are a tax on financial institutions. If this theory holds Central Banks might use reserve requirement changes during economic downturns to bolster the health and stability of financial institutions. A number of authors have tested hypotheses related to this theory and generally find evidence supporting this supposition (Black, 1975, Fabozzi and Thurston, 1986, Fama, 1985, and James, 1987).

The tax theory of reserve requirements is well accepted. However, less clear is who ultimately bears the tax. One line of literature suggests that financial institutions pass the tax onto their customers. However, there is disagreement regarding which customers, or combination of customers, ultimately bear the tax. Black (1975) and Fabozzi and Thurston (1986)) contend that bank depositors shoulder the tax in the form of lower deposit rates on reservable time deposits. Fama (1985) and James (1987) argue that borrowers shoulder the tax when they pay higher borrowing rates. A third group of researchers (Cosimano and

McDonald (1998) and Hein and Stewart (2002)) argue that financial institutions bear the tax in the form of lower profitability. They argue that competition prevents financial institutions from passing the tax on to either depositors or borrowers. Several authors find that U.S. bank stock prices move inversely to announced reserve requirement changes supporting this contention (Kolari, Mahajan, and Saunders, 1998, Slovin, Sushka and Bendeck, 1990, and Osborne and Zaher, 1992). Sellon and Weiner (1996) and (1997) provide an extended discussion of the history and goals of reserve requirements and reserve requirement changes.

While substantial research is available regarding the role of reserve requirements and reserve requirement changes in the U.S., much less evidence is available on the role of reserve requirements in other countries. Most notable is a general lack of evidence from lesser developed countries. Reinhart and Reinhart (1999) suggest that Costa Rica and other countries use reserve requirement changes to deal with capital flow problems. They contend that reserve requirement changes mitigate the impact of foreign exchange market interventions and offset the effects of large capital flows. They conjecture the precise implications of reserve requirement changes depend on whether borrowers or depositors pay the tax implied by reserve requirement increases. They examine ten countries, including Costa Rica, finding evidence that both deposit and lending rates respond to reserve requirement changes.

Stewart, Jalbert and Jalbert (2008) examine the tax hypothesis of reserve requirements in Costa Rica. Their data covers a ten year period from 1996 through 2006 including 1-, 3-, and 6-month deposit rates. Data are collected for government banks, private banks and non-bank financial institutions. Loan rates are also examined for several loan categories. They find mixed evidence regarding who bears the brunt of the tax. Evidence from earlier reserve requirement changes support the theories that the tax is passed on to depositors or to borrowers. More recent evidence supports the theory that financial institutions and their shareholders bear the tax in the form of lower profits.

The standard approach for testing the bank profitability hypothesis is to examine stock price changes around reserve requirement chance. Unfortunately in Costa Rica, reliable and accurate stock price data is not readily available. In a recent paper, Stewart, Jalbert and Jalbert (2010) develop a methodology to test the bank profitability theory in the absence of stock price data. Their approach involves using the spread between loan and deposit rates as a proxy for bank profitability. They find that private banks and non-bank financial institutions change their spreads in predictable ways around reserve requirement announcements. However, their results show that government banks do not change their interest rate spreads around reserve requirement change announcements. Moreover, they find no evidence of information leakage between the decision date and the announcement date. This finding leaves open the question of how and when Costa Rica government banks respond to reserve requirement changes. The goal of the current research is to close this gap in the literature. In this paper, we examine effective dates of reserve requirement changes. Specifically we answer the question: Do government banks change their interest rate spreads around reserve requirement changes?

As noted above, government banks respond differently to reserve requirement change announcements. To fully understand this behavior, it is necessary to identify how government banks differ from their private and non-bank financial institution competitors. Clearly privately owned banks are driven by the goal of profit maximization. Government banks however, may have a different role. Stewart, Jalbert and Jalbert (2010), conduct a survey of the mission statements of five Costa Rica Government banks. They find that each of the government banks reference some social objective in their mission or vision statements. Generally, these statements refer to improving the welfare of the Costa Rica people, or improving the economic development of the country. It is clear Costa Rica government bank objectives differ from those of private institutions.

THE COSTA RICA BANKING SYSTEM

The Costa Rica Central Bank frequently changed reserve requirements in the mid 1990's and early 2000's. Moreover sometimes reserve percentages differed for deposits denominated in Costa Rican colon and in U.S. dollars. In the mid 1990's, the Costa Rica banking system applied as many as five different required reserve rates depending on currency and maturity of the deposit. Colon denominated deposits were classified into the following groups: money market securities, time deposits of 180 days or less and time deposits greater than 180 days in maturity. U.S. dollar denominated deposits were classified as money market securities or time deposits. Table 1 shows reserve requirement rates for colon denominated securities with less than 180 days to maturity and dollar denominated time deposits. These rates are as presented in Stewart, Jalbert and Jalbert (2008 and 2010). In some instances the data reported here do not indicate a rate change, even though an effective date is presented. In these instances, reserve requirements were changed, but not for the type of deposits presented in the table.

Table 1 shows frequent changes in reserve requirements. Reserve requirements decreased from 36% and 43% on dollar and colon denominated deposits respectively in 1996, to 15% for both classes of deposits currently. Most recently changes have involved increases in reserve requirement rates. The central bank frequently announces several reserve requirement changes at one time that are phased in over several months. We take the announcement date to be the date that the reserve requirement change was published in La Gaceta, the official newspaper of the Costa Rican Government. For example, on January 22, 1997, a series of changes were announced that were phased in on eleven separate dates over a fourteen month time period. These changes resulted in a single required rate across all deposit classes beginning February 1, 2002. On June 30, 2005, the most recent reserve requirement change was announced.

DATA AND METHODOLOGY

This study uses the same data as Stewart, Jalbert and Jalbert (2010). As noted in Stewart, Jalbert and Jalbert, 2008 and 2010, stock data would be ideal for analyzing the effect of reserve requirement changes on financial institution profitability. Stock price data were obtained from Superintendencia General de Valores (SUGEVAL). SUGEVAL is the Costa Rica counterpart to the U.S. Securities and Exchange Commission. However, due to reporting problems and a lack of liquidity the data were not useable.

Liquidity issues are evident in the data as the average stock in Costa Rica trades on only 13.5 percent of all trading days (Stewart, Jalbert and Jalbert, 2008). The data also shows that stock prices sometimes remained unchanged for long periods of time. As noted in Stewart Jalbert and Jalbert (2010) these findings may be because the data was substantially incomplete. Supporting this contention is the observation that on some trading days, no individual stock transactions were reported, but there was a reported change in the Costa Rica stock index level. The incomplete nature of the data could be a result of certain regulatory and confidentiality issues that prohibit SUGEVAL from disclosing a complete transaction list.

In an attempt to obtain additional data, the authors contacted the Costa Rica stock exchange, Bolsa Nacional de Valores. Despite a personal visit to the exchange and several email follow ups requesting data, the authors were unable to obtain additional data. Collectively these outcomes suggest that tests based on stock price reactions to reserve requirement changes are not currently feasible.

To overcome these data limitations, this paper uses a proxy measure of financial institution profitability. This proxy permits the evaluation of theories based on bank profitability. We measure bank profitability as the spread between loan rates and deposit rates. This spread represents the gross profit on lending operations, and is believed to be a reasonable measure for overall bank profitability.

Table 1: Costa Rica Reserve Requirement Rates

Meeting Date	Date Published	Effective Date	Colon Denominated Time Deposits ≤ 180 Days	Dollar Denominated Time Deposits
1-31-1996	2-16-1996	Pre-3-1-1996*	30%	17%
1-31-1996	2-16-1996	3-1-1996	30%	17%
1-31-1996	2-16-1996	4-1-1996	30%	17%
1-31-1996	2-16-1996	5-1-1996	30%	17%
1-31-1996	2-16-1996	6-1-1996	30%*	17%
1-31-1996	2-16-1996	7-1-1996	28%*	17%
1-31-1996	2-16-1996	8-1-1996	26%*	17%
1-31-1996	2-16-1996	9-1-1996	24%*	17%
1-31-1996	2-16-1996	10-1-1996	22%*	17%
1-31-1996	2-16-1996	11-1-1996	20%*	17%
1-31-1996	2-16-1996	11-28-1996	17%*	17%
1-31-1996	2-16-1996	12-31-1996	17%	17%
1-22-1997	2-22-1997	3-1-1997	17%	5%*
1-22-1997	2-22-1997	4-1-1997	17%	5%
1-22-1997	2-22-1997	5-1-1997	17%	5%
1-22-1997	2-22-1997	6- 1-1997	17%	5%
1-22-1997	2-22-1997	7-1-1997	16%*	5%
1-22-1997	2-22-1997	8-1-1997	16%	5%
1-22-1997	2-22-1997	10-1-1997	16%	5%
1-22-1997	2-22-1997	12-1-1997	18%*	5%
1-22-1997	2-22-1997	1-1-1998	17%*	5%
1-22-1997	2-22-1997	2-1-1998	16%*	5%
1-22-1997	2-22-1997	3-1-1998	15%*	5%
4-1-1998	5-5-1998	5-1-1998	15%	5%
9-16-1999	09-27-1999	10-15-1999	14%*	5%
2-9-2000	2-24-2000	3-1-2000	12%*	5%
2-9-2000	2-24-2000	4-1-2001	11%*	5%
2-9-2000	2-24-2000	10-1-2001	10%*	5%
12-6-2000	12-15-2000	1-1-2001	11%*	5%
12-6-2000	12-15-2000	10-1-2001	10%*	5%
3-28-2001	4-10-2001	5-1-2001	9%*	5%
3-28-2001	4-10-2001	9-1-2001	7%*	5%
3-28-2001	4-10-2001	2-1-2002	5%*	5%
12-18-2002	1-23-2003	1-16-2003	6.5%*	6.5%*
12-18-2002	1-23-2003	2-16-2003	8%*	8%*
12-18-2002	1-23-2003	3-16-2003	10%*	10%*
7-26-2004	8-3-2004	9-1-2004	11%*	11%*
7-26-2004	8-3-2004	10-1- 2004	12%*	12%*
6-15-2005	6-30-2005	7-16-2005	13.5%*	13.5%*
6-15-2005	6-30-2005	8-16-2005	15%*	15%*

*This table shows the required reserve rates in Costa Rica banks for colon and dollar denominated accounts. Date Published is the date the reserve requirement change was announced in La Gaceta. Effective Date is the date the reserve requirement change took effect. * indicates a reserve requirement change that affected a given deposit class.*

Interest rate data were collected from the Costa Rica Central Bank on loans and deposits. The data covers the time period of February 21, 1996 through January 22, 2008 including 4,354 daily observations. For a few series data was not reported on October 31, 2007 resulting in 4,353 observations. Data were collected for Government Bank deposit and loan rates for both colon and dollar denominated accounts. Loan data is categorized as agricultural, ranching, construction, industrial, real estate and other. Daily deposit rate data for 1-month, 3-month, and 6-month time deposits were collected.

Data on loan maturity were not available from the Central Bank. The Central Bank apparently aggregates loan rate data across maturities. Because of this limitation maturity matching of deposit and loan rates was not possible. To address this issue, we compare each loan rate to several deposit rates with different maturities. The interest rate spread is computed as follows:

$$\text{Spread} = \text{Loan Rate} - \text{Deposit Rate} \qquad (1)$$

This process resulted in eighteen dollar denominated interest rate series and eighteen colon denominated interest rate series for examination. The interested reader is referred to Stewart, Jalbert and Jalbert (2008 and 2010) for summary statistics of the data.

The market reaction to changes in reserve requirements is analyzed by comparing the spread levels in the five days before the reserve requirement change, the pre-announcement period, and the five days after the reserve requirement change, the post-announcement period. The announcement day is included in the post-announcement period. We graphically depict this time frame in Figure 2. We test for differences in the spread before and after each reserve requirement change using the Mann-Whitney test.

Figure 2: Examination Period

```
                                          Change
Day  -6    -5    -4    -3    -2    -1     Day    1     2     3     4     5     6
     |-----|-----|-----|-----|-----|------|------|-----|-----|-----|-----|-----|
           ******************************        ******************************
                 Pre Change Period                     Post Change Period
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This figure shows the time frame examined for each interest rate spread.

RESULTS

Table 2 presents results of the analysis on interest rate spreads for colon denominated securities. The results are presented for 1-month, 3-month and 6-month deposit rates for agriculture, cattle ranch, industrial, construction, real estate and other loans. The analysis outcomes are reported for each of the twenty-six reserve changes that occurred between July 1, 1996 and August 16, 2005.

The results are mixed. For fourteen reserve requirement changes, no evidence of a spread change was evident. For five reserve requirement changes, there was evidence of a spread change, but the change was not significant. For seven reserve requirement changes, there was evidence of a significant change in interest rate spreads. For most reserve requirement changes with a significant result, the results are consistent for most loan types deposit maturity.

A closer examination of the spreads shows that three significant spread changes were associated with the January 31, 1996 announcement. The remaining significant spreads were each associated with a different announcement. The analysis shows that three significant spread changes involved a spread decrease and four involved a spread increase. Thus, there is no discernable pattern to draw additional inferences from.

The careful reader will recall in some instances one announcement served notice for several future reserve requirement changes. For example, the January 31, 1996 announcement provided notification of twelve future reserve requirement changes. Of interest is to determine if spread changes occur for a consistent effective date. That is, does the spread change for the first effective date after an announcement, but not for later effective dates? Such a pattern would explain the absence of a statistically significant change in rates following some of the multiple date announcements. The results show that three significant spread changes occurred on the first effective date after an announcement. Four significant changes occurred for the second or later effective date associated with an announcement. Thus, no obvious pattern is present to draw additional inferences.

Table 2: Mann-Whitney Test: Colon Denominated Interest Rate Spreads

		Effective Date								
		July 1, 1996 (12)			**August 1, 1996 (34)**			**September 1, 1996(56)**		
		Pre	Post	M-W	Pre	Post	M-W	Pre	Post	M-W
Agriculture	1 Month	12.0317	12.0317	0.000	12.112	12.295	-2.327**	11.972	11.972	0.000
Cattle Ranch	1 Month	11.999	11.999	0.000	12.078	12.263	-2.327**	12.464	12.464	0.000
Industrial	1 Month	11.831	11.831	0.000	11.911	12.094	-2.327**	11.907	11.907	0.000
Construction	1 Month	13.235	13.235	0.000	13.315	13.498	-2.327**	13.385	13.385	0.000
Real Estate	1 Month	9.110	9.110	0.000	9.190	9.374	-2.327**	9.361	9.631	0.000
Other	1 Month	14.480	14.480	0.000	14.560	14.744	-2.327**	14.537	14.537	0.000
Agriculture	3 Month	11.835	11.835	0.000	11.915	12.099	-2.327**	11.751	11.751	0.000
Cattle Ranch	3 Month	11.803	11.803	0.000	11.883	12.067	-2.327**	12.243	12.243	0.000
Industrial	3 Month	8.914	8.914	0.000	8.994	9.178	-2.327**	9.140	9.140	0.000
Construction	3 Month	11.634	11.634	0.000	11.714	11.898	-2.327**	11.686	11.686	0.000
Real Estate	3 Month	13.038	13.038	0.000	13.118	13.302	-2.327**	13.164	13.164	0.000
Other	3 Month	14.284	14.284	0.000	14.364	14.547	-2.327**	14.316	14.316	0.000
Agriculture	6 Month	10.652	10.652	0.000	10.732	10.915	-2.327**	10.592	10.592	0.000
Cattle Ranch	6 Month	10.619	10.619	0.000	10.699	10.883	-2.327**	11.084	11.084	0.000
Industrial	6 Month	10.451	10.451	0.000	10.531	10.714	-2.327**	10.527	10.527	0.000
Construction	6 Month	11.855	11.855	0.000	11.935	12.118	-2.327**	12.005	12.005	0.000
Real Estate	6 Month	7.730	7.730	0.000	7.810	7.994	-2.327**	7.981	7.981	0.000
Other	6 Month	13.100	13.100	0.000	13.180	13.364	-2.327**	13.157	13.157	0.000
		October 1, 1996(78)			**November 1, 1996(9,10)**			**November 28, 1996 (11,12)**		
Agriculture	1 Month	11.526	11.529	-2.327**	11.463	11.474	-1.833*	11.543	11.543	0.000
Cattle Ranch	1 Month	12.663	12.660	2.327**	12.675	12.675	0.000	12.882	12.882	0.000
Industrial	1 Month	11.835	11.839	-2.327**	11.881	11.881	0.000	12.026	12.026	0.000
Construction	1 Month	13.313	13.310	2.327**	13.287	13.287	0.000	12.708	12.708	0.000
Real Estate	1 Month	9.289	9.289	0.000	9.250	9.250	0.000	9.488	9.488	0.000
Other	1 Month	14.465	14.468	-2.327**	14.483	14.483	0.000	14.504	14.504	0.000
Agriculture	3 Month	11.307	11.307	-2.327**	11.241	11.252	-1.833*	11.321	11.321	0.000
Cattle Ranch	3 Month	12.441	12.438	2.327**	12.453	12.453	0.000	12.670	12.670	0.000
Industrial	3 Month	9.0671	9.0672	-2.327**	9.027	9.027	0.000	9.266	9.266	0.000
Construction	3 Month	11.613	11.616	-2.327**	11.659	11.659	0.000	11.804	11.804	0.000
Real Estate	3 Month	13.091	13.088	2.327**	13.066	13.066	0.000	12.486	12.486	0.000
Other	3 Month	14.243	14.246	-2.327**	14.261	14.261	0.000	14.282	14.282	0.000
Agriculture	6 Month	10.146	10.149	-2.327**	10.083	10.094	-1.833*	9.925	9.925	0.000
Cattle Ranch	6 Month	11.283	11.280	2.327**	11.295	11.295	0.000	11.264	11.264	0.000
Industrial	6 Month	10.455	10.459	-2.327**	10.501	10.501	0.000	10.408	10.408	0.000
Construction	6 Month	11.933	11.930	-2.327**	11.907	11.907	0.000	11.090	11.090	0.000
Real Estate	6 Month	7.909	7.909	-2.327**	7.870	7.870	0.000	7.870	7.870	0.000
Other	6 Month	13.085	13.088	-2.327**	13.103	13.103	0.000	12.886	12.886	0.000
		July 1, 1997 (13,14)			**December 1, 1997 (15,16)**			**January 1, 1998 (17,18)**		
Agriculture	1 Month	10.577	10.577	0.000	10.907	10.907	0.000	10.907	10.907	0.000
Cattle Ranch	1 Month	10.694	10.694	0.000	11.317	11.317	0.000	11.317	11.317	0.000
Industrial	1 Month	10.356	10.356	0.000	11.090	11.090	0.000	11.090	11.090	0.000
Construction	1 Month	10.909	10.909	0.000	11.524	11.524	0.000	11.524	11.524	0.000
Real Estate	1 Month	7.957	7.957	0.000	8.138	8.138	0.000	8.138	8.138	0.000
Other	1 Month	13.655	13.655	0.000	12.952	12.952	0.000	12.952	12.952	0.000
Agriculture	3 Month	9.799	9.799	0.000	10.289	10.289	0.000	10.289	10.289	0.000
Cattle Ranch	3 Month	9.195	9.195	0.000	10.700	10.700	0.000	10.698	10.698	0.000
Industrial	3 Month	7.178	7.178	0.000	7.251	7.251	0.000	7.251	7.251	0.000
Construction	3 Month	9.578	9.578	0.000	10.473	10.473	0.000	10.473	10.473	0.000
Real Estate	3 Month	10.131	10.131	0.000	10.906	10.906	0.000	10.906	10.906	0.000
Other	3 Month	12.877	12.877	0.000	12.335	12.335	0.000	12.335	12.335	0.000
Agriculture	6 Month	8.414	8.414	0.000	9.126	9.126	0.000	9.127	9.127	0.000
Cattle Ranch	6 Month	8.531	8.531	0.000	9.536	9.536	0.000	9.536	9.536	0.000
Industrial	6 Month	8.193	8.193	0.000	9.309	9.309	0.000	9.309	9.309	0.000
Construction	6 Month	8.746	8.746	0.000	9.743	7.743	0.000	9.743	9.743	0.000
Real Estate	6 Month	5.794	5.794	0.000	6.357	6.357	0.000	6.357	6.357	0.000
Other	6 Month	11.492	11.492	0.000	11.171	11.171	0.000	11.171	11.171	0.000

Table 2: (Continued)

					Effective Date					
		February 1, 1998 (19, 20)			March 1, 1998 (21,22)			October 15, 1999 (23,24		
		Pre	**Post**	**M-W**	**Pre**	**Post**	**M-W**	**Pre**	**Post**	**M-W**
Agriculture	1 Month	10.907	10.898	1.350	10.623	10.623	0.000	12.237	12.237	0.000
Cattle Ranch	1 Month	11.317	11.410	-1.350	11.399	11.399	0.000	12.427	12.427	0.000
Industrial	1 Month	11.090	11.192	-1.350	11.105	11.105	0.000	12.573	12.573	0.000
Construction	1 Month	11.523	11.616	-1.350	11.541	11.541	0.000	12.667	12.667	0.000
Real Estate	1 Month	8.138	8.230	-1.350	8.363	8.363	0.000	11.740	11.740	0.000
Other	1 Month	12.952	13.045	-1.350	12.814	12.814	0.000	14.368	14.368	0.000
Agriculture	3 Month	10.289	10.281	1.350	10.169	10.169	0.000	11.448	11.448	0.000
Cattle Ranch	3 Month	10.700	10.793	-1.350	10.945	10.945	0.000	11.638	11.638	0.000
Industrial	3 Month	7.521	7.613	-1.350	7.909	7.909	0.000	10.951	10.951	0.000
Construction	3 Month	10.473	10.566	-1.350	10.651	10.651	0.000	11.785	11.785	0.000
Real Estate	3 Month	10.906	10.999	-1.350	11.089	11.087	0.000	11.878	11.878	0.000
Other	3 Month	12.335	12.427	-1.350	12.360	12.360	0.000	13.579	13.579	0.000
Agriculture	6 Month	9.126	9.024	1.350	8.738	8.738	0.000	10.802	10.802	0.000
Cattle Ranch	6 Month	9.536	9.536	0.000	9.514	9.514	0.000	10.993	10.993	0.000
Industrial	6 Month	9.309	9.309	0.000	9.220	9.220	0.000	11.139	11.139	0.000
Construction	6 Month	9.743	7.743	0.000	9.656	9.656	0.000	11.233	11.233	0.000
Real Estate	6 Month	6.357	6.357	0.000	6.478	6.478	0.000	10.305	10.305	0.000
Other	6 Month	11.171	11.171	0.000	10.929	10.929	0.000	12.934	12.934	0.000
		March 1, 2000 (25,26)			**January 1, 2001 (31,32)**			**April 1, 2001 (27,28)**		
Agriculture	1 Month	12.354	12.354	0.000	12.150	11.810	1.833*	11.753	11.807	-1.35
Cattle Ranch	1 Month	12.419	12.419	0.000	12.208	11.879	1.833*	11.851	11.904	-1.350
Industrial	1 Month	12.127	12.127	0.000	12.132	11.724	1.833*	11.508	11.562	-1.350
Construction	1 Month	12.528	12.528	0.000	12.725	12.322	1.833*	12.120	12.174	-1.350
Real Estate	1 Month	11.043	11.043	0.000	11.144	11.200	-1.833*	11.736	11.790	-1.350
Other	1 Month	14.398	14.398	0.000	14.144	13.866	1.833*	13.956	14.001	-1.350
Agriculture	3 Month	11.989	11.989	0.000	11.690	11.378	1.830*	11.339	11.394	-1.350
Cattle Ranch	3 Month	12.055	12.055	0.000	11.748	11.448	1.830*	11.438	11.438	-1.350
Industrial	3 Month	10.679	10.679	0.000	10.684	10.768	-1.833*	11.323	11.377	-1.350
Construction	3 Month	11.762	11.762	0.000	11.672	11.291	1.833*	11.095	11.148	-1.350
Real Estate	3 Month	12.164	12.164	0.000	12.265	11.890	1.833*	11.707	11.761	-1.350
Other	3 Month	14.033	14.033	0.000	13.694	13.434	1.833*	13.542	13.596	-1.350
Agriculture	6 Month	11.432	11.432	0.000	11.412	11.071	1.833*	10.696	10.804	-1.350
Cattle Ranch	6 Month	11.497	11.497	0.000	11.469	11.140	1.830*	10.794	10.901	-1.350
Industrial	6 Month	11.205	11.205	0.000	11.394	10.985	1.830*	11.451	10.559	-1.350
Construction	6 Month	11.606	11.606	0.000	11.897	11.554	1.830*	11.063	11.171	-1.350
Real Estate	6 Month	10.121	10.121	0.000	10.405	10.461	-1.830*	10.679	10.787	-1.350
Other	6 Month	13.476	13.476	0.000	13.405	13.127	1.833*	12.899	13.007	-1.350
		October 1, 2001 (29,30)			**May 1, 2001 (33, 34)**			**September 1, 2001 (35,36)**		
Agriculture	1 Month	11.683	11.983	0.000	11.888	11.888	0.000	12.116	12.029	0.800
Cattle Ranch	1 Month	11.760	11.760	0.000	11.985	11.985	0.000	12.213	12.122	0.800
Industrial	1 Month	11.755	11.755	0.000	11.643	11.643	0.000	11.870	11.847	0.800
Construction	1 Month	12.794	12.794	0.000	12.255	12.255	0.000	12.483	12.545	-0.800
Real Estate	1 Month	12.134	13.134	0.000	11.871	11.871	0.000	11.863	11.917	-0.800
Other	1 Month	13.844	13.844	0.000	14.090	14.090	0.000	14.318	14.251	0.800
Agriculture	3 Month	11.269	11.269	0.000	11.475	11.475	0.000	11.702	11.616	0.800
Cattle Ranch	3 Month	11.347	11.347	0.000	11.572	11.572	0.000	11.800	11.709	0.800
Industrial	3 Month	11.721	11.721	0.000	11.457	11.457	0.000	11.449	11.503	-0.800
Construction	3 Month	11.342	11.342	0.000	11.229	11.229	0.000	11.457	11.434	0.800
Real Estate	3 Month	12.380	12.380	0.000	11.841	11.841	0.000	12.069	12.131	-0.800
Other	3 Month	13.431	13.431	0.000	13.677	13.677	0.000	13.905	13.838	0.800
Agriculture	6 Month	10.761	10.761	0.000	10.965	10.965	0.000	11.194	11.107	0.800
Cattle Ranch	6 Month	10.838	10.838	0.000	11.063	11.063	0.000	11.291	11.200	0.800
Industrial	6 Month	10.833	10.833	0.000	10.720	10.720	0.000	10.948	10.925	0.800
Construction	6 Month	11.871	11.871	0.000	11.333	11.333	0.000	11.561	11.663	-0.800
Real Estate	6 Month	11.212	11.212	0.000	10.949	10.949	0.000	10.941	10.995	-0.800
Other	6 Month	12.922	12.922	0.000	13.168	13.168	0.000	13.396	13.329	0.800

Table 2: (Continued)

		Effective Date								
		February 1, 2002 (37, 38)			January 16, 2003 (39, 40)			February 16, 2003 (41, 42)		
		Pre	Post	M-W	Pre	Post	M-W	Pre	Post	M-W
Agriculture	1 Month	12.293	12.293	0.000	15.031	15.323	-2.327**	15.375	15.543	-1.549
Cattle Ranch	1 Month	12.379	12.379	0.000	14.913	15.234	-2.327**	15.311	15.471	-1.549
Industrial	1 Month	12.157	12.157	0.000	15.054	15.338	-2.327**	15.485	15.627	-0.775
Construction	1 Month	13.003	13.003	0.000	15.627	15.951	-2.327**	16.271	16.393	-0.775
Real Estate	1 Month	12.134	12.134	0.000	14.509	14.509	-2.327**	14.836	14.964	-0.775
Other	1 Month	14.557	14.557	0.000	17.933	18.115	-2.327**	18.657	18.737	-0.775
Agriculture	3 Month	11.880	11.880	0.000	13.691	13.983	-2.327**	14.030	14.198	-1.549
Cattle Ranch	3 Month	11.965	11.965	0.000	13.572	13.893	-2.327**	13.966	14.126	-1.549
Industrial	3 Month	11.721	11.721	0.000	13.168	13.453	-2.327**	13.491	13.620	-0.775
Construction	3 Month	11.744	11.744	0.000	13.714	13.999	-2.327**	14.140	14.282	-0.775
Real Estate	3 Month	12.590	12.590	0.000	14.356	14.610	-2.327**	14.926	15.048	-0.775
Other	3 Month	14.144	14.144	0.000	16.593	16.774	-2.327**	17.312	17.392	-0.775
Agriculture	6 Month	11.371	11.371	0.000	12.745	13.037	-2.327**	13.086	13.254	-1.549
Cattle Ranch	6 Month	11.457	11.457	0.000	12.626	12.947	-2.327**	13.021	13.182	-1.549
Industrial	6 Month	11.235	11.235	0.000	12.768	13.052	-2.327**	13.196	13.338	-0.775
Construction	6 Month	12.082	12.082	0.000	13.411	13.665	-2.327**	13.982	14.104	-0.775
Real Estate	6 Month	11.212	11.212	0.000	12.222	12.508	-2.327**	12.547	12.675	-0.775
Other	6 Month	13.635	13.635	0.000	15.647	15.829	-2.327**	16.368	16.447	-0.775
		March 16, 2003 (43, 44)			September 1, 2004 (45, 46)			October 1, 2004 (47, 48)		
Agriculture	1 Month	15.834	15.886	-1.350	12.007	11.990	2.880***	12.240	12.240	0.000
Cattle Ranch	1 Month	15.679	15.730	-1.350	11.907	11.890	2.880***	12.140	12.140	0.000
Industrial	1 Month	15.797	15.848	-1.350	12.297	12.230	2.880***	12.480	12.480	0.000
Construction	1 Month	16.586	16.637	-1.350	12.587	12.580	2.880***	12.830	12.830	0.000
Real Estate	1 Month	15.116	15.167	-1.350	9.977	9.970	2.880***	10.120	10.120	0.000
Other	1 Month	18.874	18.925	-1.350	13.657	13.650	2.880***	13.920	13.920	0.000
Agriculture	3 Month	14.485	14.570	-1.350	11.087	11.069	2.880***	11.319	11.319	0.000
Cattle Ranch	3 Month	14.330	14.415	-1.350	10.987	10.969	2.880**	11.219	11.219	0.000
Industrial	3 Month	13.767	13.852	-1.350	9.057	9.049	2.880***	9.199	9.199	0.000
Construction	3 Month	14.449	14.534	-1.350	11.377	11.309	2.880***	11.559	11.559	0.000
Real Estate	3 Month	15.237	15.322	-1.350	11.667	11.659	2.880***	11.909	11.909	0.000
Other	3 Month	17.525	17.610	-1.350	12.737	12.729	2.880***	12.999	12.999	0.000
Agriculture	6 Month	13.543	13.611	-1.350	10.585	10.577	2.880***	10.825	10.825	0.000
Cattle Ranch	6 Month	13.387	13.455	-1.350	10.488	10.475	2.880***	10.725	10.725	0.000
Industrial	6 Month	13.506	13.574	-1.350	10.875	10.815	2.880***	11.065	11.065	0.000
Construction	6 Month	14.295	14.362	-1.350	11.165	11.165	0.000	11.415	11.415	0.000
Real Estate	6 Month	12.824	12.892	-1.350	8.555	8.555	0.000	8.705	8.705	0.000
Other	6 Month	16.582	16.650	-1.350	12.235	12.235	0.000	12.505	12.505	0.000
		July 16, 2005 (49,50)			August 16, 2005 (51,52)					
Agriculture	1 Month	12.730	12.730	0.000	12.570	12.474	2.327**			
Cattle Ranch	1 Month	12.770	12.770	0.000	12.600	12.552	2.327**			
Industrial	1 Month	12.720	12.720	0.000	12.520	12.416	2.327**			
Construction	1 Month	13.340	13.340	0.000	13.010	12.994	2.327**			
Real Estate	1 Month	10.760	10.760	0.000	10.600	10.520	2.327**			
Other	1 Month	14.170	14.170	0.000	14.010	13.922	2.327**			
Agriculture	3 Month	11.750	11.750	0.000	11.590	11.494	2.327**			
Cattle Ranch	3 Month	11.790	11.790	0.000	11.620	11.572	2.327**			
Industrial	3 Month	9.780	9.780	0.000	9.620	9.540	2.327**			
Construction	3 Month	11.740	11.740	0.000	11.540	11.436	2.327**			
Real Estate	3 Month	12.360	12.360	0.000	12.030	12.014	2.327**			
Other	3 Month	13.190	13.190	0.000	13.030	12.942	2.327**			
Agriculture	6 Month	9.020	9.050	-0.800	9.010	8.914	2.327**			
Cattle Ranch	6 Month	9.060	9.090	-0.800	9.040	8.992	2.327**			
Industrial	6 Month	9.010	9.040	-0.800	8.960	8.856	2.327**			
Construction	6 Month	9.630	9.660	-0.800	9.450	9.434	2.327**			
Real Estate	6 Month	7.050	7.080	-0.800	7.040	6.96	2.327**			
Other	6 Month	10.460	10.490	-0.800	10.450	10.362	2.327**			

*This table shows average interest rate spreads for the five day periods before (Pre) and after (Post) each reserve requirement changes for securities denominated in Costa Rica Colon. The spread is the loan rate minus the corresponding deposit rate. ***, ** and * indicate significance at the 1%, 5% and 10% level respectively. The Mann-Whitney Test (M-W) is for differences in interest rate spreads before and after the announcement.*

Table 3 shows results of tests on dollar denominated securities. Presentation of the results is in a manner analogous to the colon denominated securities analysis. The findings for dollar denominated securities show a higher level of significance than for the colon denominated securities. The results show significant spread changes for four of eight reserve requirement changes. From the remaining reserve requirement changes, evidence of a spread change is evident in two cases, but the results are not significant. No clear pattern is present with regard to the sequence of effective dates. For the July 26, 2004 reserve requirement change, a significant spread change occurred for the first effective data but not for subsequent effective dates. For the June 15, 2005 change, no significant spread change was found for the first effective date associated with the reserve requirement announcement. However, the second effective date associated with the same announcement produced significant results. Three significant results involved the spread widening and one significant result involved the spread narrowing.

Combined, the results provide some evidence that national banks change their interest rate spreads around effective dates of reserve requirement changes. The evidence of change is stronger among dollar denominated securities than for colon denominated securities. These findings explain the lack of significance around decision dates and announcement dates found by Stewart, Jalbert and Jalbert (2010). Indeed, government banks do not respond at the decision date or announcement date. However, they do respond at the effective date of the reserve requirement change.

CONCLUDING COMMENTS

This paper examines the profitability theory of reserve requirement changes using data from Costa Rica. The profitability theory is typically tested by examining how bank stock prices respond to changes in reserve requirements. In Costa Rica, bank stock data is generally not available, so the normal test approach is not feasible. To circumvent this limitation, we measure financial institution profitability through a proxy variable. Specifically, profitability is measured as the spread between interest rates earned on loans and paid on deposits. An increase in spreads indicates improved profitability for banks while declining spreads indicate lower bank profits.

The process of changing reserve requirements in Costa Rica involves three steps. First, the Central Bank makes a decision to change reserve requirements. A few days to a few weeks later the change is publicly announced along with the effective date of the change. The reserve requirement change becomes policy on the effective date. Previous studies examined interest rate level and spread responses to the Central Bank's decision to change reserve requirements and to the public announcement that reserve requirements would change. These authors note that Costa Rica Governments banks do not respond to reserve requirement changes at either the decision or announcement dates. Private banks on the other hand do respond to the announcement by changing interest rate levels and spreads. Given the lack of response by government banks on the decision and announcement dates, this paper extends the analysis to identify government bank responses to reserve requirement changes on the effective change date. The hypotheses are tested for reserve requirement changes that occurred between 1996 and 2005.

The results show that Costa Rica government banks change their interest rate spreads around the effective dates of a reserve requirement change. The results are mixed for colon denominated securities, but are remarkably strong for dollar denominated securities. This finding suggests that the careful entrepreneur can time loans and deposits to take best advantage of these predictable changes in interest rate spreads.

The tests conducted here examine only data from Costa Rica. The results should not be generalized to other countries as the Costa Rica financial system has many unique peculiarities. Future research should analyze the reaction of private and government banks in other countries. The results here are limited to the examination of a specific announcement window. Further research should examine other response windows to determine the time frame for bank response.

Table 3: Mann-Whitney Test: Dollar Denominated Interest Rate Spreads

		EFFECTIVE DATE								
		March 1, 1997 (1,2)			January 16, 2003 (3,4)			February 16, 2003 (5,6)		
		Pre	Post	M-W	Pre	Post	M-W	Pre	Post	M-W
Agriculture	1 Month	5.925	5.936	-1.350	5.166	5.738	2.327**	5.944	5.944	0.000
Cattle Ranch	1 Month	5.925	5.936	-1.350	5.602	5.548	2.327**	5.772	5.772	0.000
Industrial	1 Month	5.925	5.936	-1.350	5.622	5.802	2.327**	6.030	6.030	0.000
Construction	1 Month	5.975	5.986	-1.350	6.245	5.952	2.327**	6.446	6.446	0.000
Real Estate	1 Month	5.175	5.186	-1.350	6.184	5.810	2.327**	6.234	6.234	0.000
Other	1 Month	5.413	5.425	-1.350	7.182	6.493	2.327**	7.541	7.541	0.000
Agriculture	3 Month	5.613	5.625	-1.350	5.831	5.404	2.327**	5.611	5.611	0.000
Cattle Ranch	3 Month	5.613	5.625	-1.350	5.685	5.214	2.327**	5.439	5.439	0.000
Industrial	3 Month	5.613	5.625	-1.350	5.878	5.468	2.327**	5.696	5.696	0.000
Construction	3 Month	5.663	5.675	-1.350	5.912	5.619	2.327**	6.112	6.112	0.000
Real Estate	3 Month	4.863	4.875	-1.350	5.850	5.477	2.237**	5.900	5.900	0.000
Other	3 Month	5.102	5.116	-1.350	6.848	6.159	2.327**	7.208	7.208	0.000
Agriculture	6 Month	5.547	5.558	-1.350	5.632	5.204	2.327**	5.411	5.411	0.000
Cattle Ranch	6 Month	4.670	4.682	-1.350	9.451	9.772	-2.327**	9.912	10.072	-1.549
Industrial	6 Month	5.547	5.559	-1.350	5.568	5.268	2.327**	5.496	5.496	0.000
Construction	6 Month	5.597	5.608	-1.350	5.712	5.419	2.327**	5.912	5.912	0.000
Real Estate	6 Month	4.797	4.808	-1.350	5.650	5.277	2.327**	5.700	5.700	0.000
Other	6 Month	5.035	5.047	-1.350	6.649	5.959	2.327**	7.008	7.008	0.000
		March 16, 2003 (7,8)			September 1, 2004 (9, 10)			October 1, 2004 (11,12)		
Agriculture	1 Month	5.631	5.631	0.000	6.761	6.761	0.000	6.867	7.011	-1.833*
Cattle Ranch	1 Month	5.948	5.948	0.000	6.781	6.771	2.880***	6.883	7.021	-1.833*
Industrial	1 Month	5.904	5.904	0.000	7.601	7.841	-2.880***	8.002	8.081	-1.830*
Construction	1 Month	6.392	6.392	0.000	7.811	7.771	2.880***	7.825	7.921	-1.830*
Real Estate	1 Month	6.141	6.142	-1.350	6.721	6.671	2.880***	6.757	6.871	-1.830*
Other	1 Month	7.531	7.285	1.350	7.911	7.731	2.880***	7.835	7.931	-1.830*
Agriculture	3 Month	5.297	5.297	0.000	6.448	6.443	2.880***	6.549	6.549	-1.830*
Cattle Ranch	3 Month	5.614	5.614	0.000	6.468	6.453	2.880***	6.565	6.703	-1.830*
Industrial	3 Month	5.570	5.570	0.000	7.288	7.523	-2.880***	7.685	7.763	-1.830*
Construction	3 Month	6.058	6.058	0.000	7.798	7.453	2.880***	7.525	7.603	-1.830*
Real Estate	3 Month	5.807	5.808	-1.350	6.408	6.353	2.880***	6.439	6.553	-1.833*
Other	3 Month	7.197	6.951	1.350	7.598	7.413	2.880***	7.517	7.613	-1.833*
Agriculture	6 Month	5.097	5.097	0.000	6.180	6.170	2.880***	6.272	6.410	-1.833*
Cattle Ranch	6 Month	10.277	10.345	-1.350	8.295	8.275	2.880***	8.521	8.515	1.833*
Industrial	6 Month	5.370	5.370	0.000	7.020	7.250	-2.880***	7.408	7.480	-1.833*
Construction	6 Month	5.858	5.858	0.000	7.230	7.180	2.880***	7.248	7.320	-1.833*
Real Estate	6 Month	5.607	5.608	-1.350	6.140	6.080	2.880***	6.162	6.270	-1.833*
Other	6 Month	6.997	6.751	1.350	7.330	7.140	2.880***	7.240	7.330	-1.833*
		July 16, 2005 (13,14)			August 16, 2005) (15,16)					
Agriculture	1 Month	9.580	9.580	0.000	9.190	9.998	-2.327**			
Cattle Ranch	1 Month	8.580	8.580	0.000	8.730	8.898	-2.327**			
Industrial	1 Month	9.350	9.350	0.000	9.770	9.866	-2.307**			
Construction	1 Month	8.900	8.900	0.000	8.700	8.876	-2.327**			
Real Estate	1 Month	9.210	9.210	0.000	9.370	9.482	-2.327**			
Other	1 Month	9.870	9.870	0.000	10.000	10.104	-2.327**			
Agriculture	3 Month	9.270	9.270	0.000	9.600	9.688	-2.327**			
Cattle Ranch	3 Month	8.270	8.270	0.000	8.240	8.588	-2.327**			
Industrial	3 Month	9.040	9.040	0.000	9.460	9.556	-2.327**			
Construction	3 Month	8.590	8.590	0.000	8.390	8.566	-2.327**			
Real Estate	3 Month	8.900	8.900	0.000	9.060	9.172	-2.327**			
Other	3 Month	9.560	9.560	0.000	9.690	9.794	-2.327**			
Agriculture	6 Month	8.970	8.970	0.000	9.300	9.388	-2.327**			
Cattle Ranch	6 Month	6.810	6.840	0.000	6.790	6.742	2.327**			
Industrial	6 Month	8.740	8.740	0.000	9.160	9.256	-2.327**			
Construction	6 Month	8.290	8.290	0.000	8.090	8.266	-2.327**			
Real Estate	6 Month	8.600	8.600	0.000	8.760	8.872	-2.327**			
Other	6 Month	9.260	9.260	0.000	9.390	9.494	-2.327**			

*This table shows average interest rate spreads for the five day periods before (Pre) and after (Post) each reserve requirement change for securities denominated in U.S. dollars. The spread is the loan rate minus the corresponding deposit rate. ***, ** and * indicate significance at the 1%, 5% and 10% level respectively. The Mann-Whitney Test (M-W) is for differences in interest rate spreads before and after the announcement.*

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MARKET COMPETITION AND MERGERS IN PROFESSIONAL SERVICE FIRMS

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ABSTRACT

This study examines competition level and merger in Taiwanese audit industry over a long time interval of 1992-2008. Total public accounting firms are divided into four sub-samples in terms of market segment, including big, large, medium, and small firms. Next, based on prior studies and service attribute, this study establishes four practice sub-markets: auditing, tax, consultation, and accounting. Empirical results indicate that big firms have the highest competition level but the other three sub-samples show no significant differences in competition level. Next, the auditing, tax, and accounting sub-markets become more concentrated over time but consultation sub-market does not change significantly. Big firms exhibit the highest competition level in the four sub-markets and four sub-markets but achieve the best in three financial performance measures, net profit per partner, profit ratio, and productivity per employee. Post-merger firms financially outperform pre-merger firms for Taiwanese two big firms' mergers occurred in 1999 and 2003.

JEL: M42

KEYWORDS: Market Competition, Herfindahl-Hirschman Index, Operating Performance, Public Accounting Firms

INTRODUCTION

The global economy is changing from an industry-based to a knowledge-intensive landscape which transforms the basis of technological innovation and corporate competition (Drucker, 1994; Van de Ven, 2004). Public accounting firms (also referred to as audit firms) are a professional service organization and a knowledge-intensive entity. Following the Enron debacle, the world largest public accounting firms then, Arthur Andersen, dissolved in 2002. This caused its Taiwanese affiliate firm to combine with that of Deloitte & Touche and created the largest public accounting firm in Taiwan, Deloitte Touche Tohmatsu, in 2003. Previous research indicates that regulatory agencies are concerned about market concentration and closely monitor any mergers between large public accounting firms due to increased post-merger audit market concentration (e.g., McMeeking, Peasnell and Pope, 2007). For example, in its 2008 report on market concentration for audits of public company, the U.S. Government Accountability Office indicated that the market concentration at that time lacked significant adverse concentration effect required for immediate action (GAO, 2008). In theory, a high market concentration level denotes low competition in a market (Besanko, Dranove and Shanley, 2000). Information about market concentration of service industry is useful to governmental agencies, companies, and academics alike (Wernerheim, 2010). Because of regulations, mergers between firms, and economic development, market concentration levels change over time (Jennequin, 2008). Previous studies typically focus on the short-term or discrete audit market concentration, which motivates this study to examine the long-term concentration to fill the gap left.

Prior studies document the existence of market segmentation in the audit market owing to government regulations or the size of clients served (DeFond, Francis and Wong, 2000; Ghosh and Lustgarten, 2006). Public accounting firms are often categorized in terms of their size, service area, or practices offered. Market segmentation leads to varied market concentrations in different public accounting firm categories. Base on prior studies (Ghosh and Lustgarten, 2006; Chen, Chang and Lee, 2008), this study partitions total public accounting firms into four sub-samples to better reflect audit market attributes. The first purpose of this study is to examine the long-term audit market concentration for each sub-sample. Public accounting firms traditionally provide audit services but recently expand to other scope of services. When determining the effects of PricewaterhouseCoopers merger on audit market competition, the Australian Consumer and Competition Commission (ACCC) identifies six markets in which public accounting firms compete: audit services, accounting services, management consultancy services, corporate recovery and insolvency consulting, corporate financial consulting, and actuarial services (Goddard, 1998). The identification of six practice markets implies a varied degree of competition among different service markets. Previous studies use aggregate audit fee information to measure audit market concentration and cannot determine the concentration at any lower level (Minyard and Tabor, 1991). Audit market concentration in different practice markets provides useful information for regulators and practitioners in further. Based on the ACCC, this study divides total samples into four practice sub-markets and investigates their market concentration, which constitutes the second purpose of this study. With the primary concern that post-merger public accounting firms might possess monopoly power in the industry, regulatory agencies closely monitor mergers between large public accounting firms (e.g., McMeeking et al., 2007). Public accounting firms with monopoly power might reduce competition due to increased market concentration for audit services (Minyard and Tabor, 1991; McMeeking et al., 2007). Economic theory suggests that price-cost margins (profits) should be higher in more concentrated markets (Besanko et al., 2000). The third and final purposes of this study are to determine whether public accounting firms in a more concentrated market produce better operating performance, and whether merged public accounting firms lead to superior operating performance.

Empirical data of this study are from the 1992-2008 Survey Report of Public Accounting Firms in Taiwan. First, this study subdivides total public accounting firms into four sub-samples in terms of market segment, including big, large, medium, and small firms. Further, this study establishes four practice sub-markets: auditing, tax, consultation, and accounting. The Herfindahl-Hirschman index (HHI) is used to estimate the market concentration level for the four sub-samples and four practice sub-markets. This study establishes a standardized HHI to facilitate cross-sample market concentration comparisons. A greater standardized HHI represents a higher level of market concentration and a lower degree of market competition. Empirical results indicate varied long-term concentration levels among the four sub-samples and four sub-markets. In terms of long-term concentration level, big firms have the highest competition level but the other three sub-samples show no significant differences in competition level. Further, the long-term concentration levels do not change significantly for the four sub-samples. Next, the auditing and accounting sub-markets have significantly lower concentration level than the tax and consultation sub-markets. The market structures of the auditing, tax, and accounting sub-markets become more concentrated over time, while that of consultation sub-market does not change significantly. Big firms exhibit the lowest concentration level, the highest competition level, in the four sub-markets but achieve the best in three financial performance measures, net profit per partner, profit ratio, and productivity per employee. Post-merger firms financially outperform pre-merger firms for two big firms' mergers occurred in 1999 and 2003.

This study accesses to the audit fee information of all public accounting firms in the audit industry, and is the first to estimate audit market concentration based on a theoretically defined market share. The empirical results of this study contribute to the related literature by extending the findings of existing researches. First, previous studies tend to use a proxy variable of audit fees to examine short-term audit market concentration for large public accounting firms only (e.g., Minyard and Tabor, 1991; Thavapalan, Moroney and Simnett, 2002). This study estimates market concentration using long-term actual audit fees for both large and small public accounting firms, providing a better and more complete picture of market concentration in the audit industry. Second, previous studies only examine market concentration using aggregate data, and are unable to analyze at lower levels (e.g., Choi and Zeghal, 1999; Minyard and Tabor, 1991). Using detailed practice data for individual public accounting firm, this study examines market concentration at a lower level by identifying four practice sub-markets. This approach provides a unique evidence of audit market concentration in different practice markets. Third, economic theory suggests a higher price-cost margin (profit) in a more concentrated market (Besanko et al., 2000) and a positive association between market price and market concentration (Weiss, 1989). However, this study finds that public accounting firms in less concentrated market financially outperform those in more concentrated market do. The finding contributes to the industrial economic literature. Finally, findings of this study convey important managerial implications to the practitioners of service industry. In practice, product differentiation and overall cost leadership are two commonly used marketing strategies with which to achieve a sustainable competitive advantage and to earn abnormal rate of return in a hostile environment (Hall, 1980; Porter, 1980). Although various product differentiation alternatives exist, superior quality is the most adopted approach to characterize this strategy (Kiechel, 1981). This study finds that public accounting firms in a more competitive market produce superior financial performance to firms in less competitive market. Prior studies report that competition improves overall efficiency and technical efficiency (Lee, Park and Oh, 2000). This study argues that competition enhances service quality and thereby solicits more clients, resulting in performance improvement. For practitioners especially in service sector, upgrade of service quality is the most useful weapon to counter competition. The remainder of this study proceeds as follows. Section 2 reviews previous studies followed by the empirical data and sample classification shown in Section 3. Section 4 presents the standardized HHIs and compares the differences in market concentration for the four sub-samples and sub-markets. This study displays the comparisons of operating performance for different sub-samples in Section 5. Section 6 demonstrates the implications of the empirical results, while Section 7 makes additional tests and provides managerial implications. This study concludes in Section 8.

BACKGROUND AND LITERATURE REVIEW

Audit Market in Taiwan: The most authoritative source of information on Taiwanese audit industry comes from the Survey Report of Public Accounting Firms in Taiwan, published by Taiwanese Financial Supervisory Commission beginning in 1988. As shown in Table 1, the number of public accounting firms was 532 in 1992 and climbed to 913 in 2008, and the number of practicing Certified Public Accountants (CPAs), owners of the firms, was 1,066 in 1992 and rose to 1,910 in 2008. Growth rates of number of firms and practicing CPAs are 71.62% and 79.17%, respectively. In practice, main clients of public accounting firms include public companies (listed and non-listed) and small and medium-sized enterprises (SMEs). Table 1 indicates that the number of public companies was 1,337 in 1992 and climbed to 3,049 in 2008, resulting in a 128% growth for the period. For the same period, the number of SMEs was 871,726 in 1992 and rose to 1,234,749 in 2008 with a growth rate of 41.64%. To the extent, both the growth of number in firms and number in practicing CPAs run parallel with that of audit clients, public companies and SMEs.

Table 1: Audit Clients and Public Accounting Firms

Year	Number of Public Accounting Firms	Number of Practicing CPAs	Public Companies	Small & Medium Enterprises
1992	532	1,066	1,337	871,726
1993	613	1,208	1,444	901,768
1994	654	1,310	1,571	932,852
1995	690	1,393	1,775	991,615
1996	714	1,498	2,032	1,003,325
1997	703	1,469	2,537	1,020,435
1998	753	1,613	3,036	1,045,117
1999	783	1,720	3,470	1,060,738
2000	814	1,797	3,919	1,070,310
2001	769	1,660	3,787	1,078,162
2002	754	1,675	3,372	1,104,706
2003	719	1,622	3,312	1,146,352
2004	663	1,594	3,315	1,164,009
2005	702	1,672	3,189	1,226,095
2006	714	1,693	3,111	1,244,099
2007	774	1,738	3,091	1,236,586
2008	913	1,910	3,049	1,234,749

This table shows the information about numbers of audit clients, practicing CPAs, and public accounting firms.

Long-term cooperation between U.S. and Taiwanese audit industries has created a similar audit market structure in both countries. Many Taiwanese international public accounting firms became affiliates or members of U.S. international firms four decades ago. In addition to the international firm affiliations, many local firms are associated with other U.S. firms, such as BDO, Grant Thornton, and Baker Tilly

International. For the international firm affiliations, the Taiwanese six largest international firms, Big Six, included Arthur Andersen, KPMG, Price Waterhouse, Ernst & Young, Deloitte & Touche, and Coopers & Lybrand before 1999. The ranks of the largest international firms were further reduced to the Big Five when Taiwanese associates of Price Waterhouse and Coopers & Lybrand merged in 1999 to form the PricewaterhouseCoopers. The loss of Arthur Andersen leaves the Big Four international firms in Taiwan after 2003, including KPMG, PricewaterhouseCoopers, Ernst & Young, and Deloitte Touche Tohmatsu. The largest public accounting firm was always the Arthur Andersen except 1999 due to the merged PricewaterhouseCoopers. The successor of Arthur Andersen, Deloitte Touche Tohmatsu, has ranked the first since 2003. Some regulations over Taiwanese audit industry occurred in the past two decades. Beginning in 1988, Taiwanese authorities have raised the passing rate of the CPAs uniform examination, resulting in substantial increases in the number of qualified CPAs and in market competition.

The authorities abolished the long-standing audit fee standard to ensure fair audit market competition in 1998. Cancelling the audit fee standard adversely impacts the traditional auditing practice market. Since then, a rumor of price-cutting strategy for client solicitation has prevailed in the industry and leaded to enhanced market competition. Furthermore, the tax authorities established a tax agent system and legalized the provision of corporate registration and accounting services by tax agents to SMEs in 2004. By the end of 2008, the cumulative number of qualified tax agents who are eligible for practicing services has been 10,120, much more than the number of practicing CPAs, 1,910. Proprietorship public accounting firms have provided the same services to the SMEs for years. Tax agent legalization negatively influences proprietorship public accounting firms because of the competitive advantages the tax agents possess for a relatively lower service fees and easy service access by the clients.

Measure of Market Structure

Theoretically, the number of firms in a market and the firms' monopoly power can explain market structure. Industrial economics is particularly concerned with the relationships between market concentration, market behavior, pricing and market performance (Akehurst, 1984). Oligopoly theorists note a positive association between market price and degree of market concentration (Weiss, 1989). In a highly concentrated industry, a few suppliers dominate the industry and oligopoly behaviors appear. Conspiracy or collusion among these leading suppliers provides them with price-setting power (Yardley, Kauffman, Cairney and Albrecht, 1992). Prior accounting studies show that audit market structure is related to audit pricing, audit fees, and market power (McMeeking et al., 2007; Bandyopadhyay and Kao, 2004; Lee, 2005). Further, different market structure gives rise to varied levels of rivalry, fee-setting practices, and client turnover (Ghosh and Lustgarten, 2006). In theory, measuring a market's structure is a quick and accurate way to assess the likely nature of its competition (Besanko et al., 2000). Researchers typically use two measures of market structure. The first is the n-firm concentration ratio (CRn). This measure represents the combined market share of the n largest firms in a market,

$$CR_n = \sum_{i=1}^{n} MKS_i$$

where n is a specified number of the largest firms in a market and MKS_i represents the market share of firm i. Prior researches typically use sale revenues to calculate market share, but may also use other measures such as production capacity. The concentration ratio is often calculated for either the four (CR4), six (CR6), or eight (CR8) largest firms in the sample. Another commonly used measure of market structure is the Herfindahl-Hirschman index (HHI). This index equals the sum of squared market shares for all firms in a market,

$$HHI = \sum_{i=1}^{n} MKS_i^2$$

where n is the total number of firms in a market and MKS_i represents the market share of firm i. A market with n equal-sized firms generally has a mean HHI of 1/n. A market with n firms, either equal-sized or unequal-sized, has a mean HHI of 1/n too. The mean HHI is also called a base level HHI. In theory, the HHI can range from a minimum of close to zero (a perfectly competitive market) to a maximum of 10,000 points (a monopoly market).

As both CRn and HHI measure market concentration level, a high CRn or HHI denotes low competition in a market. The HHI is sensitive to the number of firms active in an industry and to varying activity levels across firms. The CRn suffers from a number of weaknesses, such as its lack of information about all firms in an industry and its equal weighting of the market shares of all firms (Cowling, Yusof and Vernon, 2000). In addition, CRn is not obvious about how to select the most appropriate n largest firms (Hardwick, 1996). The HHI conveys more information than the CRn (Besanko et al., 2000) and because it has the advantage that it takes into account the market share of every single firm in the industry (Hardwick, 1996). However, researchers state that to assess the circumstances surrounding the competitive interaction of firms to make conclusions about the nature of competition is essential, rather than rely solely on the CRn or HHI (Besanko et al., 2000). As annual data is available for all public accounting firms in the Taiwanese audit industry, this study uses the HHI to measure market structure.

Measurement Base for Audit Market Concentration

A multitude of prior studies use the CRn or HHI to examine audit market concentration. They use various proxies for audit fees, including sales or assets of audit clients, squared root of client sales, number of clients, and number of audits (Thavapalan et al., 2002). Concentration measures based on proxy variable often produce measurement errors Moizer and Turley (1987). As a result, some studies use actual public accounting firm revenue data, a more pertinent and homogeneous measure, to calculate CRn and HHI (e.g., Choi and Zeghal, 1999). Revenue data are preferable to any proxy variable because they provide a direct measure of audit market concentration. However, previous studies only use revenue data in aggregate amounts, and are therefore unable to examine concentration at any lower level without using proxy information (Minyard and Tabor, 1991). In contrast, this study analyzes detailed fee data of each firm, which make it possible to measure market concentration of the Taiwanese audit industry at lower level, a unique approach in the world. As a result, this study first calculates the theoretically defined market share as individual firm revenues divided by the total industry revenues. Next, to estimate the annual market concentration for each sub-sample or sub-market, this study calculates market share as individual firm revenues divided by the total revenues of each sub-sample or sub-market.

DATA METHODOLOGY

Data and Sample Classification

Taiwanese regulatory agencies began to administer public accounting firm survey in 1989 to collect business information on the audit industry for macro-economic analysis and industrial policy formation. The regulatory agencies publish the Survey Report of Public Accounting Firms in Taiwan annually but 1991, due to its data inseparable from other industry's statistics. The data used in this study are commercially available from the Financial Supervisory Commission, Taiwan. Surveyed items include total revenues and their compositions, total expenditures and their compositions, employee demographics, ending amounts of and changes in fixed assets. This survey also collects qualitative information using an open questionnaire that asks about operating difficulties encountered and future business strategies to be taken by public accounting firms. Because the survey is administered pursuant to the Statistics Act, the surveyed firms are obligated to fill out the questionnaire correctly and in a timely manner. The annual response rate, according to the Survey Report, exceeds eighty percent.

This study obtains empirical data from 1992 to 2008 to presents a continuous long-term analysis of audit market concentration. Market segmentation exists in audit industry due to either varied government regulation or size of clients served (Defond et al., 2000; Ghosh and Lustgarten, 2006). Market segmentation refers to a group of consumers within a broader market who possess a common set of characteristics. Segmentation characteristics include demographic factors, geography, buyer's industry, and size of purchasing firm (Besanko et al., 2000). Practically, the larger the company, the more complicated the organization structure, and the higher the internal agency cost. As a result, companies employ larger public accounting firms to audit their financial statements to alleviate the agency cost (Francis, Maydew and Sparks, 1999). Public companies are larger in size and revenues compared to private companies. Substantial difference in size exists between public accounting firms offering and not offering services to public companies. Hence, this study divides public accounting firms into four different categories in terms of market segmentation: big, large, medium, and small firms.

Big firms refer to the Taiwanese affiliates or members of the international firms in the U.S. during the sample period. Large firms are non-big partnership firms that provide audit services to public companies, while non-big partnership firms that do not offer this kind of service are medium firms. Small firms represent proprietorship firms. As the sample period of this study is 17 years, this study deflates all monetary variables by the yearly consumer price index to account for inflation. The Survey Report provides information about ten services that can be offered by public accounting firms. Based on the ACCC (Goddard, 1998) and Banker, Chang, and Cunningham (2003), this study groups these services into four practice sub-markets by their attributes: auditing, tax, consultation, and accounting sub-markets. In the auditing sub-market, four audit services are provided, including auditing the financial statements of public companies, auditing financial statements for granting a bank loan, auditing financial statements for other purposes, and auditing income tax returns. The tax sub-market offers such services as tax planning, administrative remedy of internal taxation, and other tax operation services. The consultation sub-market renders management advisory services. In the accounting sub-market, corporate registration and accounting and bookkeeping practices are served.

During the sample period, we delete firm-year observations: (1) newly established in the survey year; (2) with dependent variable having value more or less than three standard deviations away from its mean; and (3) with no revenue or no expenditure. The final sample consists of 12,264 firm-year observations after 143 observations are deleted. Panel A in Table 2 reports the annual number and percentage of each sub-sample. The sample includes 86 big firms and 972 large firms, which accounting for 0.72% and 8.01% of the final observations, respectively. Number of medium and small firms is 2,795 and 8,411, which depicts 22.64% and 68.63% of the final observations. Panel B in Table 2 shows the annual total revenues and market share for each sub-sample. Total revenues of big firms was NT$ 2,386 million in 1992 and soared up to NT$ 15,200 million in 2008, while the market shares increased from 41.38% in 1992 to 62.81% in 2008. Big firms earned over half of the total industry revenues with NT$ 8,010 million in 2000, and have dominated the audit market ever since, a situation similar to western countries such as the U.S. and U.K. (Daniels, Leyshon and Thrift, 1988).

A dual market structure exists in the audit market with a few large public accounting firms and many small ones (Brocheler, Maijoor and Witteloostuijn, 2004). Table 2 shows a market structure with a few large firms (i.e., big firms) and many small firms (i.e., medium and small firms) in Taiwan. Big firms, on average, account for only 0.72% of the number of observations, but earn 51.19% of the total revenues. Conversely, medium and small firms account for 91.27% (22.64%+68.63%) of the number of observations, but earn only 28.20% (14.67%+13.53%) of the total revenues. Hence, the Taiwanese audit market structure is similar to that in the U.S. and in most other western countries. Panel C in Table 2 displays annual revenues and market share for each sub-market. Auditing and accounting services are traditional practices that public accounting firms have provided for years. These traditional services are law-protected and statutory practices, auditing and accounting sub-markets occupy 73.52% and 12.61% of the total industry revenues. Total revenues of auditing sub-market was NT$ 1,879 million in 1992 increasing to NT$ 10,993 million in 2008. However, the corresponding market share of the auditing sub-market was 78.74% in 1992, and fell to 72.32% in 2008. Similarly, total revenues of the accounting sub-market increased from NT$ 327 million to 1,617 million from 1992 to 2008, while its market share dropped from 13.72% in 1992 to 10.64% in 2008. In contrast, tax and consultation services grew steadily from a market share of 7.55% (3.46%+4.09%) in 1992 to 17.03% (11.55%+5.48%) in 2008.

Table 2: Annual Numbers, Revenues, and Market Share of Public Accounting Firms

Panel A Number and Percentage of Public Accounting Firms								
Year	**Big Firms**	**%**	**Large Firms**	**%**	**Medium Firms**	**%**	**Small Firms**	**%**
1992	6	1.13	55	10.34	87	16.35	384	72.18
1993	6	0.98	54	8.81	126	20.55	427	69.66
1994	6	0.92	53	8.10	147	22.48	448	68.50
1995	6	0.87	64	9.28	142	20.58	478	69.28
1996	6	0.84	71	9.94	153	21.43	484	67.79
1997	6	0.85	61	8.68	158	22.48	478	67.99
1998	6	0.80	68	9.03	164	21.78	515	68.39
1999	5	0.64	66	8.43	176	22.48	536	68.45
2000	5	0.61	68	8.35	177	21.74	564	69.29
2001	5	0.65	56	7.28	173	22.50	535	69.57
2002	5	0.66	56	7.43	175	23.21	518	68.70
2003	4	0.56	54	7.51	175	24.34	486	67.59
2004	4	0.60	48	7.24	162	24.43	449	67.72
2005	4	0.57	52	7.41	172	24.50	474	67.52
2006	4	0.56	46	6.44	184	25.77	480	67.23
2007	4	0.52	48	6.20	198	25.58	524	67.70
2008	4	0.44	52	5.70	226	24.75	631	69.11
1992-2008	86	0.72	972	8.01	2,795	22.64	8,411	68.63
Panel B Total Revenues and Market Share for Four Sub-Samples (in Million New Taiwan Dollars)								
Year	**Big Firms**	**%**	**Large Firms**	**%**	**Medium Firms**	**%**	**Small Firms**	**%**
1992	2,386	41.38	1,601	27.77	673	11.67	1,107	19.19
1993	2,710	39.37	1,741	25.29	1,142	16.59	1,291	18.75
1994	3,061	38.14	1,845	22.99	1,646	20.51	1,473	18.35
1995	3,539	38.97	2,192	24.14	1,731	19.06	1,620	17.84
1996	4,202	40.53	2,669	25.75	1,732	16.71	1,765	17.02
1997	4,951	43.06	2,826	24.58	1,951	16.97	1,770	15.39
1998	5,867	44.90	3,388	25.93	1,937	14.82	1,876	14.35
1999	6,647	47.75	3,431	24.65	1,969	14.14	1,874	13.46
2000	8,010	50.58	3,599	22.73	2,254	14.23	1,972	12.46
2001	8,937	57.47	2,759	17.74	1,966	12.64	1,888	12.14
2002	9,217	58.32	2,770	17.53	2,080	13.16	1,737	10.99
2003	9,533	58.75	2,885	17.78	2,059	12.69	1,749	10.78
2004	10,100	61.47	2,410	14.67	2,270	13.82	1,650	10.04
2005	10,800	60.30	2,890	16.14	2,400	13.40	1,820	10.16
2006	12,500	62.72	2,880	14.45	2,600	13.05	1,950	9.78
2007	14,100	63.66	3,160	14.26	2,810	12.69	2,080	9.39
2008	15,200	62.81	3,400	14.05	3,060	12.64	2,420	10.00
1992-2008	131,760	51.19	46,446	20.61	34,280	14.67	30,042	13.53
Panel C Total Revenues and Market Share for Four Sub-Markets (in Million New Taiwan Dollars)								
Year	**Auditing**	**%**	**Tax**	**%**	**Consultation**	**%**	**Accounting**	**%**
1992	1,879	78.74	83	3.46	97	4.09	327	13.72
1993	2,114	78.02	90	3.32	155	5.72	351	12.94
1994	2,359	77.08	129	4.23	170	5.56	402	13.13
1995	2,754	77.83	135	3.81	189	5.35	461	13.02
1996	3,144	74.82	204	4.86	290	6.90	564	13.42
1997	3,722	75.17	302	6.10	273	5.51	655	13.22
1998	4,342	74.01	341	5.80	341	5.81	843	14.37
1999	4,960	74.62	421	6.33	361	5.43	905	13.62
2000	5,961	74.41	566	7.07	514	6.42	970	12.10
2001	6,728	75.29	652	7.29	546	6.11	1,011	11.31
2002	6,871	74.55	724	7.86	561	6.08	1,060	11.50
2003	7,038	73.83	844	8.86	561	5.88	1,090	11.43
2004	7,331	72.58	1,042	10.32	396	3.92	1,331	13.18
2005	7,885	73.01	1,051	9.73	440	4.07	1,424	13.18
2006	9,076	72.61	1,362	10.89	507	4.06	1,556	12.45
2007	10,225	72.52	1,614	11.44	682	4.84	1,579	11.20
2008	10,993	72.32	1,756	11.55	833	5.48	1,617	10.64
1992-2008	94,108	73.52	14,591	8.50	6,916	5.37	16,145	12.61

Table 2 reports the annual numbers, percentage, revenues, and market share of public accounting firms. Panel A shows the number and percentage for four sub-sample firms including big, large, medium, and small audit firms. Total revenues and market shares of the four sub-samples are appeared in Panel B. The Panel C reports the total revenues and market share for four sub-markets, auditing, tax, consultation, and accounting.

Measurement and Comparisons of Market Structure

Standardized HHI: Traditionally, previous studies often calculate the Herfindahl-Hirschman index (HHI) for the four (HHI4), six (HHI6), or eight (HHI8) largest firms in the sample under analysis. Because annual data are available for all public accounting firms in the Taiwanese audit industry, this study includes all firms in estimating the theoretically defined HHI and does not report the traditional ones for brevity. Theoretically, the HHI includes a base level (or mean HHI) that depends on the number of firms estimated. For example, if n firms are used to calculate the HHI, the base level is 1/n. This study estimates the annual HHI by the total number of public accounting firms for each sub-sample, shown in Panel A of Table 2. For example, there were 55 large firms in 1992, creating a year's base level HHI of 1/55, or 1.82%. As the annual number of observation for each sub-sample is different, this results in different base level HHI. Consequently, this study estimates a standardized HHI as a consistent benchmark for cross-sample comparisons. First, this study calculates the difference between actual HHI and the base level HHI to allow a concentration assessment above the base level. This difference represents an excess concentration:

Excess concentration = actual HHI − base level HHI = actual HHI − 1/n (1)

Then, the excess concentration is divided by the base level for each sub-sample. The resulting index is referred to as a standardized HHI as follows.

Standardized HHI = Excess concentration/base level HHI (2)

A greater standardized HHI denotes a relatively higher market concentration degree and a relatively lower market competition level. This study uses the same procedure to calculate standardized HHI for the four practice sub-markets. The following sections use these standardized HHIs to compare the concentration (competition) levels among sub-samples and sub-markets.

RESULTS

Comparisons of Competition Levels among Sub-Samples

Panel A in Table 3 shows the annual standardized HHIs for the four sub-samples. Big firms have the least mean standardized HHIs (0.10) followed by large firms (0.78) and small firms (0.79). Medium firms have the largest mean standardized HHIs (0.95). Panel B in Table 3 displays the testing results of differences in standardized HHIs among different sub-samples. This panel shows that big firms have less standardized HHIs than large, medium, and small firms, all significantly at the 1% level (Z = -4.902, -4.981, and -4.903). However, the pair-wise differences in standardized HHIs between large, medium, and small firms are statistically insignificant. This indicates that big firms have the least concentrated and the most competitive market in the long run among the four sub-samples. However, no significant difference exists in long-term competition levels among large, medium, and small firms.

Next, Figure 1 illustrates the annual standardized HHIs to present the long-term concentration level during the sample period. This figure shows that the standardized HHIs of big firms present a smooth upward movement. The standardized HHIs of large firms move upward until 2001, fall briefly, and then continue to rise after 2004. The long-term trend of standardized HHIs for medium firms varies, exhibiting a sharp decline in 1996, a significant increase after 2001, and then another fall after 2004. Finally, the standardized HHIs of small firms drift downward until 2006, when they began to rise.

To assess the long-term tendency of standardized HHIs statistically, this study performs the following linear regression.

$$YEAR = \gamma_0 + \gamma_1 \, SHHI + \varepsilon \quad (3)$$

where *YEAR* denotes year of 1992, 1993,…, and 2008, and *SHHI* is the standardized HHIs for each sub-sample. Regression results show that t-statistics of coefficients on the *SHHI* for big, large, medium, and small firms are 1.621, 1.189, 0.840, and -1.696. All coefficients are insignificant at the 10% level (two-tailed), indicating that the long-term level of market competition does not change significantly for the four sub-samples.

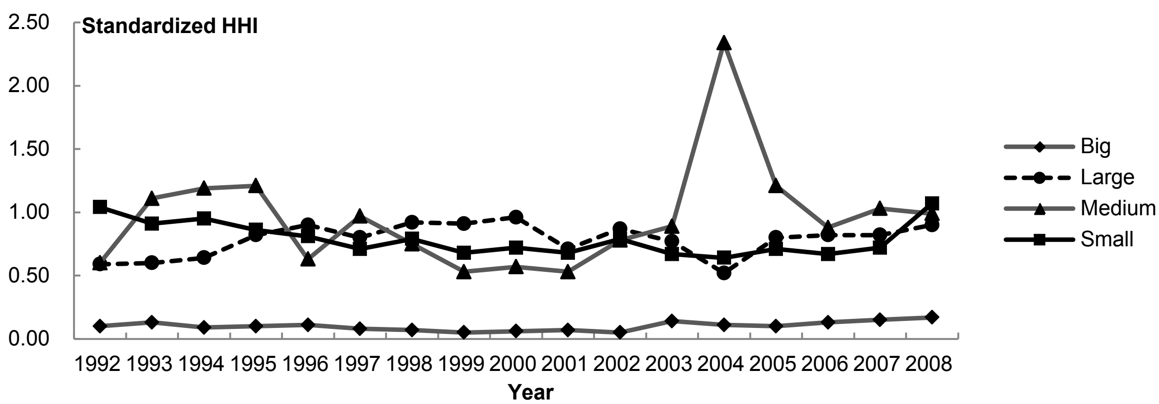
Table 3: Standardized HHIs and Differences between Sub-Samples

Panel A Standardized HHIs				
Year	Big Firms (n=86) (A)	Large Firms (n=972) (B)	Medium Firms (n=2,795) (C)	Small Firms (n=8,411) (D)
1992	0.10	0.59	0.60	1.04
1993	0.13	0.60	1.11	0.91
1994	0.09	0.64	1.19	0.95
1995	0.10	0.82	1.21	0.86
1996	0.11	0.90	0.63	0.81
1997	0.08	0.80	0.97	0.71
1998	0.07	0.92	0.75	0.79
1999	0.05	0.91	0.53	0.68
2000	0.06	0.96	0.57	0.72
2001	0.07	0.71	0.53	0.68
2002	0.05	0.87	0.78	0.79
2003	0.14	0.77	0.89	0.67
2004	0.11	0.52	2.34	0.64
2005	0.10	0.80	1.21	0.71
2006	0.13	0.82	0.88	0.67
2007	0.15	0.82	1.03	0.72
2008	0.17	0.90	0.99	1.07
Mean	0.10	0.78	0.95	0.79
Ho: Std.HHI=0	0.000***	0.000***	0.000***	0.000***

Panel B Test of Difference in Standardized HHIs						
	(A)-(B)	(A)-(C)	(A)-(D)	(B)-(C)	(B)-(D)	(C)-(D)
Wilcoxon Signed-Rank Test Z Statistics	-4.902***	-4.981***	-4.903***	-1.206	-0.362	-1.017
Asymptotic Significant Level	0.000	0.000	0.000	0.228	0.717	0.309

Panel A of this table displays the annual standardized HHIs estimated for four sub-samples, big, large, medium, and small audit firms. Panel B reports the testing results of differences in standardized HHIs among the four different sub-samples. *, **, *** *Denote two-tailed significance at the 10 %, 5 % and 1 % levels. Both t-test and two-sample sign test were performed to analyze the differences in standardized HHIs between sub-samples with the same results (un-tabulated) as reported here.*

Figure 1: Tendency of Standardized HHIs for Each Sub-Sample



This figure illustrates the long-term concentration level during the sample period by the annual standardized HHIs for four sub-samples, big, large, medium, and small audit firms.

Comparisons of Competition Levels among Sub-Markets

This study classifies four practice markets based on the sub-markets identified by the ACCC: auditing, tax, consultation, and accounting sub-markets. Panel A in Table 4 shows the annual standardized HHIs for the four sub-markets. As can be seen, the accounting sub-market has the least mean standardized HHIs (0.063) followed by the auditing (0.073) and tax (0.086) sub-markets. The consultation sub-market has the largest mean standardized HHIs (0.099). Next, this study uses the standardized HHIs to compare the differences in long-term concentration levels among the four sub-markets. Panel B in Table 4 shows that the auditing sub-market has significantly lower standardized HHIs than the tax and consultation sub-markets at the 10% and 5% levels, respectively (Z = -1.292 and -1.809). Similarly, the accounting sub-market has significantly lower standardized HHIs than the tax and consultation sub-markets at the 1% level (Z = -2.550 and -2.850). However, the differences in standardized HHIs between the auditing and accounting sub-markets and between the tax and consultation sub-markets are statistically insignificant. This indicates that the auditing and accounting sub-markets are the most competitive markets among the four sub-markets.

Table 4: Standardized HHI and Test of Difference between Sub-Markets

Panel A Standardized HHI				**(n=12,264)**
Year	**Auditing (A)**	**Tax (B)**	**Consultation (C)**	**Accounting (D)**
1992	0.034	0.087	0.032	0.056
1993	0.029	0.074	0.123	0.047
1994	0.027	0.054	0.093	0.042
1995	0.028	0.052	0.103	0.037
1996	0.029	0.066	0.159	0.036
1997	0.035	0.060	0.043	0.048
1998	0.037	0.074	0.084	0.050
1999	0.050	0.104	0.051	0.046
2000	0.057	0.098	0.095	0.054
2001	0.074	0.101	0.191	0.061
2002	0.076	0.083	0.128	0.060
2003	0.107	0.106	0.103	0.076
2004	0.113	0.115	0.056	0.105
2005	0.108	0.093	0.067	0.092
2006	0.117	0.118	0.086	0.103
2007	0.124	0.133	0.125	0.090
2008	0.202	0.047	0.149	0.072
Mean	0.073	0.086	0.099	0.063
Ho: Standardized HHI=0	0.000^{***}	0.000^{***}	0.000^{***}	0.000^{***}

Panel B Difference Test of Standardized HHI						
	(A)-(B)	**(A)-(C)**	**(A)-(D)**	**(B)-(C)**	**(B)-(D)**	**(C)-(D)**
Wilcoxon Signed-Rank Test Z Statistics	-1.292^{*}	-1.809^{**}	-0.034	-0.792	-2.550^{***}	-2.850^{***}
Asymptotic Significant Level	0.196	0.071	0.973	0.428	0.011	0.011

Panel A of this table shows the annual standardized HHIs estimated for the four sub-markets including auditing, tax, consultation, and accounting. The testing results of differences in standardized HHIs among the four different sub-markets are shown in Panel B. *, **, *** *Denote two-tailed significance at the 10 %, 5 % and 1 % levels. Both t-test and two-sample sign test were performed to analyze the difference in standardized HHIs between sub-markets with the same results (un-tabulated) as reported here.*

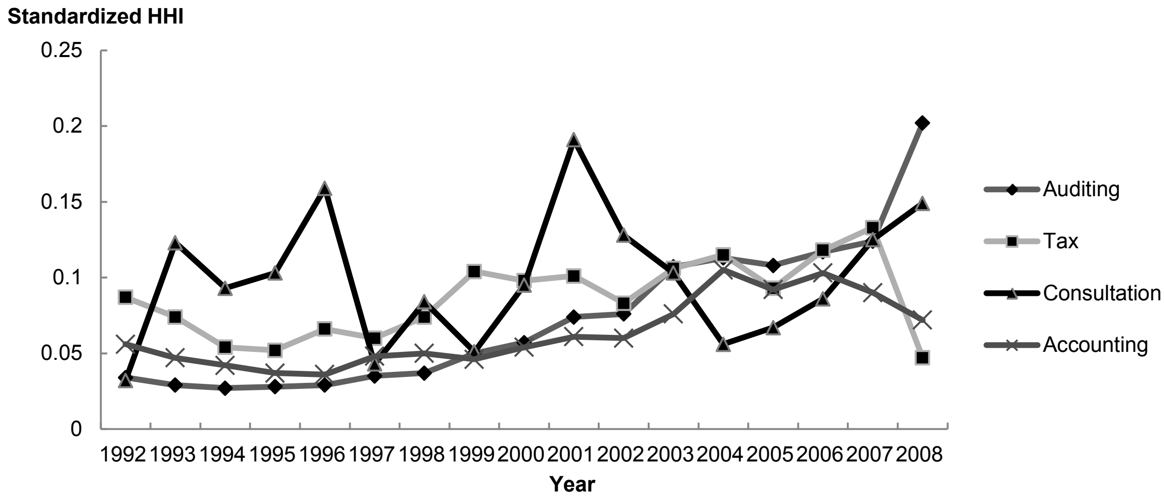
Next, Figure 2 demonstrates the annual standardized HHIs to present the long-term concentration levels during the sample period for the four sub-markets. As shown, the standardized HHIs of auditing sub-market present a steep upward movement. Both the standardized HHIs of taxing and accounting sub-markets move upward and fall after 2007 and 2006, respectively. The long-term trend of standardized HHIs for consultation sub-markets varies, exhibiting a sharp decline in 1996, a significant increase after 1999, another fall after 2001, and then beginning to rise in 2004.

To assess the long-term tendency of standardized HHIs statistically for the four sub-markets and based on model (3), this study performs a linear regression as follows.

$$YEAR = \gamma_0 + \gamma_1 \, SHHI + \varepsilon \qquad (4)$$

where *YEAR* denotes year of 1992, 1993,…, and 2008, and *SHHI* is the standardized HHIs for each sub-market. Coefficients on the *SHHI* of the auditing, tax, and accounting sub-markets have t-statistic of 8.630, 2.255, and 5.261, respectively. The coefficients are significant at either the 1% or 5% level (two tailed). However, the coefficient on the *SHHI* of consultation sub-market is positive but insignificant (t = 0.820). Empirical results show that while the long-term market structure of auditing, tax, and accounting sub-markets become more concentrated and less competitive, the consultation sub-market does not change significantly.

Figure 2: Tendency of Standardized HHIs for Each Sub-Market



This figure depicts the annual standardized HHIs to present the long-term concentration levels during the sample period for the four sub-markets, auditing, tax, consultation, and accounting.

Comparisons of Competition Levels Among Sub-Samples in Different Practice Sub-Markets

Standardized HHIs for the Four Sub-Samples in the Four Sub-Markets: Based on the results in Tables 3 and 4 for the sub-samples and sub-markets, this section further compares the concentration levels among sub-samples in different sub-markets and lists the results in Table 5. Panel A presents the mean standardized HHIs for each sub-sample in the four practice sub-markets. Specifically, big firms have the lowest standardized HHIs (0.12) in the four sub-markets. For example, the standardized HHI of big firms in the consultation sub-market is 1.34, followed by medium (5.69), large (5.88), and small firms (9.52).

Test of Differences in Standardized HHIs: Univariate test is used to compare the differences in concentration level among the four sub-samples in the four practice sub-markets. Panel B in Table 5 displays the results. In the auditing sub-market, big firms have significantly lower standardized HHIs than large, medium, and small firms at the 1% level (Z = -4.502, -4.508, and -4.501). However, the pair-wise differences in standardized HHIs between large, medium, and small firms are statistically insignificant in the auditing sub-market. These results indicate that the competition level of big firms is the highest in the auditing sub-market. Similarly, big firms have higher competition level than the other three sub-samples in the tax sub-market.

In the consultation sub-market, big firms have lower standardized HHIs than large, medium, and small firms significantly at the 1% level (Z = -3.676, -4.503, and -4.513). However, large and medium firms show statistically insignificant differences in standardized HHIs. Large firms have lower standardized HHIs than small firms at the 5% level (Z = -2.343), while medium firms are lower than small firms at the 1% level (Z = -4.135). These results reveal that big firms are the least concentrated in the consultation sub-market, followed by both large and medium firms and then by small firms.

In the accounting sub-market, big firms have significantly lower standardized HHIs than large, medium, and small firms at the 1% level (Z = -4.505, -4.535, and -4.537). Large firms have significantly lower standardized HHIs than medium and small firms at the 1% level (Z = -3.219 and -4.412), while medium firms are significantly lower than small firms at the 10% level (Z = -1.884). These findings indicate that competition levels of the four sub-samples in the accounting sub-market follow the increasing order of big, large, medium, and small firms.

In summary, big firms have the highest competition level in the auditing, tax, consultation, and accounting sub-markets. There is no significant difference in competition level among large, medium, and small firms in the auditing and tax sub-markets. Small firms have the lowest competition level in both consultation and accounting sub-markets.

Table 5: Standardized HHIs and Difference between Sub-Samples in Sub-Markets

Panel A Standardized HHI

Sub-Sample / Sub-Market	Big Firms (n=86) (A)	Large Firms (n=972) (B)	Medium Firms (n=2,795) (C)	Small Firms (n=8,411) (D)
Auditing	0.12***	1.30***	1.44***	1.29***
	(0.000)	(0.000)	(0.000)	(0.000)
Tax	0.13***	1.22***	3.09*	1.53***
	(0.001)	(0.002)	(0.080)	(0.001)
Consultation	1.34***	5.88***	5.69***	9.52***
	(0.000)	(0.000)	(0.000)	(0.000)
Accounting	0.29***	1.62***	3.30***	3.36***
	(0.000)	(0.000)	(0.001)	(0.000)
Ho: Mean Standardized HHI=0				

Panel B Test of Differences in Standardized HHIs between Sub-Samples

Sub-Market	Wilcoxon Signed-Rank Test Z-statistics (Asymptotic Significant Level)					
	(A)-(B)	(A)-(C)	(A)-(D)	(B)-(C)	(B)-(D)	(C)-(D)
Auditing	-4.502***	-4.508***	-4.501***	-0.851	-0.552	-0.965
	(0.000)	(0.000)	(0.000)	(0.395)	(0.581)	(0.335)
Tax	-4.408***	-4.518***	-4.521***	-0.230	-1.409	-1.175
	(0.000)	(0.000)	(0.000)	(0.818)	(0.159)	(0.240)
Consultation	-3.676***	-4.503***	-4.513***	-0.781	-2.343**	-4.135***
	(0.000)	(0.000)	(0.000)	(0.435)	(0.019)	(0.000)
Accounting	-4.505***	-4.535***	-4.537***	-3.219***	-4.412***	-1.884*
	(0.000)	(0.000)	(0.000)	(0.001)	(0.000)	(0.060)

*Panel A of this table shows the annual standardized HHIs estimated for the four sub-samples in the four sub-markets. The four sub-samples are big, large, medium, and small audit firms. The four sub-markets include auditing, tax, consultation, and accounting. Panel B reports the testing results of differences in standardized HHIs among the four different sub-samples in four sub-markets. *, **, *** Denote two-tailed significance at the 10 %, 5 % and 1 % levels. Both t-test and two-sample sign test were performed to analyze the differences in standardized HHIs with the same results (un-tabulated) as reported here.*

DISCUSSION OF THE RESULTS

Tables 3, 4, and 5 reveal some results with managerial implications. For example, big firms have the highest competition level among the four sub-samples and in the four practice sub-markets. Taiwanese big firms have associated with international public accounting firms in the U.S. for more than four decades. The members of these big firms share abundant resources, including professional auditing techniques and expertise, human resource development, and continuing professional education. Further, the headquarters of international firms determine the services offered by their worldwide members, who often exchange valuable information. With this systematic mechanism of professional development, big firms have become a symbol of high quality auditors, and their reputation remains strong in Taiwan.

For example, the 2008 Public Company Accounting Oversight Board (PCAOB) reports on two Taiwanese big firms, PricewaterhouseCoopers and Ernst and Young, indicate that the inspection team did not identify any quality control defects worthy of mention (PCAOB, 2008). As stated above, both big and large firms render audit services to public companies in Taiwan. As a result, the big firms in this study account for approximately 84% of the revenues generated from offering audit services to public companies, while large firms account for the remaining 16%. Because the general public rates each big firm with equal service quality, big firms compete with each other for business within the same market. For example, to solicit new client, a Taiwanese big firm sets a zero audit fee on initial audit engagements (China Times, April 7, 2003). The firm then maneuvers a low-balling practice to expand its market share and increases future profit by realizing client-specific quasi-rents to the incumbent firms. As big firms provide services with homogeneous quality, they exhibit the highest relative competition level among the four sub-samples and in the four practice sub-markets.

Although the long-term market structures of auditing and accounting sub-markets become less competitive, Table 4 indicates that both sub-markets have the highest long-term competition level among the four sub-markets. The two most important sub-markets for the sample period are auditing and accounting, which depicts approximately 73.52% and 12.61% of total industry revenues (see the Panel C in Table 2). Auditors lend credibility to financial statements by rendering auditing services and facilitate the sound functioning of capital market. Because the auditing and accounting services are a general requirement by various governmental agencies, they are services that clients need but do not necessarily want (Istvan, 1984). Earlier entrants to this market gain a competitive advantage over subsequent ones. The entry barrier makes the auditing and accounting sub-markets become less competitive in the long run. When compared with other sub-markets, however, both sub-markets have the highest long-term competition level due to the following regulations in Taiwanese audit industry. One is the raise of passing rate of the Certified Public Accountants (CPAs) uniform examination in 1988. Next, the Taiwanese Fair Trade Commission abolished the long-standing audit fee standard to ensure fair audit market competition in 1998. The other is the establishment of tax agent system to provide accounting services to small and medium-sized enterprises by tax agents in 2004.

In 1998 the Taiwanese Fair Trade Commission abolished the long-standing audit fee standard, established by the Taiwan Institute of CPAs, to ensure fair audit market competition. Since that time, market competition has increased. Further, the Ministry of Finance established a tax agent system and legalized the provision of corporate registration and accounting services by tax agents to small and medium-sized enterprises in 2004. The legalization of tax agents negatively affects the accounting sub-market due to the competitive advantages enjoyed by tax agents for their relatively lower fees and ease of service access by client. In contrast, the provision of consultation services typically requires greater involvement and communication between auditors and clients to meet specific service demands. Because consultation businesses are not regulated by the auditing standards, they are more flexible in formats, timing, and places of service provisions. Consultation services are often tailor-made and have no service fee standard, thereby making them more profitable than auditing and accounting services (Banker, Chang, and Cunningham, 2005). Consequently, consultation services create better business opportunities for auditors to expand their scope of services. Panel A in Table 4 shows that the consultation sub-market has the highest mean standardized HHIs (0.099) among the four sub-markets. Further, Panel A in Table 5 shows that consultation sub-market has highest mean standardized HHIs in the big firms (1.34), large firms (5.88), medium firms (5.69), and small firms sub-samples (9.52). These results indicate that the consultation sub-market is the least competitive of the four practice sub-markets for the four sub-samples.

COMPARISONS OF OPERATING PERFORMANCE FOR DIFFERENT SUB-SAMPLES

As shown in Tables 3 and 5, big firms have the least concentrated and the most competitive market among the four sub-samples in the long run. Economic theory suggests that price-cost margins (profits) should be higher in more concentrated markets (Besanko et al., 2000). Does a highly concentrated market lead to superior operating performance for firms in that market? Specifically, whether the big firms are inferior in operating performance to the large, medium, and small firms?

Operating performance can be assessed by non-financial measures, such as product quality and customer satisfaction index, or by financial measures, such as profit, return on assets (ROA), and return on invested capital (ROI). As a professional service organization, public accounting firms rarely possess the fixed assets typically owned by a manufacturing or merchandising company (Collins-Dodd, Gordon, and Smart, 2004). This study omits ROA and ROI and estimates three financial measures used in previous studies related to public accounting firms: net profit per partner (Chen et al., 2008), profit ratio (Fasci and Valdez, 1998), and productivity/product per employee (Collins-Dodd et al., 2004). The operational definitions of these financial performance measures are as follows.

1. Net profit per partner = (total revenues – total expenses + salaries paid to partners)/ number of partners;
2. Profit ratio = (total revenues – total expenses + salaries paid to partners)/ total revenues;
3. Productivity per employee = total revenues/ number of employees.

Partners are the owners and residual interest claimants of a public accounting firm. Their annual income comprises salaries and share of operating profits of the firm. The salaries of partners, weekly or monthly, are a part of total expenses of the firm. The more the salaries of the partners, the less the operating profits of the firm. It makes no difference to the partners whether they receive salaries or not in terms of their total annual income. In addition, the criteria for salary payments to partners vary across firms. Based on prior study (Chen et al., 2008), in calculating the net profit per partner and profit ratio, this study adds the partner salaries back to net income to reduce such an artificial noise.

Table 6 compares the operating performance for different sub-samples. Panel A shows that big firms have much higher net profit per partner (\$6,012,216) than large firms (\$1,619,321), medium firms (\$834,083), and small firms (\$671,045). Big firms have the highest profit ratio (0.243), followed closely by large firms (0.242), and then medium firms (0.216) and small firms (0.184). On average, productivity per employee is much higher in the big firms (\$1,414,679) than in the large firms (\$913,113), medium firms (\$743,546), and small firms (\$662,980).

Table 6: Comparisons of Operating Performance between Sub-Samples

Panel A Mean Operating Performance

	Big Firms (n=86) (A)	**Large Firms (n=972) (B)**	**Medium Firms (n=2,795) (C)**	**Small Firms (n=8,411) (D)**
Net Profit Per Partner	6,012,216	1,619,321	834,083	671,045
Profit Ratio	0.243	0.242	0.216	0.184
Productivity Per Employee	1,414,679	913,113	743,546	662,980

Panel B Test of Difference in Operating Performance

	(A)−(B) (t-statistic)	**(A)−(C) (t-statistic)**	**(A)−(D) (t-statistic)**	**(B)−(C) (t-statistic)**	**(B)−(D) (t-statistic)**	**(C)−(D) (t-statistic)**
Net Profit Per Partner	4,392,895	5,178,133	5,341,171	785,238	948,276	163,038
	(16.181)***	(19.239)***	(19.862)***	(18.924)***	(23.722)***	(8.957)***
Profit Ratio	0.001	0.027	0.059	0.026	0.058	0.032
	(0.259)	(4.088)***	(9.201)***	(5.090)***	(12.425)***	(7.207)***
Productivity Per Employee	501,566	671,133	751,699	169,567	250,133	80,566
	(9.453)***	(10.509)***	(8.499)***	(9.022)***	(14.310)***	(5.436)***

*Table 6 compares the operating performance among the four sub-samples. Measures of operating performance include net profit per partner, profit ratio, and productivity per employee. Panel A lists the mean operating performance for the four sub-samples and Panel B shows the testing results of difference in operating performance between different sub-samples. *, **, *** Denote two-tailed significance at the 10 %, 5 % and 1 % levels. Both net profit per partner and productivity per employee are expressed in new Taiwan dollars. A two-sample sign test was performed to analyze the difference in standardized HHIs with the same results (un-tabulated) as that reported here.*

Panel B shows performance differences between sub-samples. Big firms have higher net profit per partner than large firms (t = 16.181), medium firms (t = 19.239), and small firms (t = 19.862), all statistically significant at the 1% level. Large firms have significantly higher net profit per partner than medium firms (t = 18.924) and small firms (t = 23.722) at the 1% level. Finally, medium firms have significantly better net profit per partner than small firms at the 1% level (t = 8.957). The differences in profit ratio between big and large firms are positive but insignificant. However, big firms have a significantly higher profit ratio than medium and small firms at the 1% level (t = 4.088 and 9.201). Likewise, large firms have a significantly higher profit ratio than medium and small firms at the 1% level (t = 5.090 and 12.425).

The profit ratio of medium firms is significantly better than that of small firms (t = 7.207). Finally, in terms of productivity per employee, big firms are better than large firms (t = 9.453), large firms are better than medium firms (t = 9.022), and medium firms are better than small firms (t = 5.436). All of these results are statistically significant at the 1% level. Collectively, big firms perform the best in net profit per partner, profit ratio, and productivity per employee. Large firms outperform medium and small firms, and medium firms outperform small firms in the three performance measures. Classification of the four sub-samples used in this study is equivalent to audit firm size category. Hence, the larger the size of public accounting firms, the better the financial performance of the firms. Although big firms experience the highest long-term competition level, they achieve the best operating performance among the four sub-samples.

COMPARISONS OF OPERATING PERFORMANCE BETWEEN PRE- AND POST-MERGER PUBLIC ACCOUNTING FIRMS

Mergers between big international firms are notable events for regulators and academics (e.g., McMeeking et al., 2007). Two mergers between Taiwanese big firms occur during the sampling period. The first is the 1999 merger of Coopers and Lybrand (CL) and Price Waterhouse (PW), which shrinks the Big Six to the Big Five (hereafter, the PwC event). The second merger is the 2003 Arthur Andersen (AA) and Deloitte & Touche (DT) merger, which creates the largest firms in Taiwan, Deloitte Touche Tohmatsu (hereafter, the DTT event). The 1999 merger provides PricewaterhouseCoopers (PwC) with the largest market share and a two-firm concentration ratio (CR2) that climbed from 42.69% in 1998 to 49.19% in 1999. The Deloitte Touche Tohmatsu ranked first in market share in 2003 and a CR2 soared up from 48.94% in 2002 to 64.08% in 2003. These results confirm that mergers between public accounting firms yield increased concentration within the audit market (e.g., McMeeking et al., 2007).

Does the more concentrated and less competitive market lead to superior operating performance in the post-merger public accounting firms? Panel A in Table 7 displays the three financial performance figures of pre-merger and post-merger public accounting firms for the PwC and DTT events, including net profit per partner, profit ratio, and productivity per employee. The first two figures represent measures of individual firms, while the last figure represents their weighted average measure. In the PwC event, the first figure is for Price Waterhouse (PW) and the second is for Coopers and Lybrand (CL). In the DTT event, the first figure is for Arthur Andersen (AA) and the second is for Deloitte & Touche (DT). Size of PW is larger than CL, and AA is larger than DT before the merger. However, financial performance does not always the case, and depends upon the measures used. In the DTT event, for example, mean net profit per partner for AA ($5,817,561) in 1992 was higher than that of DT ($4,989,567), but mean productivity per employee for AA ($1,104,140) was lower than that of DT ($1,282,850).

Table 8 shows differences in operating performance between pre and post-merger accounting firms. Table 8, Panels 1, 2, and 3 report the statistical testing results of differences in the three financial performance measures between pre-merger and post-merger public accounting firms. Panel B-1 shows that post-merger firms have significantly higher net profit per partner than pre-merger firms for the PwC and DTT events at the 1% and 5% levels (t = 3.932 and 2.192; Z = 3.191 and 1.868). Panel B-2 indicates that post-merger firms have a lower profit ratio than pre-merger firms but insignificant for both the PwC and the DTT events.

Panel B-3 displays the differences in productivity per employee. Similar to the trend in Panel B-1, post-merger firms have significantly higher productivity per employee than pre-merger firms in the PwC and DTT events at the 1% level (t = 4.500 and 4.199; Z = 4.450 and 3.270). For the PwC and DTT events, mergers lead to increased concentration and superior operating performance in both net profit per partner and productivity per employee, but produce an immaterial change in the profit ratio. Combining two firms achieves substantial savings in fixed costs for knowledge sharing and support personnel. This is because a post-merger firm may be better able to exploit opportunities and generate additional revenues because of its size, professional skills, and experience. The synergy between these factors may reduce costs, increase revenues, and create economies of scale (Banker et al., 2003).

Additional Test and Managerial Implications

Additional Test of Audit Market Concentration: To provide audit services to public companies in Taiwan, public accounting firms must have more than two partners to practice. In the previous sample classification, both big and large firms meet this requirement. Tyranski (2008) states that while the reputation of Big 4 firms remains strong in the U.S. audit market, many large and medium firms continue to gain market share among public companies. To some extent, big and large firms in effect compete for clients in the same market (Elder, Beasley, and Arens, 2008). Hence, this study groups the big and large firms into a category called public firms for additional analysis.

Table 7: Comparisons of Performance between Pre- and Post-Merger Public Accounting Firms

Annual Performance of The Pre- and Post-Merger Public Accounting Firms						
Year	PwC event			DTT event		
	Net profit Per Partner	Profit Ratio	Productivity Per Employee	Net profit Per Partner	Profit Ratio	Productivity Per Employee
1992	$4,960,386	0.231	$1,068,122	$5,817,561	0.269	$1,104,140
	6,183,084	0.258	1,106,578	4,989,567	0.218	1,282,850
	5,296,628	0.245	1,079,292	5,281,800	0.244	1,215,893
1993						
	4,249,714	0.217	1,026,354	5,118,411	0.276	1,007,275
	2,935,850	0.143	973,783	5,324,113	0.177	1,660,700
	3,858,350	0.180	1,009,562	5,236,499	0.227	1,379,175
1994						
	6,112,756	0.225	1,353,847	5,494,696	0.277	1,250,062
	3,751,976	0.201	1,240,609	3,479,095	0.122	1,744,499
	5,188,972	0.213	1,317,045	4,342,924	0.199	1,536,481
1995						
	5,429,694	0.198	1,160,871	5,656,901	0.223	1,367,214
	4,983,434	0.264	1,224,082	5,898,877	0.175	1,787,767
	5,251,190	0.231	1,180,064	5,803,984	0.199	1,624,464
1996						
	6,032,793	0.281	1,048,273	5,933,573	0.176	1,767,183
	6,560,655	0.245	1,308,749	3,171,739	0.187	1,165,726
	6,371,421	0.263	1,215,165	4,591,015	0.182	1,515,011
1997						
	7,197,075	0.227	1,268,167	5,232,794	0.275	1,126,407
	8,583,860	0.329	1,181,675	7,332,175	0.222	1,793,716
	7,746,556	0.278	1,236,394	6,296,864	0.248	1,479,231
1998						
	—	—	—	5,697,949	0.180	1,750,797
	—	—	—	4,631,143	0.208	1,120,871
	5,277,024	0.202	1,387,415	5,195,923	0.194	1,440,177
1999						
	—	—	—	5,834,061	0.215	1,795,315
	—	—		5,338,826	0.249	1,283,270
	7,644,118	0.257	1,418,019	5,602,059	0.232	1,542,203
2000						
	—	—	—	7,425,935	0.296	1,586,026
	—	—	—	7,577,200	0.230	2,162,901
	7,093,224	0.217	1,476,282	7,509,971	0.263	1,901,168
2001						
	—	—	—	9,491,483	0.269	2,575,964
	—	—	—	4,586,726	0.195	1,777,322
	7,837,157	0.213	2,030,910	7,887,729	0.232	2,230,435
2002						
	—	—	—	7,374,727	0.205	2,657,304
	—	—	—	8,134,741	0.287	1,589,983
	8,117,745	0.231	2,197,096	7,713,907	0.246	2,107,851
2003	8,089,860	0.243	2,157,043	9,189,091	0.268	2,473,793
2004	8,101,415	0.252	3,412,142	8,136,970	0.296	2,638,740
2005	7,881,657	0.239	1,730,970	6,902,613	0.267	2,096,446
2006	5,682,412	0.149	1,722,005	5,180,018	0.157	2,194,549
2007	9,578,168	0.231	1,866,581	6,731,827	0.158	2,289,222
2008	8,959,812	0.216	1,951,902	7,429,128	0.158	2,276,921

Table 7 displays annual operating performance for the pre- and post-merger public accounting firms. It list nine (three) figures for the pre-merger (post-merger) years for both PwC and DTT events. In the pre-merger years in each column, the first two figures are shown for individual firms and the last figure for their weighted average measure. After merger, the three figures are for the merged firm. Measures of performance include net profit per partner, profit ratio, and productivity per employee.. Both net profit per partner and productivity per employee are expressed in new Taiwan dollars.

Table 8: Differences in Operating Performance between Pre and Post-Merger Accounting Firms

	Post-Merger Mean (Median)	Pre-Merger Mean (Median)	Difference	t-statistic (Z-statistic)
Panel 1 Net Profit Per Partner				
PwC	7,660,236	5,594,133	2,066,103	3.932***
	(7,881,657)	(5,363,161)	(2,518,496)	(3.191***)
New DT	7,261,608	5,909,241	1,352,366	2.192**
	(7,165,871)	(5,656,901)	(1,508,970)	(1.868**)
Panel 2 Profit Ratio				
PwC	0.223	0.236	-0.013	-0.875
	(0.231)	(0.231)	(0.001)	(-0.860)
New DT	0.220	0.225	-0.005	-0.194
	(0.215)	(0.220)	(-0.005)	(-0.371)
Panel 3 Productivity Per Employee				
PwC	1,940,942	1,166,591	774,352	4.500***
	(1,866,581)	(1,180,870)	(685,711)	(4.450***)
New DT	2,328,279	1,616,042	712,237	4.199***
	(2,283,072)	(1,586,026)	(697,046)	(3.270***)

*Table 8 shows the testing results of difference in operating performance between pre- and post-merger public accounting firms. *, **, *** Denote two-tailed significance at the 10 %, 5 % and 1 % levels. Both net profit per partner and productivity per employee are expressed in new Taiwan dollars.*

After calculating the standardized HHIs for the newly established sub-sample, this study compares the HHI differences between the three sub-samples: public firms, medium firms, and small firms. The differences in standardized HHIs and operating performance for the three sub-samples are then statistically tested in the four practice sub-markets. The un-tabulated empirical results are similar to those reported in Panel B of Tables 3 through 6. Public firms have the lowest concentration and highest competition levels for the three sub-samples and in the four practice sub-markets. Public firms perform the best in net profit per partner, profit ratio, and productivity per employee.

Managerial Implications of the Results

The empirical results of this study show that the auditing and accounting sub-markets are the most competitive in the four sub-markets, while the tax and consultation sub-markets are the least competitive. This study defines tax practices as tax planning, administrative remedy of internal taxation, and other tax operation services. Practically, tax and consultation practices are often referred to as a broadly defined management advisory service (MAS). For the past few decades, traditional practice market, such as auditing and accounting sub-markets of this study, has been increasingly competitive and less lucrative for practitioners. The findings of this study suggest that practitioners may expand their services to the MAS because it creates unlimited business opportunity and has no adverse effect on their auditor independence, especially for practitioners in the medium and small firms which provide no audit services to public companies. Next, this study demonstrates that mergers between big firms increase market concentration level. However, the long-term concentration level of big firms sub-sample does not change significantly. This indicates that mergers between big firms do not adversely change market structure and post-merger firms achieve better operating performance than pre-merger ones. This suggests that practitioners take into account the mergers between public accounting firms because combining two firms leads to synergy, substantial cost savings, increased revenue, and economies of scale (Banker et al., 2003). Long-term cooperation between U.S. and Taiwanese audit industries has created a similar audit market structure in both countries. Hence, the implications of this study apply to practitioners in any countries with audit markets similar to that in the U.S.

CONCLUSIONS

Based on long-term audit fee information for all public accounting firms in Taiwan, this study presents a direct measure of market share for estimating the Herfindahl-Hirschman index (HHI). To facilitate cross-sample comparisons, this study creates a standardized HHI and obtains the following results. First, this study presents empirical results for four sub-samples: big, large, medium, and small firms. Standardized HHIs indicate variance in the long-term market competition levels for the four sub-samples. Big firms have the highest competition level, but large, medium, and small firms exhibit no significant difference in the competition level. However, the long-term level of market competition does not change significantly in the four sub-samples.

Next, this study reports findings on the long-term market structure for the four practice markets, including auditing, tax, consultation, and accounting sub-markets. The auditing and accounting sub-markets have significantly higher competition level than the tax and consultation sub-markets. Further, the long-term market structure of auditing, tax, and accounting sub-markets becomes less competitive, while that of the consultation sub-market does not change significantly. Big firms have the highest competition level in the four sub-markets. Given the highest competition level, big firms have better financial performance than the other three sub-samples. Post-merger firms outperform pre-merger firms for both the 1999 PwC and the 2003 DTT events.

The results above should be interpreted in light of the following limitations. First, this study utilizes the Herfindahl-Hirschman index (HHI) to assess audit market concentration and then use it to determine the competition level. Prior studies state that it is essential to assess the circumstances surrounding the competitive interaction of firms to make conclusions about the nature of competition, rather than rely solely on the HHI (Besanko et al., 2000). Second, this study uses a univariate test to compare the operating performance for different sub-samples and for pre- and post-merger public accounting firms. This method does not control for other factors affecting operating performance. Comparison of operating performance has important implications for practitioners, academics and regulators alike. Adopting a more rigorous method, such as a multiple regression model, on this issue constitutes a promising avenue for future studies.

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AFTERMARKET RISK AND UNDERPRICING OF INITIAL PUBLIC OFFERS IN THE ARABIAN GULF COUNTRIES: AN EMPIRICAL ANALYSIS

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ABSTRACT

The recent evidence has shown that IPO uncertainty continues in aftermarket until a normal market is established. Evidence also indicates that aftermarket risk measured by stock's beta is related to the degree of underpricing in the US market. This may imply that additional underpricing may be required to compensate for aftermarket risk, assuming that the aftermarket risk is important for the investors who want to buy primary shares from the public offers and/or aftermarket trading. Therefore, we examine the relationship between aftermarket risk and underpricing by using data from six Arabian Gulf countries, an economic region where all personal incomes including capital gains are tax-free. The evidence based on 147 samples depicts that IPO firm's aftermarket risk measured by stock's beta has significant relationship with the degree of underpricing. Thereby, it is confirmed that the relationship between aftermarket risk and underpricing also exists in the Arabian Gulf countries, an important economic region outside the US. Paper concludes that that underwriters and/or issuers may need to forecast the expected aftermarket risk while determining the offer price.

JEL: G30 and G32

KEY WORDS: IPO Underpricing, Aftermarket Risk, and Arabian Gulf Markets

INTRODUCTION

The underpricing of initial public offers (IPO) is a global phenomenon that has been investigated by numerous researchers from different dimensions of underpricing (see Loughran *et al.,* 1994 for global evidence on IPO underpricing and Jenkinson and Ljungqvist (2001) for a critical review of the vast IPO literature). Researchers offer several theories suggesting that underpricing is inevitable due to information asymmetry among the parties involved in IPO process. These include winner's curse theory (Rock, 1986); signaling theory (Allen and Faulhaber, 1989; and Grinblatt and Hwang, 1989); and price delegation theory (Baron, 1982). In addition, researchers have offered explanation of underpricing in the light of institutional factors, such as underwriter price supports and pre-issue information gathering (Ruud, 1993; and Benveniste and Spindt, 1989), need for ownership dispersion and secondary market liquidity (Brennan and Frank, 1997; Bodnaruk *et. al*., 2008; and Booth and Chua, 1996), and listing delay after offer pricing, (Chowdhry and Sherman, 1996). Explanations of underpricing also appear with respect to fads and divergence of opinion in primary market (Aggarwal and Rivoli, 1990; Miller, 1977; and Gao *et. al*., 2006). Of these theories, many of them have basically considered an ideal market condition while explaining IPO underpricing as an equilibrium phenomenon due to ex-ante uncertainty in primary market.

Plenty of empirical studies have examined IPO underpricing across the world, and found evidence of ex-ante uncertainty in the primary markets using numerous proxy variables, but recent evidence has found that market uncertainty continues in the aftermarket period until normal market is established (Chen and Wilhelm, 2008 and Falconieri, *et. al*., 2009). Evidence is also found that IPO aftermarket beta and initial underpricing are correlated in the US (Gleason, *et. al.,* 2008). These findings may imply that underpricing is needed to compensate IPO aftermarket risks in addition to ex-ante uncertainty if investors want to buy shares from the public offers and/or aftermarket trading. Therefore, the evidence of correlation between aftermarket stock's beta and underpricing needs confirmation from different markets outside the US. Hence, we examine the relationship between aftermarket risk and the level of underpricing using data from the markets of Arabian Gulf countries (e.g. Saudi Arabia, United Arab Emirates, Kuwait, Qatar, Bahrain, and Oman), an important economic region where all personal incomes and capital gains are tax-free. It is noted that none of the Arabian Gulf Countries imposes taxes on the personal incomes (including capital gains) of their local and expatriates residents.

However, the foreign companies doing business in the region usually need to pay corporate taxes at different rates depending on the nature of their business. The local companies are however exempted from the corporate taxes. In Saudi Arabia, both the individuals and corporations (except foreigners) need to pay 2.5 percent of the surplus liquid assets (not annual income) as 'Zakat', a religious obligatory donation from the rich Muslim citizens to poor citizens of the country. Hence this is not a tax on the current income. It is pertinent to note that none of the gulf countries impose taxes on the capital gains from portfolio investments. This is an important issue because capital gain is a taxable income in many countries. Hence, IPO underpricing may be needed to cover any potential tax effect in the countries where initial return is taxable (Rydqvst, 1997; Guenther; Willenborg, 1999; and Riano et al, 2007). Thus Arabian Gulf markets are different from the other markets.

The findings on these markets will give us a better understanding about the relationship between aftermarket risk and underpricing of initial public offers, particularly whether such relationship prevails in the different markets. We hypothesized that positive relationship between the aftermarket risks and underpricing may exist in the Arabian Gulf markets, because uncertainty about IPO value may not be fully resolved until a normal market is established after listing. Hamada's (1972) theory also justifies a positive relationship between the aftermarket risk and underpricing, because the aftermarket beta may have a link to the firm's business and financial risks that have been inherited from the period before listing.The study based on 147 IPOs showed that average IPO underpricing in the Arabian Gulf markets is about 250 percent, suggesting high risk in the primary markets of this region. The finding is consistent with the previous finding based on 47 Arabian Gulf IPOs that is reported by Al-Hassan, *et al*. (2010). It is found underpricing is significantly correlated with the aftermarket risk factor (stock's beta), supporting the prior findings in the US market that is reported by Gleason, *et. al.* (2008). While the direction of relationship (causality) between the aftermarket risks and underpricing is yet to be thoroughly investigated, the regression models that include beta as an explanatory variable with underpricing as dependant variable have more explanatory power than that of the models which include beta as dependant variable with underpricing as an explanatory variable. Given the above results, it is suggested that underwriters and/or issuers may need to forecast the expected aftermarket risk while determining the offer price, and thereby, underpricing may reflect the level of aftermarket risk.

The rest of the paper is organized as follows: in the next section, the relevant literature and main hypothesis is discussed. In the subsequent sections, we describe respectively the research methodology, sample characteristics and empirical findings. Finally, the conclusion is given in the last section

LITERATURE REVIEW AND HYPOTHESIS

A common premise of major IPO theories noted earlier is that the underwriters and issuers have difficulties in assessing fair value of shares to be offered, which is known as IPO ex ante uncertainty. A review on IPO studies presented by Jenkinson and Ljungqvist (2001) overwhelmingly shows that higher IPO ex-ante uncertainty results in higher initial returns across the countries. These studies used different proxy variables (e.g. underwriter reputation, offer size, ownership retention by directors, industry classification etc.) to examine the effect of ex-ante uncertainty on IPO underpricing. The issuers and underwriters face uncertainty of new issue valuation before its listing on the stock exchange, but efforts are usually made to set a best possible offer price by book building process. There are questions here as to whether it is possible to correctly estimate the degree of ex-ante uncertainty, or uncertainty remains after IPO listing and if the price discount offered through underpricing of IPOs can justify the level of uncertainty at the time of and after the IPO. In this regard, literature shows that the degree of uncertainty and information asymmetry among investors is not completely resolved prior to secondary market trading, despite underwriter's book building efforts. Indeed, uncertainty is found to be significantly prevalent in early aftermarket period until a normal market condition appeared (Chen and Wilhelm, 2008 and Falconieri, *et. al*., 2009).

In the mainstream literature, the effect of IPO ex-ante uncertainty on underpricing is well known from the findings of numerous studies that used different ex-ante uncertainty proxy variables. Earlier studies however recognized that uncertainty in the primary market should be measured based on the risks which are priced in the market, but they did not consider the stock's beta as a suitable proxy for it by arguing that the ex-ante risk is different from the stock's beta (Johnson and Miller, 1988; Beatty and Ritter, 1986; and Ritter, 1984). While a direct measure of risk for ex-ante uncertainty is difficult to find, some researchers used aftermarket standard deviation as ex-post measure of IPO ex-ante uncertainty (e.g., Ritter 1984, Beatty and Ritter, 1986). They found a strong relationship between the aftermarket return standard deviation and IPO initial return. The relationship between the aftermarket volatility and IPO initial returns also appeared in recent literature (Jog and Wang, 2009 and Pettway *et. al.* 2008). These findings seemed to be consistent with the above discussion that the degree of uncertainty is not completely resolved prior to secondary market trading, but rather it prevails in the aftermarket period.

According to asset pricing theory 'standard deviation' measures the total risk of investment that consists of 'unsystematic' and 'systematic' components. Beneda and Zhang (2009) found a negative relationship between the initial unsystematic volatility level and post-IPO volatility change. Their results showed a significant positive association of underpricing with the initial systematic variance, but insignificant association is observed between the initial returns and unsystematic variance in the aftermarket period. The finding suggests that the aftermarket standard deviation would not be very relevant in determining IPO underpricing though it measures the total risk of newly listed stocks. If uncertainty is the reason for underpricing, then underwriter/issuers should consider the risk element that is priced in the market. Since trading price data is not available in primary markets, it is not possible to estimate market beta that is priced in the market. An earlier study by Almisher and Kish (2000) used 'accounting beta' as a proxy for IPO ex-ante uncertainty by arguing that there is an association between the market and accounting betas.

They found that accounting beta of IPO companies have significant positive effect on the level of underpricing giving initial support to the idea that IPO uncertainty proxy should accommodate the risk that is priced in the market. However, Beatty and Ritter (1986) argued that the ex-ante uncertainty which causes IPO underpricing does not conform to the notion of systematic risk (beta) due to limitation in diversifying the IPO investments by uninformed investors. Therefore, earlier studies did not consider the 'beta' as a useful proxy for measuring IPO risk on this theoretical ground. As a result of this and because of the difficulty in measuring the beta coefficient for IPO companies prior to their listing, researchers paid less attention on this issue. Recently, Gleason, *et. al,* (2008) documented using US data that aftermarket risk following IPO is higher for the firms that experienced a higher level of underpricing. This finding motivates us to revisit the beta issue and examine it in a different location. Gleason, *et. al.,* 2008 analyzed that underpricing not only reflects the uncertainty at the time of the offering, but also is useful indicator of aftermarket risk. We can also analyze the evidence of positive relationship between the underpricing and aftermarket risk in the light of Hamada's (1972) theory, which suggests that a security's market beta is due to the business and financial risks of the corporate firm. If an IPO firm inherits such business and financial risks from the prelisting period, then beta may display relationship with the initial underpricing.

Based on the above analyses, we draw test hypothesis as follows:

H_0: There is no positive relationship between IPO firm's Aftermarket risk and initial return.

H_A: There is positive relationship between IPO firm's Aftermarket risk and initial return.

If IPO firm's aftermarket beta has positive relationship with the level of underpricing then it can be suggested that that the level of underpricing needs to justify for market's uncertainty both at the time of and after IPO listing until the normal market is established over the long run. This is particularly important for the investors who want to retain IPO stocks in the aftermarket period. The above hypotheses have been examined using data from the markets of Arabian Gulf region. This helps us confirming whether the relationship between the IPO aftermarket risk and underpricing also exists in different economies outside US where personal incomes and capital gains are not taxable.

METHODOLOGY

The study applied regression method to examine the IPO underpricing levels in Arabian Gulf markets. The dependent variable, level of underpricing, is estimated on the day of listing using follows measures.

$$MAIR_1 = \frac{1}{N}\sum_{i=1}^{N} IR_i - R_m \tag{1}$$

Where, $MAIR_1$ is the market-adjusted average initial return on the 1st day of IPO listing. IR_i is the initial return for IPO *i* calculated as $(P_{it} - P_{io})/P_{io}$ from the day of public offer to the close of the first trading day following listing. P_{it} is the closing market price of IPO *i* on the 1st trading day and R_m is the market return calculated for the same period, using the respective market's value-weighted price index. In a similar manner, underpricing is also estimated on the 30th day after listing in order to examine the consistency of results. This is denoted as $MAIR_{30}$ and estimated as under:

$$MAIR_{30} = \frac{1}{N}\sum_{i=1}^{N} IR_i - R_m \tag{2}$$

Where, $MAIR_{30}$ is the market-adjusted average initial return on the 30th day of IPO listing. IR_i is the initial return for IPO *i* calculated as $(P_{it} - P_{io})/P_{io}$ from the day of public offer to the close of the 30th trading day following listing. P_{it} is the closing market price of IPO *i* on the 30th trading day and R_m is the market return calculated for the same period, using the respective market's value-weighted price index.

Following explanatory variables are selected based on the mainstream IPO literature, our observation on Arabian Gulf markets, and data availability.

Beta: this is the *ex-post* measure of IPO's market risk factor calculated by using market model time series regression with weekly returns over 52 weeks (one year) following IPO listing. The beta estimate is adjusted for thin-trading effect by using Dimson (1979) method with three lag and lead returns. This variable has emerged from the background discussion of this paper that leads to our test hypothesis.

Log_Size: this is natural log of the size of public offer measured by the gross issue proceeds, which is calculated as the number of shares in offer times the offer price. This variable is used in literature as proxy measure of IPO ex-ante uncertainty. For example, Beatty and Ritter (1986), Clarkson and Merkley (1994), McGuiness (1992) are among others.

Log_Age: this is the natural log of the number of days between the date of company incorporation and the date of listing. This variable is used in literature as a proxy measure of IPO ex-ante uncertainty. David (2002), Kirkulak and Davis (2005), Islam, et. al., (2010) are among others.

Urp: this is the measure of issue underwriters' reputation estimated by examining the tombstones placed in the financial section of local newspapers. We examined how many times a particular investment banker worked as the lead/co-lead underwriter since its incorporation. The higher the frequency as lead or co-lead underwriter the higher is the reputation. The frequency has been divided by the age (years) of the banker in order to get adjusted reputation score. After preparing an adjusted score table we have classified them into four reputation class (Outstanding, Very High, High, and Low). This variable is widely used by researchers as proxy for IPO uncertainty. For example, Beatty and Ritter, Cater and Manaster (1990) and Johnson and Miller (1988) among many others.

Retn: this variable measures the percentage of total equity retained by the existing owners after share floatation. This variable is suggested by Leland and Pyle (1977) and Grinblatt and Hwang (1989) to examine signaling hypothesis of IPO underpricing and is found significant in many empirical studies.

Sub: this is the level of IPO subscription on the day of public offer, which measures the market demand of IPO shares. It has been argued that IPO initial return usually depends on the level of market demand and many studies have found significantly positive relationship between the level of initial return and the degree of oversubscription. Among the most recent works, McGuinnes's (2009) paper is notable.

Fin: this is a dichotomous variable. FIN = 1 if IPO is listed under Financial Services category, else FIN=0. About 36 percent of the sample includes IPOs listed under Financial Services sector. Therefore, it is needed to control for sample bias (if any).

Crisis: this is a dichotomous variable. CRISIS = 1 if IPO is listed during recent global financial crisis period (period after July 2007), else CRISIS = 0. We observed that IPO activities and marker performance in the Gulf region are very sluggish. Therefore, it is needed to control for abnormal sample period.

A number of regression tests are conducted using the above variables in order to examine the relationship between the levels of underpricing and IPO firms' aftermarket risk. The first set of test models specified to measure the marginal effect of aftermarket risk on the level of initial underpricing as follows:

$$MAIR_{i1} = \alpha + \beta_1 BETA_i + \beta_2 LOGSIZE_i + \beta_3 LOGAGE^i + \beta_4 URP_i + \beta_5 RETN_i + \beta_6 SUB_i \quad \beta_7 FIN_i + \beta_8 CRISIS_i + \varepsilon_i \tag{3}$$

This model is estimated by using full sample and sub-sample data and the relevant results are presented in Table 3 and Table 4 respectively. The consistency of the results is also tested by replacing the $MAIR_{i1}$ with $MAIR_{i30}$. Since there is no confirmed prior knowledge about the direction of relationship between IPO underpricing and aftermarket risk, we tested another model as follows.

$$BETA_i = \alpha + \beta_1 MAIR_{i1} + \beta_2 LOGSIZE_i + \beta_3 LOGAGE^i + \beta_4 URP_i + \beta_5 RETN_i + \beta_6 SUB_i \quad \beta_7 FIN_i + \beta_8 CRISIS_i + \varepsilon_i \tag{4}$$

This model is also estimated by using full sample and sub-sample data and the relevant results are presented in Table 5 and Table 6 respectively

Samples and Data

The sample includes 147 IPOs listed on seven stock markets of six Arabian Gulf countries during the period from January 2003 to February 2010, which covers about 21.45 percent of all securities listed on these markets as on February 2010. The sample characteristics showed that 53 IPOs are listed under Financial Services sector, which comprises about 36 percent of the sample. This is followed by 17 IPOs (12 percent) listed under Oil and Gas sector, 16 (11 percent) under Industrial Manufacturing sector, 12 (eight percent) under Real Estate sector and 9 (six percent) in respectively construction and transport sectors. Overall it shows that samples are somewhat dominated by the IPOs listed under Financial Services sector, through total sample has representation from all the major industrial sectors in the Arabian Gulf countries.

Table 1: Industry Distribution of IPOS Across the Six Arabian Gulf Countries over the Period from 2003 to 2010

Industry Sector	Saudi Arabia	UAE	Kuwait	Qatar	Oman	Bahrain	Total
Agriculture		1					1
Construction	1	5	1		1	1	9
Consumer Goods	1					1	2
Education	1						1
Financial Services	30	12	2	4	2	3	53
Food and Beverages	4	1		1			6
Health Care	2						2
Industrial Manufacturing	8	2	2	1	1	2	16
Leisure and Tourism	1					1	2
Media	1						1
Mining and Metals	1						1
Oil and Gas	9	2	2	3	1		17
Power and Utilities		1			4		5
Real Estates	3	6		1		2	12
Retail	2						2
Services	1						1
Telecommunications	4	1		1	1		7
Transport	1	4	3	1			9
Total	70	35	10	12	10	10	147

A total of 685 companies listed on the exchanges of six Arabian Gulf countries as of February 2010. Therefore, the sample size covers about 21.45 percent of the total market size in terms of number of listings while it covers about 35 percent of the market in terms of total capitalization value. About 48 percent of total sample is from Saudi Arabian market, which is the largest market in the region with a total listing of 140 companies. Although Kuwait is the second largest market in the region we are able to get only 10 IPO data from the database

The country distribution presented in Table 1 shows that a total 70 IPOs belong to Saudi Arabia, which covers about 48 percent of the sample. This is followed by 35 IPOs from United Arab Emirates, comprising about 24 percent of the sample. The remaining 42 IPOs are from four countries: Kuwait, Bahrain, Qatar, and Oman that accounts for the remaining 28 percent of the sample. The general distribution of sample indicates that a major fraction of the samples are from the Saudi Arabian market which is the largest market (in terms of the market capitalization and number of securities) with in the region. Although Kuwait is the second largest market, we are unable to get enough samples for this market from the database. For other markets the respective samples covers 20 to 25 percent of their total market size. The required data for the samples are extracted from various sources that includes the databases of the respective stock markets, Gulf base, and Sharjah University's digital library.

EMPIRICAL FINDINGS

Results presented in Panel A of Table 2 show that average underpricing, measured by market adjusted initial return (MAIR), in the Arabian Gulf region is over 250 percent on the day of listing. The level of underpricing appeared to be over 248 percent if MAIR is computed on the 30th day after listing. Both the measures of underpricing, $MAIR_1$ and $MAIR_{30}$, provide the evidence of very high IPO underpricing in the Arabian Gulf markets. This new finding is comparable with the similar high underpricing found in China (see Loughran, *et. al*, 1994 for IPO underpricing in different international markets). Also it is consistent with the another finding based on 47 IPOs from the same region that is reported by Al-Hassan, *et al*. (2010) While average underpricing is very high in the Arabian Gulf region, wide variations are also observed from the high average standard deviation (about 300 percent).

Table 2: Characteristics of IPO Underpricing in Arabian Gulf Countries over the Period from 2003 to 2009

Panel A: Descriptive Statistics for Underpricing Across the Arabian Gulf Region		
	$MAIR_1$	**$MAIR_{30}$**
Average	250.17	248.26
Standard Deviation	297.07	332.77
Median	133.57	125.42
Maximum	1,430.00	1,720.00
Minimum	-30.00	-37.00
Panel B: IPO Underpricing in Different Industrial Sectors		
Industry Sectors	**$MAIR_1$**	**$MAIR_{30}$**
Agriculture	0.01	1.70
Construction	89.50	87.00
Consumer Goods	130.40	118.55
Education	178.46	151.92
Financial Services	394.86	415.33
Food and Beverages	157.44	134.20
Health Care	15.33	7.80
Industrial Manufacturing	117.50	62.47
Leisure and Tourism	-5.25	-13.62
Media	106.50	57.61
Mining and Metals	52.50	28.75
Oil and Gas	216.56	221.69
Power and Utilities	168.93	124.26
Real Estates	226.03	245.86
Retail	31.48	-3.01
Services	126.14	113.64
Telecommunications	133.59	137.31
Transport	175.44	113.32
Panel B: Level of Underpricing in Different Arabian Gulf Countries		
	$MAIR_1$	**$MAIR_{30}$**
Saudi Arabia	283.89	297.80
United Arab Emirates	287.62	263.51
Kuwait	246.13	224.90
Qatar	264.97	248.68
Oman	57.84	41.93
Bahrain	21.40	8.09

MAIR stands for market adjusted initial return that measures the level of IPO underpricing after adjustment of market return over the period from the day of offer opening to the listing day (or 30 day after listing). We used respective market's value-weighted price index to calculate market returns. $MAIR_1$ measures the level of underpricing on the listing and $MAIR_{30}$ measures the underpricing on 30-day after listing.

The underpricing levels of different industrial sectors presented in the Panel B of Table 2 show that underpricing level in the Financial Services sector is around 400 percent ($MAIR_1$ is 394 percent and $MAIR_{30}$ is 415 percent) which is far larger than the level of underpricing in other sectors. This result seems to reflect the effects of regional financial sectors boom during the recent past period. Since 2003 a total of 53 IPOs have been floated in the market comprising nearly 36 percent of the sample. Among other industrial sectors, the average underpricings of Real Estate and Oil & Gas sectors are found to be 226 and 217 percents respectively.

In other sectors, the average underpricing varies from 100 to 200 percent with a few exceptions. Results also depict variations of underpricing across the different markets of Arabian Gulf area. The Panel C of Table 2 shows that average underpricing of two smaller markets, Bahrain (21.40 percent) and Oman (57.84 percent), are much lower than those of the other four markets.

Having examined the characteristics of IPO underpricing, we have tested ten regression models to determine the relationship between the level of underpricing and the aftermarket risk (beta) of the newly listed stocks. Models 1 and 4 applied univariate tests using the aftermarket risk (IPO stock's aftermarket beta) as the sole explanatory variable. Results show that BETA significantly affect the IPO initial underprcing measured on the day of listing ($MAIR_1$) as well as that is measured on the 30_{th} day after listing ($MAIR_{30}$). It is also found that BETA alone can explain about 6 percent of $MAIR_1$ and 5 percent of $MAIR_{30}$. Model 2 employs multivariate tests by including all the selected explanatory variables to measure their effect on the level of initial underpricing measure by $MAIR_1$.

The multivariate test results of Model 2 that presented in Table 3 show that BETA positively affect the $MAIR_1$ and the coefficient is significant at 5 percent level. Among other variables, the offer size (LOG_SIZE), age of IPO companies (LOG_AGE), and underwriter reputation (URP) have significant negative relationship with level of initial underpricing. These variables could explain about 29 percent of initial underpricing in the Arabian Gulf markets. Results of Model 5 also show that same variables significantly affect the level of underpricing measured on the 30^{th} day after IPO listing, and all variables together explain about 31 percent of $MAIR_{30}$.

The Models 3 and 6 used step-wise regression method to determine the most relevant variables to explain the underpricing level on the day of listing as well as on the 30^{th} day after IPO listing. It is found that all the significant variables of Models 2 and 5 remain significant as before, and additionally the issue subscription appears to be significant. The adjusted R^2 increases to 35 and 36 percent respectively for Model 3 and Model 6. Overall the results of Models 1 through 6 demonstrated that the IPO aftermarket risk measured by stock's beta has significant positive relationship with the level of initial undepricing, and thereby the null hypothesis is rejected.

The consistency of results is tested by estimating another set of four regressions with different sub samples, and results are presented in Table 4. The results of Model 5 that used 70 Saudi Arabian IPO data showed that BETA is significant at one percent level. Among the other variables, LOG_SIZE, LOG_AGE, URP, and FIN are significant at five and ten percent levels. The results of Model 6 that used 35 UAE IPO data showed that BETA is significant at ten percent level. Among the other variables, LOG_SIZE, LOG_AGE, URP, and CRISIS are significant. The BETA variable also appears to be significant in the Model 7 that uses 77 IPO data from the Arabian Gulf region except Saudi Arabia. Same results are also revealed in Model 8 that used IPO data from all countries except Saudi Arabia and UAE.

Table 3: Full sample regression results with MAIR as Dependent variable (N = 147)

	$MAIR_1$ as Dependent Variable			$MAIR_{30}$ as Dependent Variable		
Explanatory Variables	**Model-1**	**Model -2**	**Model -3**	**Model -4**	**Model- 5:**	**Model- 6:**
constant	1.44 (2.35)**	7.18 (5.10)***	7.25 (7.49)***	1.66 (2.36)**	8.04 (5.06)***	8.13 (6.35)***
beta	1.04 (1.97)**	1.13 (2.05)**	1.19 (2.23)**	0.78 (1.72)*	0.94 (2.35)**	1.13 (2.46)**
log_size		-1.25 (-3.50)***	-1.42 (-4.45)***		-1.61 (-4.01)***	-1.44 (-3.70)***
log_age		-0.80 (-3.60)***	-0.90 (-4.37)***		-0.91 (-3.67)***	-0.90 (-3.63)***
urp		-0.35 (-2.05)**	-0.35 (-2.09)**		-0.41 (-2.12)**	-0.40 (-2.10)**
retn		-0.011 (-0.81)			-0.001 (-0.08)	
sub		0.004 (1.58)	0.004 (1.91)*		0.004 (1.53)	0.005 (2.02)**
fin		0.62 (1.20)			0.78 (1.34)	
crisis		-0.24 (-0.56)			-0.62 (-1.26)	
f value	3.77**	7.6***	11.70***	2.96*	8.28***	12.44***
adjusted r^2	0.06	0.29	0.35	0.05	0.31	0.36

*The dependent variable $MAIR_1$ is the market adjusted initial return on the 1st day of IPO listing, and $MAIR_{30}$ measures the market adjusted initial return on the 30th day after listing. Among the explanatory variables, BETA is the measure of IPO stock's market risk factor calculated over the one year in aftermarket period. LOG_SIZE is the natural log of the offer size measured by the number of shares in offer times the offer price. LOG_AGE is the natural log of the age of the company measured by the time period between the day of incorporation and the day of exchange listing. URP is the measure of underwriters' reputation. RETN is that percent of ownership retained by the existing directors following share floatation. SUB is that issue subscription times relative to the offer size. FIN is a dichotomous variable where FIN = 1 if IPO is listed in Financial sector, else FIN = 0. CRISIS is a dichotomous variable where CRISIS =1 if IPO is listed during financial crisis period (after July 2007) else CRISIS = 0. Numbers in parentheses shows the t-values of respective co-efficient. Asterisks ***, **, and * denote the level of significance at respectively 1, 5, and 10 percents.*

Finally, results presented in Table 4 depicts that BETA is consistently significant in all markets, while some variation occurs with respect to other variables. For example, CRISIS variable is not significant in Saudi Arabia but it is highly significant in UAE and other regional markets. This is indeed an interesting to find that the global financial crisis did not affect the Saudi Arabian market while it significantly affected the UAE market. It suggests that Saudi Arabian market is somewhat insulated from global economic upheavals.

The ownership retention by the founding directors is not significant in the Saudi Arabia and UAE, suggesting that IPO signaling hypothesis is not very useful in these markets. In addition, LOG-SIZE and URP variables are only significant in Saudi Arabia and UAE. Since BETA is consistently significant in all the six countries, we can accept the alternative hypothesis that suggests a positive relationship between the initial underpricing and aftermarket risk.

Therefore, this study confirms that aftermarket risk is positively related to IPO underpricing in the non-tax economies of Arabian Gulf countries that provides support to similar findings by Gleason, *et. al,* (2008) in the US market. Finally, findings suggest that underwriters and/or issuers may need to consider IPO aftermarket risk in determining the offer price. This suggestion is pertinent because aftermarket risk may be important for the investors who want to purchase IPO shares from public offers and/or from aftermarket trading.

Table 4: Sub-sample Regression Results with $MAIR_1$ as Dependent Variable

Explanatory Variables	Saudi Arabia (n = 70)	United Arab Emirates (UAE) (n = 35)	Arabian Gulf Region Excluding Saudi Arabia (n = 77)	Arabian Gulf Excluding Saudi Arabia and UAE (n = 42)
	model 7	model 8	model 9	model 10
constant	1.99 (0.51)	7.28 (2.64)***	7.05 (3.74)***	7.54 (3.87)***
beta	1.78 (2.70)***	0.23 (1.88)*	0.42 (2.00)**	1.46 (1.99)**
log_size	-1.36 (-2.51)**	-3.33 (-2.72)***	-0.14 (-0.24)	-0.23 (-0.42)
log_age	-0.65 (-2.31)**	-0.24 (1.87)*	-0.86 (-2.32)**	-1.04 (-2.42)**
urp	-0.16 (-1.66)*	-0.22 (-2.12)**	-0.18 (-0.72)	-0.30 (-1.41)
retn	0.05 (1.35)	-0.037 (-020)	-0.01 (-2.65)***	-0.01 (-3.49)***
sub	-0.00 (-0.014)	0.05 (2.44)**	0.01 (1.88)*	-0.00 (-0.86)
fin	1.68 (1.72)*	-0.86 (-1.23)	0.32 (0.50)	0.30 (0.44)
crisis	0.18 (0.33)	-2.11 (4.77)***	-1.34 (-1.97)**	-0.88 (-1.33)
f value	7.97***	8.78***	3.69***	3.49***
adjusted r^2	0.451	0.39	0.20	0.37

*The dependent variable $MAIR_1$ is the market adjusted initial return on the 1st day of IPO listing, and among the explanatory variables, BETA is the measure of IPO stock's market risk factor calculated over the one year in aftermarket period. LOG_SIZE is the natural log of the offer size measured by the number of shares in offer times the offer price. LOG_AGE is the natural log of the age of the company measured by the time period between the day of incorporation and the day of exchange listing. URP is the measure of underwriters' reputation. RETN is that percent of ownership retained by the existing directors following share floatation. SUB is that issue subscription times relative to the offer size. FIN is a dichotomous variable where FIN = 1 if IPO is listed in Financial sector, else FIN = 0. CRISIS is a dichotomous variable where CRISIS =1 if IPO is listed during financial crisis period (after July 2007) else CRISIS = 0. Numbers in parentheses shows the t-values of respective co-efficient. Asterisks ***, **, and * denote the level of significance at respectively 1, 5, and 10 percents.*

The above findings by using BETA as explanatory variable may be criticized as only 'prima facie' evidence, because the direction of relationship (causality) between the aftermarket risk and underpricing needs separate in-depth study. However, this study assumes that the IPO aftermarket risk may have effect on the level of underpricing. This is because prior evidence has shown that market uncertainty for new stocks does not fully resolve on the day of listing, and in the light of Hamada's (1972) theory, the IPO firm's inherent business and financial risk may have a connection to the aftermarket stock's beta. Hence, the level of aftermarket risk may affect the underpricing.

Alternatively, it can also be analyzed that underpricing can be a useful indicator of the aftermarket risk as suggested by Gleason, *et. al*. (2008). Therefore, we re-run the multivariate test models by using aftermarket risk (BETA) as the dependant variable and underpricing (MAIR) as the explanatory variable. The selected ex-ante uncertainty proxy variables are also included in new regression models as the control variables, because the ex-ante uncertainty may also influence the aftermarket risk.

The results of the models 11 and 12 presented in Table 5 depict that $MAIR_1$ LOG_SIZE, and FIN variables affect the BETA while the other variables are insignificant. The explanatory power of these new models appeared to be only 9.5 and 10 percents respectively, which are far lower than those of the models 2 and 3 in Table 3 that used BETA as explanatory variable.

The results of the models 13 and 14 that used underpricing at the 30th day after listing ($MAIR_{30}$) also reveals similar findings. In addition, the sub-sample results of the models 15 through 18 presented in

Table 6 depict that initial underpricing consistently affect the aftermarket risk (BETA) in all the Arabian Gulf markets, although the explanatory powers of these models are far lower than those of the models 7 through 10 in Table 4 that used BETA as explanatory variables.

Table 5: Full sample regression results with BETA as Dependent variable (N = 147)

Explanatory Variables	Model-11 All Variables Except $MAIR_{30}$	Model -12 Most Relevant Variables From Model 11	Model -13 All Variables Except $MAIR_1$	Model -14 Most Relevant Variables From Model 13
constant	0.17 (0.68)	0.38 $(2.75)^{**}$	0.23 (0.91)	0.40 (2.84)
$mair_1$	0.0003 $(2.05)^{**}$	0.0003 $(2.19)^{**}$		
$mair_{30}$			0.0002 $(2.35)^{**}$	0.0002 $(2.41)^{**}$
log_size	0.20 $(3.44)^{***}$	0.21 $(3.73)^{***}$	0.20 $(3.32)^{***}$	0.21 $(3.63)^{***}$
log_age	0.01 (0.35)		0.01 (0.20)	
urp	0.02 (0.62)		0.02 (0.55)	
retn	0.002 (0.89)		0.001 (0.76)	
sub	0.0006 (1.53)		0.0006 (1.63)	
fin	0.15 $(1.73)^{*}$	0.15 $(1.90)^{*}$	0.15 $(1.78)^{*}$	0.16 $(2.03)^{**}$
crisis	0.06 (0.81)		0.06 (0.88)	
f value	2.72^{***}	5.87^{***}	2.45^{**}	5.15^{***}
adjusted r^2	0.095	0.10	0.082	0.09

*The dependent variable BETA is the measure of IPO stock's market risk factor calculated over the one year in aftermarket period. Among the explanatory variables, $MAIR_1$ is the market adjusted initial return on the 1^{st} day of IPO listing, and $MAIR_{30}$ measures the market adjusted initial return on the 30^{th} day after listing LOG_SIZE is the natural log of the offer size measured by the number of shares in offer times the offer price. LOG_AGE is the natural log of the age of the company measured by the time period between the day of incorporation and the day of exchange listing. URP is the measure of underwriters' reputation. RETN is that percent of ownership retained by the existing directors following share floatation. SUB is that issue subscription times relative to the offer size. FIN is a dichotomous variable where FIN = 1 if IPO is listed in Financial sector, else FIN = 0. CRISIS is a dichotomous variable where CRISIS =1 if IPO is listed during financial crisis period (after July 2007) else CRISIS = 0. Numbers in parentheses shows the t-values of respective co-efficient. Asterisks ***, **, and * denote the level of significance at respectively 1, 5, and 10 percents*

Therefore, the suggestion that initial underpricing may be a useful predictor of aftermarket risk is also supported by our evidence, although the explanatory powers of the test models are lower. Set aside the direction of the relationship, it is confirmed that aftermarket risk and underpricing are correlated in the Arabian Gulf markets. This may imply that IPO underpricing may be needed to compensate for the aftermarket risk in addition to ex-ante uncertainty, and thereby the aftermarket stock beta is higher for the firms that experience a higher level of underpricing - or vice versa.

CONCLUSION

IPO underpricing is known as inevitable because of ex ante uncertainty about the fair valuation of new issue. There is abundance of empirical studies supporting the existence of IPO ex ante uncertainty in different markets across the world. However, some evidence shows that uncertainty also prevails in the aftermarket period - particularly, over the short term period. Recently, it is also documented with the US data that aftermarket risk measured by stock's beta is positively related to the level of underpricing. In this paper, we have examined whether such relationship between the IPO aftermarket risk and underpricing also exist in different markets outside the US.

Table 6: Sub-Sample Regression Results with BETA as Dependent Variable

Explanatory Variables	Saudi Arabia (n = 70)	United Arab Emirates (UAE) (n = 35)	Arabian Gulf Region Excluding Saudi Arabia (n = 77)	Arabian Gulf Excluding Saudi Arabia and UAE (n = 42)
	model 15	model 16	model 17	model 18
constant	1.78 (4.31)***	0.90 (3.12)***	-0.20 (-0.48)	-0.90 (-157)
$mair_1$	0.0003 (2.70)***	0.0003 (1.88)*	0.0002 (2.00)**	0.0009 (1.99)**
log_size	-0.056 (-0.81)	0.39 (2.38)**	0.36 (3.59)***	0.34 (2.93)***
log_age	0.002 (0.05)	-0.05 (-0.47)	0.014 (0.17)	0.14 (1.27)
urp	-0.004 (-0.14)	0.05 (0.35)	0.04 (0.82)	0.034 (0.65)
retn	-0.011 (-2.72)***	0.01 (1.22)	0.002 (0.49)	0.003 (0.78)
sub	0.001 (0.86)	-0.0004 (-0.07)	0.001 (1.10)	0.011 (1.87)*
fin	-0.19 (-1.62)	0.24 (1.89)*	0.16 (1.20)	0.10 (0.63)
crisis	-0.01 (-0.16)	-0.10 (-0.36)	0.04 (0.31)	0.10 (0.63)
f value	3.59**	2.55**	2.53**	2.16**
adjusted r^2	0.05	0.05	0.16	0.21

*The dependent variable BETA is the measure of IPO stock's market risk factor calculated over the one year in aftermarket period. Among the explanatory variables, $MAIR_1$ is the market adjusted initial return on the 1st day of IPO listing, and $MAIR_{30}$ measures the market adjusted initial return on the 30th day after listing LOG_SIZE is the natural log of the offer size measured by the number of shares in offer times the offer price. LOG_AGE is the natural log of the age of the company measured by the time period between the day of incorporation and the day of exchange listing. URP is the measure of underwriters' reputation. RETN is that percent of ownership retained by the existing directors following share floatation. SUB is that issue subscription times relative to the offer size. FIN is a dichotomous variable where FIN = 1 if IPO is listed in Financial sector, else FIN = 0. CRISIS is a dichotomous variable where CRISIS =1 if IPO is listed during financial crisis period (after July 2007) else CRISIS = 0. Numbers in parentheses shows the t-values of respective co-efficient. Asterisks ***, **, and * denote the level of significance at respectively 1, 5, and 10 percents*

This is important because the existence of such relationship in the different parts of the world may imply that underpricing of public offers would additionally be needed to compensate for aftermarket risk, besides the ex-ante uncertainty risk Hence, we test the relationship between the IPO aftermarket risk and underpricing in the markets of Arabian Gulf countries (Saudi Arabia, United Arab Emirates, Kuwait, Qatar, Bahrain, and Oman) a different economic region outside the US, where personal incomes and capital gains are tax-free. While the direction of relationship (causality) between the aftermarket risk and underpricing is yet to be investigated, we assume that IPO value uncertainty may not be fully resolved until a normal market is established after the listing of stock, and the aftermarket beta may have a link to the IPO firm's business and financial risks that has been inherited from the period before listing.

The empirical examination based on 147 IPOs in the Arabian Gulf markets showed that the average underpricing is about 250 percent. The level of underpricing reported in this study is consistent with the previous finding based on 47 Arabian Gulf IPOs that is reported by Al-Hassan, *et al*. (2010). The regression results consistently depict that IPO aftermarket risk (stock's beta) maintains a positive relationship with the level of initial underpricing in all the markets of Arabian Gulf region. While the causality is yet to be known, based on the assumptions mentioned above, we first test the relationship between aftermarket risk and underpricing by using the stock's beta as the explanatory variables along with other ex-ante uncertainty proxy variables. The results showed that beta significantly affects the underpricing in all the Arabian Gulf markets. The relationship is re-examined by using stock beta as the dependant variable and underpricing as an explanatory variable, along with other ex-ante uncertainty proxies. These results displayed that underpricing can also affect the aftermarket risk. Hence, in both ways, a significant relationship between the aftermarket risk and underpricing is found to exist in the Arabian Gulf markets. However, the models that include beta as an explanatory variable with underpricing as the dependant variable have more explanatory power than that of the models which include beta as the dependant variable with underpricing as an explanatory variable.

Therefore, this study confirms that a positive relationship between the IPO aftermarket risk and underpricing also exists in the Arabian Gulf countries - an important economic region outside the US where personal incomes including capital gains are tax-free. Since it is found that IPO aftermarket stock beta significantly affects the level of underpricing, with higher explanatory power in the multivariate tests, it may be concluded that underwriters and/or issuers in the Arabian Gulf countries may need to forecast the expected aftermarket risk while determining the offer price. Thereby, the degree of initial underpricing may be a useful predictor of the level of aftermarket risk as it was suggested by an earlier study. Given the present results, the future studies will explore the sources of the positive relationship between the aftermarket risk and underpricing.

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