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THE RISK FACTORS ASSOCIATED WITH INVESTING IN AN EMERGING EQUITY MARKET DURING THE EU MEMBERSHIP PROCESS

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ABSTRACT

This paper identifies the risks associated with investing in the Turkish stock market. We find that Turkish firms are more volatile than firms in countries that have recently joined the EU (our control group) and that the excess volatility is significantly associated with higher financial and economic risks rather than fundamentals. Additionally, firms' fundamentals are as important as country risk factors in explaining stock risk premiums for the control group, while the combined effect of country risk scores has a greater impact on risk premiums than firms' fundamentals alone for Turkish firms. Also, while Turkish stocks are sensitive to all country risk factors — economic conditions, international openness, investment profile, conflicts, and social tensions — stocks of the control group are mostly affected by only two factors, namely social tensions and economic conditions. Finally, some risks have become less relevant as a result of the changes in legal, political, and economic policies that occurred from 1999 to 2004 (the candidacy period for EU membership). Overall, Turkey has been quite successful at pursuing reforms since it began its candidacy for the EU. It has liberalized its political system and relaxed restrictions on freedom. It has also reduced hyperinflation, strengthened its currency, lowered interest rates, and provided more stable growth in GDP. However, political stability and financial and economic development appear to be issues for Turkey in its quest to become an EU member.

JEL: F3; G1; N2

INTRODUCTION

Turkey formally applied to join the European Community (now, the European Union) on April 14, 1987. It was officially recognized as a candidate for membership on December 10, 1999. The hope of joining the EU has driven major reforms in Turkey, including economic liberalization, human rights protection, and greater civilian oversight of the military. In 2002, the EU outlined the political and economic conditions that Turkey would have to satisfy before formal accession talks could begin. The criteria required that Turkey have a functioning market economy and stable institutions that guarantee democracy, the rule of law, and human rights.

Since commencing its official candidacy for membership in the EU, Turkey has pursued reforms involving liberalizing the political system and relaxing restrictions on freedom and human rights. It has abolished the death penalty, adopted measures to promote independence of the judiciary, and reformed the prison system. In addition, Turkey has modified its penal legal environment. Military and police powers have been somewhat lessened and the administration of justice has been strengthened. Turkey has also started economic and financial reforms leading to reduced hyperinflation, a more fairly valued currency, lower interest rates, and a decreasing amount of past-due loans which used to account for more than 20 percent of all banking system credit. With $39.5 billion of assistance from the International Monetary Fund, it has shrunken the pension system, downsized the public sector, and reformed bankruptcy law. By mid-2004, inflation was reduced to 13 percent, its lowest level in almost 30 years, and Turkey's GDP growth for 2004 was around 5 percent.

However, the EU still perceives Turkey as too unstable politically, and too underdeveloped financially and economically to become a member. Turkish general opinion does not perceive positively the EU's hesitation to incorporate their country into the union. Ten countries (the Greek part of Cyprus, the Czech Republic, Estonia, Hungary, Latvia, Lithuania, Malta, Poland, Slovakia and Slovenia) started their candidacy for membership in the EU around the same time as Turkey; all of them joined the European Union in May of 2004. Turkish opinion widely believes that the underlying reason for rejection is associated with cultural differences rather than economic, political and financial weaknesses.

The purpose of this paper is to shed some light on the reasons behind EU's decision by identifying the risks associated with investing in the Turkish stock market. Indeed, capital markets are mirrors of the expected changes in the political, economic and financial landscape of a country[1] . If cultural differences are at the source of the discord between Turkey and the EU, then the excess volatility found in the Turkish capital market compared to the new EU members' capital markets during their common candidacy period should be driven by differences in firms' fundamentals rather than by relative political, financial and economic underdevelopment.

We first investigate the source of excess volatility found in a portfolio of 78 Turkish firms compared to a portfolio of 176 stocks traded in a control group comprised of the Czech Republic, Estonia, Hungary, Latvia, Lithuania, Poland, Slovakia and Slovenia (8 of the 10 new EU members) during their candidacy period. We notice that Turkish firms are indeed much more volatile and that the excess volatility can be significantly associated with higher financial and economic risks rather than fundamentals. We further examine the relationship between risks and stock returns using a multifactor extension of the CAPM and show that firms' fundamentals are as important as country risk factors in explaining stock risk premiums for our control group, while the combined effect of country risk scores has a greater impact on risk premiums than firms' fundamentals alone for Turkish firms. Also, while Turkish stocks are sensitive to all country risk factors — economic conditions, international openness, investment profile, conflicts, and social tensions — stocks of the control group are mostly affected by only two factors, namely social tensions and economic conditions.

Finally, some risks have become less relevant as a result of changes in legal, political, and economic policies that occurred from 1999 to 2004. We conclude that Turkey has been quite successful at pursuing reforms since commencing its candidacy for membership in the EU. It has liberalized its political system and relaxed restrictions on freedom. As well it has reduced hyperinflation, strengthened its currency, lowered interest rates, and provided more stable growth in GDP. However, political instability and financial and economic underdevelopment appear to be major issues for Turkey in its quest to become an EU member. The remainder of the paper is organized as follows. Section 2 briefly discusses the relevant literature. Data selection, research methodology, and empirical models are described in Section 3. Section 4 provides analysis and interpretations of the empirical findings and Section 5 concludes the paper.

LITERATURE REVIEW

Finance theory suggests that pricing of assets always starts by evaluating the risks involved with investing in them. When it comes to stocks traded in emerging markets, finance literature suggests that risks are both fundamentals-related and country-specific. For instance, Erb, Harvey and Viskanta (1995) show how a country risk rating model explains the return generating process in world markets. The authors use composite risks such as political, economic and financial risk ratings and country credit ratings from the International Country Risk Guide (ICRG) and explore how these are correlated with wealth. They also observe that a lower rating (higher risk) is associated with higher expected returns. In a related article, Erb, Harvey and Viskanta (1996b) investigate how ICRG composite risk scores (political, financial and economic risk) explain the cross-sections of expected returns on IFC country indexes. They find that economic and financial risks convey the most information about expected returns in developed markets, while political risk has some marginal explanatory power in emerging equity markets. They also investigate the relationship between the world beta, the index volatility, one fundamental attribute at the country level (index aggregate book-to-price value) and composite risk scores. Their findings suggest that composite risk scores are highly correlated with country fundamentals. Similar conclusions have been reached by other authors. For instance, Oijen and Perotti (2001) indicate that changes in political risk are a priced factor and tend to have a strong effect on local stock market development and excess returns in emerging economies. La Porta, Lopez-de-Silanes, Shleifer, and Vishny (1997) find that countries with lower quality of legal rules and law enforcement have smaller and narrower capital markets. Demirgüç-Kunt and Maksimovic (1998) show that firms traded in countries with high ratings for the effectiveness of their legal systems are able to grow faster by relying more on external capital.

At the firm level, empirical research has shown that some fundamental firm-specific factors (such as size or book value to market value of equity) are more suited to describe the cross-sections of stock returns. Many studies have shown that high beta, small, value and high momentum firms have higher cross-sectional risk premiums in developed markets (Chan, Hamao, and Lakonishok, 1991; Aggarwal, Hiraki, and Rao, 1992; Fama and French, 1992 and 1996). As for the risks explaining the return-generating processes of stocks traded in emerging capital markets, findings are dichotomous. While Fama and French (1998), Patel (1998) and Rouwenhorst (1999) argue that risk premiums in emerging markets exhibit the same characteristics as in developed markets (i.e., significant momentum, small stocks outperform large stocks and value stocks outperform growth stocks) others report mixed results for the relationship between fundamental attributes and returns in emerging markets. These include Claessens, Dasgupta, and Glen (1995, 1998), Lyn and Zychowicz (2004), Ramcharran (2004) and Girard and Omran (2006). In some instances the above authors find positive relationships between size and returns as well as a positive relationship between price-to-book value and returns, contrary to the conventional belief that small and value firms are riskier. Several arguments may account for these findings. Daniel and Titman (1997) propose that firms' characteristics explain the return premium — i .e., a value premium will exist in emerging markets if value stocks are less liquid than growth stocks. Harvey and Roper (1999) argue that market growth has led to the mobilization of new capital and an increase in the number of firms rather than an increase in value. Furthermore, due to either the restrictions on debt financing or the immature debt markets, small firms have a capital structure comprised principally of equity, while larger firms with their international exposure can more easily access leverage. For instance, Bolbol and Omran (2005) and Girard and Omran (2006) indicate that only large firms have higher leverage ratios in Arab markets. Claessens, Dasgupta, and Glen (1998) also suggest that market microstructure causes these substantial differences and that regulatory and tax regimes force investors to behave differently in nascent markets. The authors also hypothesize that the positive relationships between returns and size and market-to-book value can be attributed to the segmentation of financial markets.

In a recent article, Girard and Omran (2006) investigate how firms' fundamentals and country risk ratings provide an explanation for the return-generating process of individual stocks traded in emerging markets. Their study shows that a constant beta is not a good proxy for risk in thinly traded emerging markets, and firms' fundamentals and country risk rating factors are important in explaining the cross-sections of stock returns. Furthermore, they suggest that a pricing model including firms' fundamentals and country risk rating factors has significantly better explanatory power than either a CAPM, or a model which only includes a firms' fundamentals, or a model based only on country composite risk ratings. The authors conclude that financial transparency and political instability are still powerful obstacles to investing in emerging markets.

DATA AND METHODOLOGY

The sample consists of firms traded at the Istanbul Stock Exchange during the period September 1988 to June 2004. Monthly indexes, stock prices and firms' fundamentals are obtained from the S&P/IFC Emerging Markets Data Base[2] . The country risk ratings are obtained from the ICRG[3] managed by the Political Risk Group. ICRG country risk scores are grouped into three categories, which consist of 12 political risk, 5 financial risk and 5 economic risk scores. ICRG scales rank risks from a high score, indicating a low risk, to a low score, indicating a high risk. We retrieve the Turkish IFCG market indexes[4], individual firms' monthly stock returns, market capitalization and price-to-book values. We choose monthly prices in US dollars to circumvent the problem of high inflation. All monthly indexes and stock returns are then deflated[5] using the US 90-day T-bill rate in the following formulae:

$$R_{i,t,deflated} = \frac{(r_{i,t} - r_{f,t})}{(1 + r_{f,t})} \tag{1a}$$

$$R_{m,t,deflated} = \frac{(r_{m,t} - r_{f,t})}{(1 + r_{f,t})} \tag{1b}$$

where $r_{i,t}$ and $r_{m,t}$ are the monthly stock and market returns[6] , while $r_{f,t}$ is the monthly US T-bill rate.

We compute local betas by regressing each stock dollar's returns on a country index to which the firm belongs as in Rouwenhorst (1999). This equally weighted country index is comprised of dollar-denominated stock returns averaged each month[7]. One lag of the equally weighted country index is included to allow for a delayed response due to non-synchronous trading. Betas are computed with a minimum of two years and a maximum of five years of historical monthly returns. Each stock return is matched by a monthly size (market capitalization in US dollars) and a price-to-book value (PB). The total number of Turkish firms included in the IFC is 91. However, out of the original sample, 13 stocks had less than 24 months of data; hence, we have to exclude those firms from the analysis. Our sample consists of 78 firms traded from September 1988 to June 2004.

We investigate the cross-sections of risk premiums of stocks traded in the Turkish capital market with k-risk factors comprised of three groups of firm risk components (beta, the logarithm of a firm's market capitalization, and the logarithm of price-to-book value) and 22 risk scores (12 ICRG political risk scores, 5 ICRG economic risk scores, and 5 ICRG financial risk scores). Our approach is similar to that of Girard and Omran (2006). We follow a principal component analysis methodology to reduce the factor loading, and identify the significance of each risk factor's effects on stock risk premiums. Finally, we test the information content of our multifactor expression as compared to a simpler nested model — a three-factor composite risk model. In order to avoid arbitrary weighting of risk scores by using a composite measure, we utilize a principal component analysis to select the main risk drivers within a risk category [8]. Our fundamentals and country risk factor model should have each asset return linearly related to *k* factors plus its own idiosyncratic disturbance as follows:

$$R_{i,deflated} = b_0 + b_1\beta_i + b_2 \log(PTBV) + b_3 \log(size) + \sum_{i=1}^{k} \lambda_i \tilde{Z}_i + \varepsilon_i \tag{2}$$

where $R_{i,deflated}$ is a vector of monthly deflated stock returns; b_1, b_2 and b_3 are the risk premiums associated with beta, the price-to-book ratio and the market capitalization of a stock; Z_i is a vector of common country risk score factors determined using a principal component analysis; and λ_i is a vector of risk premiums associated with the country risk score factors. Finally, we compare equation 2 to a model proposed in the literature at an aggregate level and on a country basis. As in Erb, Harvey and Vistanka

(1996a), a country risk composite model relates return to political risk (PR), economic risk (ER), and financial risk (FR), i.e.,

$$R_{i,deflated} = \lambda_0 + \lambda_1 Ln(PR) + \lambda_2 Ln(ER) + \lambda_3 Ln(FR) + \varepsilon_i \qquad (3)$$

We use three tests[9] to compare the explanatory power of equation 3 to that of equation 2. These include the Davidson and MacKinnon test (1981), the posterior odds ratio and a partitioned residual analysis.

EMPIRICAL RESULTS

We first provide information about the monthly IFCG index's return, the monthly standard deviation of the index return, the market capitalization, the price-to-book ratio, four composite macro risk ratings, and twenty-two individual macro risk ratings[10] (see Table 1). Metrics are reported over three periods: the overall sample period (1988:09 to 2004:06), a period prior to candidacy for EU membership (1988:09 to 1999:11), and the current period of candidacy for EU membership (1999:12 to 2004:06). The average monthly index return is 0.65% during the overall sample period. It decreased from 0.88% (1988-1999) to 0.27% (1999-2004). The average monthly standard deviation is 16.33% for the overall period; it increased from 15.91% for the first period to 17.02% for the second period. Hence, Turkish stocks appear to be providing less and have become riskier. The average market capitalization of the index is $40 billion and it has increased from $26 billion to $65 billion. The number of firms traded at the ISE increased from 164 in the pre-candidacy period to 290 during the candidacy period. Turkish stocks are traded at 4.66 times their book values during the overall period.

This figure decreased from 5.16 during the pre-candidacy period to 3.82 during the candidacy period. This is an indication of how Turkish stocks have become riskier and value-oriented. While composite risk ratings (50 percent weighted in political risk rating) and political risk ratings are higher during the candidacy period compared to the pre-candidacy period, financial and economic risk ratings are lower during the latter period. This indicates that Turkey has improved its political landscape but has failed to do the same at a financial and economic level. More specifically, issues related to government stability, investment profile, trade deficit, inflation and stability of GDP growth have dramatically improved from the pre-candidacy to the candidacy period. However, risks associated with socioeconomic conditions, corruption, democratic accountability, ethnic tensions and debt servicing have slightly increased from one period to the other.

The descriptive statistics of the 78 firms used in our analysis are presented in Table 2. The final sample consists of 7,806 monthly observations for firms traded from 1988:09 to 2004:06. The median monthly return is -0.04 percent and the median monthly standard deviation is 0.29 percent. Firms have a median market capitalization of $190.031 million and a median price-to-book ratio of 3.494. By EU standards, the firms traded on the Turkish capital market are relatively small and their returns are extremely volatile

We start our analysis by comparing Turkey's risk ratings during the candidacy period with those of two groups of countries. Group 1 is comprised of the 15 EU member states as of April 2004 (Austria, Belgium, Denmark, Finland, France, Germany, Greece, Ireland, Italy, Luxembourg, the Netherlands, Portugal, Spain, Sweden, and the United Kingdom). Group 2 is comprised of the ten countries that joined the EU in May of 2004 (the Greek part of Cyprus, the Czech Republic, Estonia, Hungary, Latvia, Lithuania, Malta, Poland, Slovakia and Slovenia). Composite, political, financial and economic risk ratings are much lower in Turkey as compared to groups 1 and 2. For example, Turkey has the lowest ratings in 9 out of 12 political risks, 2 out of 5 economic risks, and 2 out of 5 financial risks. Out of the remaining 9 risk ratings, 7 are in the lower range of group 2 and the remaining 2 are close to medians of groups 1 and 2. Thus, during the candidacy period for joining the EU, Turkey shows greater political, financial and economic risks as compared to the norm of the 15 EU members. Furthermore it has lower ratings compared to the group of 10 countries that joined the EU in May of 2004 (Table 3).

Table 1: The Turkish Capital Market: Risk and Return from January 1988 to June 2004

	Overall Period	**Pre-candidature[a] Period**	**Candidature[a] Period:**
Median # of companies traded	229	164	290
Median # of companies included in IFCG	44	38	53
Average IFCG Turkey Return	0.65%	0.88%	0.27%
IFCG Turkey Standard Deviation	16.33%	15.91%	17.02%
IFCG Turkey Market Value	40,374.69	25,611.10	64,980.67
IFCG Turkey Price-to-Book Ratio	4.66	5.16	3.82
Composite Risk Rating	56.66	56.31	57.25
Political Risk Rating	56.29	54.88	58.63
Financial Risk Rating	29.57	29.71	29.34
Economic Risk Rating	27.35	27.88	26.45
Political Risk Ratings Variables			
Government Stability	7.33	6.45	8.80
Socioeconomic Conditions	3.96	4.04	3.83
Investment Profile	6.42	5.42	8.09
Internal Conflict	7.19	6.74	7.95
External Conflict	9.24	9.30	9.14
Corruption	2.65	2.82	2.38
Military in Politics	3.23	3.58	2.64
Law & Order	3.74	3.59	3.99
Religious Tensions	3.71	3.40	4.22
Ethnic Tensions	2.26	2.42	2.00
Democratic Accountability	4.22	4.59	3.60
Bureaucracy Quality	2.35	2.56	2.00
Financial Risk Rating Variables			
Budget Balance	3.58	4.06	2.79
Current Account as % of GDP	8.40	6.75	11.14
Current Account as % of XGS	11.73	11.65	11.86
Debt Service	6.38	6.62	5.99
Exchange Rate Stability	5.31	5.31	5.32
Economic Risk Rating Variables			
Foreign Debt	4.91	4.78	5.12
GDP Growth	5.84	4.84	7.50
GDP per Head of Population	1.89	2.01	1.68
Inflation	2.57	2.11	3.34
International Liquidity	0.60	0.32	1.06

Composite Risk Rating: ½ of the sum of political, financial and economic risk ratings. Politic Risk Rating is the sum of the following risk ratings: Government Stability: risk associated with a government's ability to carry out its declared program(s), and its ability to stay in office. Socioeconomic Conditions: risk associated with the general public satisfaction with the government's economic policies. Investment Profile: risk associated with expropriation, taxation, repatriation of capital, and labor costs. Internal Conflict: risk associated with political violence and its impact on governance. External Conflict: risk to both the incumbent government and inward investment. Corruption risk: risk associated with corruption within the political system. Military in Politics: risk associated with military involvement in politics. Religious Tensions: risk associated with the domination of a single religious group or the suppression of religious freedom. Law and Order: risk associated with the weakness and partiality of a legal system, and the lack of observance of the law. Ethnic Tensions: risk associated with tensions within a country attributable to racial, nationality, or language divisions. Democratic Accountability: risk associated with a government that is not responsive to its people. Financial Risk Rating is the sum of the following risk ratings: Foreign Debt as a % of GDP: risk associated with gross foreign debt in a given year, converted into US dollars. Foreign Debt Service as a % of Exports of Goods and Services: risk associated with foreign debt service per year, in $US. Current Account as a % of Exports of Goods and Services: risk associated with the annual current account deficit, in $US. Net International Liquidity as Months of Import: risk associated with the total estimated official reserves for a given year, in $US. Exchange Rate Stability: risk associated with the appreciation/ depreciation of a currency against the $US (against the DM for the US). Economic Risk Rating is the sum of the following risk ratings: GDP Per Head: Risk associated with a low GDP per head for a given year, converted into $US. Real GDP Growth: risk associated with a % increase or decrease in the estimated GDP, at constant 1990 prices. Annual Inflation Rate: Risk associated with annual inflation rate. Budget Balance as a % of GDP: Risk associated with a government budget deficit for a given year in the national currency. Current Account as a % of GDP: risk associated with the current account balance deficit for a given year, converted into $US.[a:] Turkey was officially recognized as a candidate for membership on December 10, 1999

Table 2: Companies Traded on the Turkish Capital Market: Risk and Return September 1988 -June 2004

Name	Start	End	Obs.	Return	Stdev	Beta	MV	PB
Adana Cimento-A	199302	200210	117	-0.048	0.278	0.963	62.511	10.509
Ak Enerji	200112	200406	31	-0.059	0.314	1.061	295.377	1.469
Akal Tekstil	199402	199910	69	-0.069	0.280	1.162	55.177	2.553
Akbank	199101	200406	162	-0.044	0.258	0.966	2,596.994	3.345
Akcansa	199611	200406	92	-0.039	0.273	0.758	411.752	3.657
Akcimento	198809	199609	97	-0.051	0.377	0.920	100.121	3.598
Aksa	199302	200406	77	-0.070	0.341	0.219	392.508	4.346
Aksigorta	199612	200406	91	-0.026	0.274	1.137	337.093	4.014
Aktas Electrik	199612	200009	46	-0.032	0.274	0.650	362.524	17.511
Alarko Gayrimenkul	199712	200210	59	-0.022	0.213	0.855	41.013	0.911
Alarko Holding	200012	200406	43	-0.011	0.219	1.014	291.905	5.048
Alcatel-Bearer	199602	200310	93	-0.042	0.345	1.196	154.629	8.323
Anadolu Efes Biracilik	199402	200406	125	-0.037	0.279	0.635	657.434	11.542
Anadolu Isuzu	199712	200310	71	-0.033	0.309	1.163	115.332	4.144
Arcelik	198902	200406	185	-0.032	0.275	1.058	897.736	5.172
Aselsan-Bearer	199602	200406	101	-0.027	0.265	1.012	174.939	3.280
Aygaz	199302	200406	137	-0.043	0.255	0.831	402.865	5.275
Bagfas	198809	200210	170	-0.026	0.310	0.979	78.707	2.896
Bosch Fren-Bearer	199602	200210	81	-0.005	0.192	0.596	33.149	4.614
Brisa	198809	200406	190	-0.025	0.317	0.970	243.432	2.998
Celebi Hava Servisi	199712	200406	79	-0.055	0.357	0.781	70.045	6.528
Celik Halat	198809	200210	170	-0.045	0.286	1.012	26.000	2.867
Cimentas	199402	199710	45	-0.057	0.270	0.474	127.802	4.407
Cimsa	198809	200406	190	-0.044	0.332	0.810	157.980	3.407
Cukurova Elektrik	198809	200009	145	-0.013	0.251	0.726	469.724	10.261
Dardanel-Bearer	199602	200110	69	-0.070	0.298	0.990	30.074	1.389
Dogan Holding	199402	200406	125	-0.060	0.369	1.326	517.939	3.561
Dogan Yayin Holding	200012	200406	43	-0.019	0.331	1.442	494.086	2.748
Eczacibasi Ilac	199101	200406	162	-0.053	0.329	1.103	189.974	2.974
Eczacibasi Yatirim	199302	200210	117	-0.048	0.359	1.115	57.452	2.898
Ege Bira	199302	199710	57	-0.065	0.319	0.460	562.801	13.248
Erdemir	198809	200406	190	-0.036	0.339	0.767	795.729	1.928
Goodyear	198809	200110	158	-0.051	0.314	0.988	138.496	5.526
Guney Bira	199402	199910	69	-0.052	0.290	0.574	84.265	6.309
Hurriyet Gazette	200012	200406	43	-0.023	0.311	1.288	424.014	2.818
Ihlas Holding	199612	200109	58	-0.079	0.301	0.890	270.407	3.555
Is Gayrimenkul	200012	200406	43	-0.024	0.205	0.928	170.366	0.752
Izmir Demir Celik-Bearer	199502	200210	93	-0.046	0.293	1.064	83.057	0.867
Kartonsan	198809	200110	158	-0.020	0.228	0.628	100.338	2.444
Kepez Elektrik-Bearer	199602	200009	56	0.006	0.224	0.947	157.163	13.757
Koc Holding	198902	200406	185	-0.032	0.303	1.151	2,055.346	10.527
Koc Yatirim	199302	199708	55	-0.034	0.257	0.773	180.417	9.332
Kordsa	198809	200406	190	-0.040	0.273	0.861	163.272	3.240
Koruma Tarim	198809	199312	64	-0.043	0.287	0.801	13.901	1.984
Mardin Cimento	199302	200210	117	-0.054	0.275	0.858	50.223	3.359
Medya Holding	199612	200103	52	-0.011	0.289	0.779	175.669	3.763
Mensucat Santral	198902	199312	59	-0.070	0.296	0.811	90.697	1.786
Migros	199402	200406	125	-0.063	0.324	0.708	725.510	15.146
Net Holding	199612	200210	71	-0.066	0.327	1.242	67.124	1.158
Net Turizm	199712	200210	59	-0.097	0.360	1.184	39.944	1.612
Netas	199402	199710	45	-0.040	0.220	0.853	357.578	5.706
Netas Telekom	200012	200406	43	-0.037	0.235	1.143	190.088	3.236
Otosan	198902	200406	185	-0.029	0.375	1.061	618.540	7.194
Petkim	199101	200406	162	-0.032	0.402	0.998	1267.614	2.271
Petrol Ofisi	199302	200211	118	-0.039	0.349	0.939	1,435.994	11.018
Rabak	198809	199312	64	-0.037	0.317	1.075	17.721	1.581

Name	Start	End	Obs.	Return	Stdev	Beta	MV	PB
Raks Elektronik-Bearer	199502	199910	57	-0.078	0.411	0.246	65.671	1.821
Sabanci Holding	199712	200406	79	-0.037	0.285	1.200	3,264.424	4.578
Sarkuysan	198809	200310	182	-0.033	0.293	1.007	75.665	3.434
Sasa Dupont	199712	200406	79	-0.050	0.319	1.019	221.024	2.119
Sifas-Bearer	199602	199910	45	-0.082	0.259	0.640	20.828	1.443
T. Demir Dokum	198809	199710	110	-0.030	0.269	1.000	141.138	5.017
T. Garanti Bankasi	199101	200406	162	-0.037	0.303	1.187	1287.168	2.667
T. Is.bank (C)-Bearer	199602	200406	101	-0.032	0.311	1.068	5,020.097	4.838
T. Sise Cam	199101	200406	162	-0.050	0.310	1.216	434.984	3.917
T.Is Bank	199101	199312	36	-0.038	0.328	1.315	559.306	1.545
Tansas	200112	200406	31	-0.026	0.199	1.037	228.763	0.926
Tat Konserve-Bearer	199502	200210	93	-0.060	0.282	0.861	127.668	4.890
THY	199402	200406	125	-0.038	0.298	1.146	1,689.879	9.946
Tire Kutsan-Bearer	199602	200210	81	-0.052	0.261	0.611	42.821	2.359
Tofas Oto Fab	199302	200406	137	-0.051	0.304	0.937	822.434	5.253
Tofas Oto Tic	199302	200105	100	-0.067	0.351	1.121	100.658	9.554
Trakya Cam	200112	200406	31	-0.019	0.175	0.667	366.524	2.155
Tupras	199302	200406	137	-0.029	0.333	1.082	2,218.224	13.460
Turkcell	200008	200406	47	-0.036	0.280	1.124	4310.721	12.524
Ucak Servisi-Bearer	199502	200310	105	-0.060	0.534	1.129	86.025	7.378
Vestel	200012	200406	43	-0.088	0.591	0.761	420.424	1.775
YKB	199101	200406	162	-0.042	0.304	1.292	1453.307	2.060
Median				-0.040	0.298	0.984	190.031	3.494

Table 3: Comparison of Risk Ratings Between EC Member States Turkey (Dec 1999[a] - Jun 2004)

	Group 1[b]			Group 2[c]			Turkey
Risk Ratings	**Max.**	**Med.**	**Min.**	**Max.**	**Med.**	**Min.**	**Med.**
Composite Risk Rating	90.4	83.4	75.3	79.9	75.8	74.0	57.2
Political Risk Rating	93.5	88.0	77.9	86.5	77.7	71.4	58.6
Financial Risk Rating	42.3	41.8	34.1	42.8	39.0	36.1	29.3
Economic Risk Rating	45.4	38.9	36.6	39.6	38.6	33.1	26.5
Government Stability	10.6	9.4	8.6	10.4	9.0	7.3	8.8
Socioeconomic Conditions	10.8	9.1	6.9	9.7	6.2	5.1	3.8
Investment Profile	11.6	10.8	10.3	10.8	10.2	9.2	8.1
Internal Conflict	12.0	11.3	8.4	11.4	11.0	8.9	8.0
External Conflict	12.0	11.1	8.8	11.6	10.9	8.3	9.1
Corruption	6.0	4.3	2.7	4.0	3.4	2.5	2.4
Military in Politics	6.0	5.0	4.7	6.0	5.5	4.5	2.6
Law & Order	6.0	5.0	3.0	5.3	4.6	4.0	3.9
Religious Tensions	6.0	5.0	4.6	6.0	5.0	4.1	4.2
Ethnic Tensions	6.0	4.0	2.8	6.0	3.6	2.2	2.0
Democratic Accountability	6.0	5.9	4.1	6.0	5.6	5.0	3.6
Bureaucracy Quality	4.0	3.5	2.7	4.0	3.0	2.3	2.0
Budget Balance	9.2	7.6	6.7	7.8	6.0	4.6	2.8
Current Account as % of GDP	13.9	12.0	9.4	12.3	10.0	9.2	11.1
Current Account as % of XGS	13.1	12.4	10.1	12.1	11.2	10.5	11.9
Debt Service	10.0	8.0	5.0	9.3	8.9	6.0	6.0
Exchange Rate Stability	9.7	9.2	9.0	9.9	9.2	8.7	5.3
Foreign Debt	10.0	8.5	4.6	9.8	7.0	3.5	5.1
GDP Growth	9.3	7.3	6.8	9.2	8.6	7.5	7.5
GDP Per Head of Population	5.0	4.4	3.8	3.8	2.5	1.9	1.7
Inflation	10.0	9.6	8.8	9.7	8.9	7.4	3.3
International Liquidity	3.3	1.5	0.3	3.5	2.1	1.5	1.1

[a] *Turkey was officially recognized as a candidate for membership on December 10, 1999 .*[b] *Group 1 is composed of the 15 EC member states as of April 2004: Austria, Belgium, Denmark, Finland, France, Germany, Greece, Ireland, Italy, Luxembourg, Netherlands, Portugal, Spain, Sweden, and United Kingdom.* [c] *Group 2 is composed of the ten countries that joined the EU in May of 2004: Cyprus (Greek part), the Czech Republic, Estonia, Hungary, Latvia, Lithuania, Malta, Poland, Slovakia and Slovenia.*

We compare 78 Turkish median monthly stock returns, median monthly fundamentals (price-to-book value, beta and size), and monthly standardized composite risk ratings (economic, financial and political risk ratings) with those of a portfolio made of 176 stocks traded in the countries making group 2 -- 28 Czech, 11 Estonian, 18 Hungarian, 17 Latvian, 26 Lithuanian, 43 Polish, 16 Slovakian, and 17 Slovenian stocks--during the EU candidacy period (1999-2004)[11]. To investigate whether the difference in fundamentals and composite ratings between Turkey and a portfolio containing 176 stocks traded in the 8 other markets explains the excess volatility observed in the Turkish market as compared to these other markets, we estimate the following equation:

$$\begin{aligned}(R_{Turkey,deflated} - R_{Group\,2,deflated})^2 = {} & a_0 \\ & + a_1(\beta_{Turkey} - \beta_{Group\,2}) \\ & + a_2[\log(PTBV_{Turkey}) - \log(PTBV_{Group\,2})] \\ & + a_3[\log(size_{Turkey}) - \log(size_{Group\,2})] \\ & + a_4(SEconR_{Turkey} - SEconR_{Group\,2}) \\ & + a_5(SFinR_{Turkey} - SFinR_{Group\,2}) \\ & + a_6(SPolR_{Turkey} - SPolR_{Group\,2}) + \varepsilon_i\end{aligned} \tag{4}$$

where $R_{Turkey,deflated}$ and $R_{Group2,deflated}$ are each month's median deflated stock returns (69 stocks for Turkey and 176 stocks for group 2); β_{Turkey} and β_{Group2} are each month's median betas; $PTBV_{Turkey}$ and $PTBV_{Group\,2}$ are each month's median price-to-book value; $\mathrm{size}_{\mathrm{Turkey}}$ and $\mathrm{size}_{\mathrm{Group2}}$ are each month's median market capitalization; $SEconR_{Turkey}$ and $SEconR_{Group2}$ are each month's median standardized economic risk ratings; $\mathrm{SFinR}_{\mathrm{Turkey}}$ and $\mathrm{SFinR}_{\mathrm{Group2}}$ are each month's median standardized financial risk ratings; and $SPolR_{Turkey}$ $SPolR_{Group2}$ are each month's median standardized political risk ratings. Results are reported in Table 4 suggesting that differences in economic and financial risk ratings are significantly related to the excess volatility. Furthermore, differences in fundamentals do not explain excess volatility. Thus, excess volatility is mainly driven by differences in relative financial and economic differences between Turkey and these countries. These findings support the argument that the EU's rejection of Turkey into its membership stems from concerns over Turkey being financially and economically less developed, rather than cultural differences.

Next, we investigate the difference in sensitivity to fundamentals and composite ratings between Turkey and a portfolio containing 176 stocks traded in the 8 other markets during their common candidacy period for becoming members of the EU. We first conduct a principal component analysis to select the risk scores by factors[12]. Table 5 presents the results from the factor analysis. The first row shows the number of common factors found using a VARIMAX rotation. In the second row, the eigenvalues represent the proportion of total variance in all the variables accounted for by that factor. To decide the number of factors to retain, we use the Kaiser criterion which involves dropping the eigenvalues less than one — i.e., unless a factor extracts at least as much as the equivalent of one original variable, we drop it. In the third row (% of variance), these values are expressed as a percentage of the total. As we can see, factor 1 accounts for 18.14 percent of the variance, factor 2 comprises 16.84 percent, and so on. The fourth row (cumulated %) contains the cumulative variance extracted and shows that the six dominant factors whose eigenvalues are more than one, add up to 72.59 percent of the total variance. These factors can be considered as the 5 major country risk factors that characterize groups 1 and 2 and Turkey.

The following rows show the loading of each risk score variable within each factor. The construction of the factors is not straightforward as it depends on the particular combination of observed variables that correlate highly with each factor. In order to minimize the subjective nature of the principal component analysis, we only consider individual risk score loadings with "good" correlation. Comrey and Lee (1992) define a "good" condition for a loading greater than 0.5 (or smaller than -0.5) — i.e., 50 percent overlapping variance. Each factor's composite score is determined by taking into account the risk scores that load highly in it. Accordingly, following Seiler (2004), each factor's score is computed using a summated scale methodology where selected loadings within each factor are added to determine a factor score. Since risk scores are not on a standardized scale, we have to ensure that each risk score selected for the composition of a risk factor is standardized so that equal importance is given to all risk scores in the summation process. Finally, the factor is computed using the logarithm of the sum. Our factors form coherent groups of selected associated variables. Also, each factor has positive loadings and follows the ICRG scale — i.e., a high value indicates a low risk and a low value indicates a high risk. Each of the six constructs is briefly reviewed below.

Table 4: Relationship between Excess Volatility and Fundamental and Country Risk Differentials December 1999 - June 2004

	Coefficient	Std. Error	t-stat	Prob.
a(0)	0.02818	0.097434	0.29	0.7737
a(1)	-0.09001	0.000196	-0.48	0.6327
a(2)	-0.92828	0.876114	-1.06	0.2947
a(3)	0.02150	0.013528	1.59	0.1185
a(4)	-0.04038	0.019841	-2.04**	0.0474
a(5)	-0.09044	0.021688	-4.17***	0.0001
a(6)	-0.06276	0.039260	-1.60	0.1165
Observations	55 months			
Adjusted R^2	0.33			
Durbin-Watson Stat.	1.99			

*where $R_{Turkey,deflated}$ and $R_{Group2,deflated}$ are each month's median deflated stock returns (64 stocks for Turkey and 159 stocks for group 2); β_{Turkey} and β_{Group2} are each month's median betas; $PTBV_{Turkey}$ and $PTBV_{Group\ 2}$ are each month's median price-to-book value; $size_{Turkey}$ and $size_{Group2}$ are each month's median market capitalization; $SEconR_{Turkey}$ and $SEconR_{Group2}$ are each month's median standardized economic risk ratings; $SFinR_{Turkey}$ and $SFinR_{Group2}$ are each month's median standardized financial risk ratings; and $SPolR_{Turkey}$ $SPolR_{Group2}$ are each month's median standardized political risk ratings. To correct for the presence of autocorrelation and heteroskedasticity, standard errors and t-statistics are calculated using the Newey-West heteroskedasticity and autocorrelation consistent (HAC) covariance matrix. Variance inflation factors are less than 2 suggesting the absence of multicolinearity. ***, **, and * indicate significance at the 1, 5 and 10 percent level.*

$$\begin{aligned}(R_{Turkey,deflated} - R_{Group\,2,deflated})^2 = {} & a_0 \\ & + a_1(\beta_{Turkey} - \beta_{Group\,2}) \\ & + a_2[\log(PTBV_{Turkey}) - \log(PTBV_{Group\,2})] \\ & + a_3[\log(size_{Turkey}) - \log(size_{Group\,2})] \\ & + a_4(SEconR_{Turkey} - SEconR_{Group\,2}) \\ & + a_5(SFinR_{Turkey} - SFinR_{Group\,2}) \\ & + a_6(SPolR_{Turkey} - SPolR_{Group\,2}) + \varepsilon_i\end{aligned}$$

The first factor's contributing variables are corruption, democratic accountability and religious and ethnic tensions. We name the first factor *social tension rating* because the contributing variables emphasize issues associated with corruption, democratic accountability, and religious and ethnic tensions. This factor accounts for 18.14 percent of the variance. The second factor consists of five significant variables: current accounts as a percentage of exchange of goods and services, current accounts as a percentage of GDP, GDP per inhabitant, GDP growth, and socio-economic conditions. Ultimately, this risk factor is named *economic condition rating*. This factor accounts for 16.8 percent of the variance. We name the third factor *conflict rating* because of its loading with internal and external conflicts. This factor accounts for 14.68 percent of the variance. The fourth factor relates to exchange rate stability and inflation and is summed up as an *investment profile rating.* It accounts for 12.59 percent of the variance. The fifth factor is an *international openness rating* as it addresses the stability of the current regime and the longevity of the laws passed or initiated as well as international liquidity and foreign investments. This factor accounts for 10.35 percent of the variance.

Table 5: Factor Analysis on Europe's 22 Country Risk Ratings (1999:12 to 2004:06) — Group1, Group 2, and Turkey

Name	Factor 1 Social Tensions	Factor 2 Economic Conditions	Factor 3 Conflicts	Factor 4 Investment Profile	Factor 5 International Openness
Corruption	0.757	0.381	0.038	-0.087	0.053
Democratic Accountability	0.701	-0.147	0.144	0.352	-0.158
Religious Tensions	0.668	0.190	0.058	0.200	0.053
Ethnic Tensions	0.653	0.039	0.258	0.143	-0.145
Risk Points for Budget Balance	0.495	0.095	0.202	0.292	0.443
Bureaucracy Quality	0.467	0.441	0.238	0.231	0.148
Risk Points for Current Account as % of XGS	0.118	0.850	0.15	-0.081	0.125
Risk Points for Current Account as % of GDP	0.184	0.765	0.136	-0.074	0.206
Risk Points for GDP per Head of Population	0.398	0.676	0.06	0.344	0.236
Risk Points for GDP Growth	0.168	0.667	0.134	-0.194	0.224
Socioeconomic Conditions	0.448	0.661	0.207	0.427	0.249
Law & Order	0.407	0.636	0.442	-0.007	0.287
Internal Conflict	0.265	0.096	0.795	0.100	0.038
Risk Points for Debt Service	-0.171	0.001	0.710	0.220	0.143
External Conflict	0.279	0.179	0.675	0.025	-0.167
Military in Politics	0.428	0.134	0.660	0.498	0.015
Risk Points for Exchange Rate Stability	0.112	-0.171	0.263	0.706	0.107
Investment Profile	0.133	0.311	-0.016	0.670	-0.230
Risk Points for Inflation	0.397	0.102	0.229	0.655	0.269
Risk Points for Foreign Debt	-0.107	0.123	-0.004	0.306	0.813
Risk Points for International Liquidity	0.077	-0.42	-0.143	0.213	0.798
Government Stability	0.461	-0.002	-0.177	-0.219	0.622
Eigenvalue	3.99	3.704	2.569	2.55	2.056
% of Variance	18.138	16.838	14.677	12.593	10.346
Cumulative	18.138	34.976	49.652	62.245	72.591

This table provides the statistics related to factor analysis. The top portion of the table provides the details with respect to eigenvalues, and the proportion of the total variation that these factors explain. "Cumulative" is the cumulative variation explained by these factors. The bottom portion of the table provides information about the rotated factor pattern. We use a VARIMAX rotation which minimizes the number of factors on which a single variable has a high loading. Kaiser's Measure of Sampling Adequacy is 0.8545 and the Barlett test of sphericity is significant at the 1 percent level. Highlighted figures indicate the factors selected (the cut-off is 0.5).

The next step is to identify which combination of these factors and firms' fundamentals can explain monthly stock risk premiums in Turkey and in a portfolio of stocks traded in 8 countries (group 2) which were recently accepted into the EU. We investigate equation 2 and its nested model, equation 3, to identify the significant factors that explain Turkish and group 2 stock returns from December 1999 to June 2004. The regression findings are shown in Table 6. We report all coefficients, standard errors, and standardized coefficients. At the bottom of the table, the sum of the absolute value of the standardized coefficients is reported; the significance of the sum is determined by a Wald test.

independent variable is not shared by other independent variables) indicating that our regressions are not likely affected by multi-collinearity. Equation 2 always provides a better fit than equation 3 — i.e., equation 2 has R^2 ranging from 0.11 to 0.18 and equation 3 has R^2 ranging from 9% to 12%. In addition, all residual tests unambiguously demonstrate that equation 2 is the best model. For instance, alphas in Davidson and MacKinnon (DM) tests are close to unity (equation 2 is 100 percent more effective than equation 3) and are significant at the 1% level. Posterior odd ratios are highly in favor of equation 2. The residual tests indicate that, while equation 2's factors provide significant additional information to supplement that found in equation 3, the reverse is not true. As Girard and Omran (2006) point out, political, economic and financial risk ratings assume a fixed weighting scheme among their respective constituents and there is no obvious empirical rationale for this. Furthermore, several risk scores constituting each composite rating are negatively correlated; indicating that, at the composite level, the effect of some risk scores will offset other risk scores.

Table 6: Determinant of Returns in Turkey and Group 2

	(Equation 3) $R_{i,deflated} = b_0 + b_1\beta_i + b_2 \ln(PB) + b_3 \ln(size) + \lambda_1 Ln(PR) + \lambda_2 Ln(ER) + \lambda_3 Ln(FR) + \varepsilon_i$		(Equation 2) $R_{i,deflated} = b_0 + b_1\beta_i + b_2 \ln(PB) + b_3 \ln(size) + \lambda_1 factor_1 + \lambda_2 factor_2 + \lambda_3 factor_3 + \lambda_4 factor_4 + \lambda_5 factor_5 + \varepsilon_i$	
	Group 2	***Turkey***	***Group 2***	***Turkey***
(Constant)	0.511	0.920***	1.040**	3.921***
Std. Error	0.430	0.328	0.500	0.607
Beta	0.006*	-0.002	0.018***	-0.011
Std. Error	0.004	0.013	0.006	0.016
SCoef	0.024	-0.002	0.046	-0.014
Ln(PB)	0.009***	0.012*	0.008***	0.014*
Std. Error	0.002	0.007	0.003	0.007
SCoef	0.074	0.036	0.068	0.044
Ln(size)	0.003**	0.011***	0.003*	0.013***
Std. Error	0.001	0.004	0.001	0.004
SCoef	0.050	0.055	0.047	0.053
Ln(economic risk)	0.013	-0.208***		
Std. Error	0.053	0.045		
SCoef	0.004	-0.096		
Ln(financial risk)	-0.107**	0.173***		
Std. Error	0.052	0.066		
SCoef	-0.029	0.055		
Ln(political risk)	-0.047	-0.250***		
Std. Error	0.066	0.076		
SCoef	-0.011	-0.060		
Factor1			-0.188***	0.078*
Std. Error			0.047	0.045
SCoef			-0.073	0.045
Factor2			-0.185***	-0.803***
Std. Error			0.066	0.130
SCoef			-0.065	-0.194
Factor3			0.063	-0.315**
Std. Error			0.073	0.128
SCoef			0.017	-0.073
Factor4			-0.010	0.197***
Std. Error			0.030	0.043
SCoef			-0.006	0.157
Factor5			-0.009	-0.331***
Std. Error			0.031	0.060
SCoef			-0.006	-0.166
R^2	0.088	0.118	0.111	0.180
N	4240	2808	4240	2808
# of stocks	176	78	176	78
F	6.099***	6.927***	7.854***	10.438***
Fundamentals premiums	0.148***	0.093***	0.161***	0.111***
Country premiums	0.044***	0.211***	0.167***	0.635***
Residuals Test 1	0.090	0.047	0.941***	0.978***
Residuals Test 2	>10^23	>10^25	0	0
Residuals Test 3	F1***,F2***,F3***,	F1***,F2***,F3***,	None	None

To correct for the presence of autocorrelation and heteroskedasticity, standard errors and t-statistics are calculated using the Newey-West heteroskedasticity and autocorrelation consistent (HAC) covariance matrix. Variance inflation factors are less than 2, suggesting the absence of multicolinearity. Standardized coefficients (SCOEF) are the coefficients obtained after standardizing the variables and they indicate that an increase in 1 standard deviation on one of the factors affects "beta" standard difference in Ri, holding constant the other predictors in the model. The sum of the absolute value of the standardized coefficients is also reported; the significance of the sum is determined by a Wald test. RT1 (Residual test 1) is the Davidson and MacKinnon (1981) equation which estimates the proportion of information (α , the effectiveness of a model as compared to a competing model) unexplained by eq. 2 which is explained by eq. 3 —.e., the test consists in measuring α in $Ri = \alpha R_{eq2} + (1-\alpha) R_{eq3} + \varepsilon_i$. RT2 (Residual test 2) is the ratio for posterior odds (POR) of eq. 2 over eq. 3 represented as $POR = \left[ESS_{eq2} / ESS_{eq3}\right]^{N/2} N^{(K_{eq2} - K_{eq3})/2}$, where ESS is the error sum of squares, N the number of observations, and K is the dimension of respective models. RT3 (Residual test 3) is the partitioned residual test and it consists of running a regression between $\varepsilon_{eq3,i}$ and the factor loadings of eq. 2; we also run a regression of the residuals of eq. 2 ($\varepsilon_{eq2,i}$) with the factor loadings of eq. 3 to check for information missed by eq. 2. ***,** and * indicate significance at the 1, 5 and 10 percent level.

The variance inflation factors (for the sake of brevity, unreported) for each independent variable are extremely low for each period (less than 1.5, that is, more than 67 percent of the variance of each

The first interesting finding is that firms' fundamentals are as important as country risk factors in explaining stock risk premiums for the group 2 portfolio — e.g., a one-standard deviation shock on fundamentals leads to a 0.15 to 0.16 standard deviation shock on R_i, and a one-standard deviation shock on country risk ratings leads to a 0.04 to 0.17 standard deviation shock on R_i. However, the combined effect of country risk scores has a greater impact on risk premiums than firms' fundamentals alone in Turkey — e.g., a one-standard deviation shock on fundamentals leads to a 0.09 to 0.11 standard deviation shock on R_i, and a one-standard deviation shock on country risk factors leads to a 0.21 to 0.64 standard deviation shock on R_i.

Fundamentals such as size and price-to-book affect returns of both Turkish stocks and those traded in the countries forming group 2. However, there is a significant difference between the impacts of country risk factors in Turkish and group 2 stocks. For instance, while Turkish stocks are sensitive to all country risk factors, group 2 is mostly affected by only two factors. More specifically, listed from most to least important based on the standardized coefficients, Turkish stocks are sensitive to factor 2 (economic conditions), factor 5 (international openness), factor 4 (investment profile), factor 3 (conflicts) and factor 1 (social tensions). Group two is mostly affected by factor 1 (social tensions) and factor 2 (economic conditions), in that order.

In sum, we have identified that size, price-to-book value, and mostly all country risk factors affect Turkish stocks. On the other hand, group 2 stocks are equally affected by fundamentals and country risk factors. These findings are in accordance with our earlier discussion in the introduction — i.e., EU concerns of Turkey being politically too unstable, and financially and economically too underdeveloped, may be warranted.

CONCLUSION

Turkey formally became a candidate for EU membership on December 10, 1999. The hope of joining the EU has driven major reforms in Turkey, including economic liberalization, human rights protection, independence of the judiciary system, as well as economic and financial reforms leading to reduced hyperinflation, a more fairly valued currency, and lower interest rates. Nevertheless, the EU still perceives Turkey as politically too unstable, and financially and economically too underdeveloped to become a member. Turkish general opinion widely believes that the underlying reason for rejection by the EU is cultural differences rather than economic, political and financial weaknesses.

This paper attempts to explore the issue by identifying the risks associated with investing in the Turkish stock market. We first compare the source of excess volatility found in a portfolio of Turkish firms with those of new EU members which shared the same candidacy period as Turkey. We find that Turkish firms are indeed much more volatile and that the excess volatility can be significantly associated with higher financial and economic risks rather than fundamentals. We further investigate whether the differences in fundamentals and composite ratings between Turkey and a comparison group explain the excess volatility observed in the Turkish market as compared to other markets. The findings confirm that excess volatility is indeed driven by differences in relative financial and economic development rather than firms' fundamentals.

We further investigate the difference in sensitivity fundamentals and composite ratings between Turkey and a portfolio containing 176 stocks traded in the 8 other markets during their common candidacy period for EU membership. The first interesting finding is that firms' fundamentals are as important as country risk factors in explaining stock risk premiums for group 2, while the combined effect of country risk scores has a greater impact on risk premiums than firms' fundamentals alone in Turkey. While Turkish stocks are sensitive to all country risk factors — economic conditions, international openness, investment profile, conflicts, and social tensions — group 2 is mostly affected by only two factors, namely social tensions and economic conditions.

In sum, we have identified that size, price-to-book value, and mostly all country risk factors affect Turkish stocks. On the other hand, group 2 stocks are equally affected by fundamentals and country risk factors. These findings are in accordance with our earlier discussion in that the EU concerns of Turkey being politically unstable, and financially and economically less developed may be warranted. Overall, Turkey has been quite successful at pursuing reforms since commencing its candidacy for EU membership. It has liberalized its political system and relaxed restrictions on freedom, reduced hyperinflation, strengthened its currency, lowered interest rates, and provided a more stable growth in GDP. However, political, financial, and economic instabilities appear to be dominant issues throughout the study period.

ENDNOTES

1. Because the profit motive induces participants in financial markets to use every piece of readily available data to infer the current state of the economy, a capital market reflects investors' best collective forecast of future profits and is forward-looking on the economy. In fact, Fama (1990), Fama and French (1989), Hamilton and Lin (1996), Schwert (1990), Steven and Robert (1998), and Whitelaw (1994), among many others, find evidence of systematic movements in excess stock returns that are related to estimates of the underlying state of the business cycle — i.e., contractions in the stock market usually begin months before an economic recession and end before the trough and, therefore, anticipate the economic recovery. That is, stock market fluctuations lead the business cycle and are driven by expectations about changes in future economic activity.

2. EMDB does not represent a random sample of firms; thus, a selection bias can be seen in favor of larger and more actively traded firms (Rouwenhorst, 1999).

3. This is the same risk provider used and recommended in Erb, Harvey and Viskanta (1995, 1996a, 1996b and 1998) and Girard and Omran (2006). Indeed the authors examine many providers of country risk data (Bank of America World Information Services, Business Environment Risk Intelligence, Control Risks Information Services, the Economist Intelligence Unit, Euromoney, Institutional Investor, S&P Rating Group, the ICRG, Coplin-O'Leary Rating System, and Moody's Investors Services) and conclude that only the ICRG composite, political, financial and economic risk scores contain information that explains index returns.

4. The Turkish IFCG index is value-weighted and intended to represent a target of 60 to 75 percent of a country's total market capitalization and an industrial composition similar to that of the overall market.

5. This is the standard procedure used to estimate deflated excess return and has been used by a number of researchers including Kraus and Litzenberger (1976) and Fang and Lai (1997).

6. The monthly returns are estimated by dividing the difference in the price (or index values) between two consecutive months by the first month i.e., $R_{i(\text{or } m).t} = \left(P_{i(\text{or } m).t} - P_{i(\text{or } m).t-1}\right) / \left(P_{i(\text{or } m).t-1}\right)$

7. Rather than using the value weighted indices (IFCG indices) provided by EMDB, we build an equally weighted index for each country by averaging, each month, the returns for the stocks available. Indeed, the database is already biased towards large and liquid firms and, more generally, value weighted indices of emerging markets are more likely dominated by a few very large stocks. Thus, the use of an equally weighted index can minimize this size bias.

8. Girard and Omran (2006) show that (i) risk scores include information that cannot be aggregated in a composite measure and (ii) some risk measures have a greater bearing on business or investments, and a composite risk rating should place greater weight on those factors.

9. These have been suggested by Chen (1983) and used by Girard and Omran (2006) among others.

10. ICRG risk ratings consist of 12 political risk ratings, 5 financial risk ratings, 5 economic risk ratings, and 4 composite risk ratings. The second column of the table indicates the scale of each measure — i.e., 'bureaucracy quality' rating is out of 4, 'government stability' is out of 12, and so on. All political risk ratings scales add up to 100 possible points. The 'composite political risk' rating is then computed by summing all 12 individual political risk ratings. It has a maximum of 100 points. In the same vein, financial and economic risk ratings add up to 50 possible points, and their sum constitutes the composite financial and economic risk ratings. The composite rating (out of 100 points) is half of the sum of the composite political (100 possible points), financial (50 possible points) and economic risk ratings (50 possible points).

11. These are all stocks available in EMDB in each of the markets mentioned above from 1999:12 to 2004:06. There are no data available for the Greek part of Cyprus and for Malta, so they have been excluded from the analysis.

12. The Kaiser-Meyer-Olkin test (KMO) value for the sample is high (0.8545) and the Barlett test of sphericity is significant at the 1% level, indicating that the factor analysis is an appropriate technique for our data.

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AN ANALYSIS OF THE INTRA-REGIONAL TRADE IN THE MIDDLE EAST AND NORTH AFRICA REGION

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ABSTRACT

This paper analyzes the intra-regional trade and investment flows in the Middle East and North Africa (MENA) region using an augmented gravity model applied to panel data. The study uses annual trade and investment data for the period 1980-2006. There is a growing awareness among countries in the MENA region regarding the importance of international trade and foreign direct investment for stimulating growth and integrating into the world economy. The research will attempt to achieve the following objectives: (a) analyze the intra-regional trade and investment flows in the MENA region; (b) identify the major determinants of trade and investment flows in the MENA region using an augmented gravity model applied to panel data; and (c) measure the effect of preferential trading arrangements in the region on members' trade and investment with other MENA countries.

JEL: F14, F21

INTRODUCTION

The Middle East and North Africa (MENA) region is an economically diverse region that includes countries with a common heritage, shared religion, culture, and language, at various stages of economic development, vastly different levels of per capita income, and with very different endowment of natural resources. As Bolle (2006) points out, the countries in the MENA region are divided into four subgroups based mostly on geographical place and production foundation specifically the Maghreb countries (Algeria, Libya, Mauritania, Morocco and Tunisia), the Gulf Cooperation Countries (GCC) (Bahrain, Kuwait, Oman, Qatar, Saudi Arabia and United Arab Emirates), the Mashreq countries (Egypt, Jordan, Lebanon, Syria and Sudan), and other countries (Djibouti, Somalia, and Yemen). Arab countries possess various connections comprising shared religion, culture and language. Conversely, they are distinct in terms of size, crude source endowments, and standard of living (Al-Atrash and Yousef, 2000). Several are mainly farming and rural countries (Mauritania and Sudan), others are chiefly energy creators (members of the Gulf Cooperation Council (GCC), and others hold a promising and rising industrial foundation (Egypt and Morocco).

Trade policy has frequently been mentioned as the major policy provoked barrier to intra- MENA trade. Even as several countries in the area, particularly the GCC countries, sustain a moderately open trade regimes, others have faced considerable impediments to trade. Still several countries utilize a range of procedures comprising of restraining licensing, embargos and sanctions, state trading/monopolies, restraining foreign exchange provision and multiple exchange rates, to depress imports (Al-Atrash and Yousef, 2000).

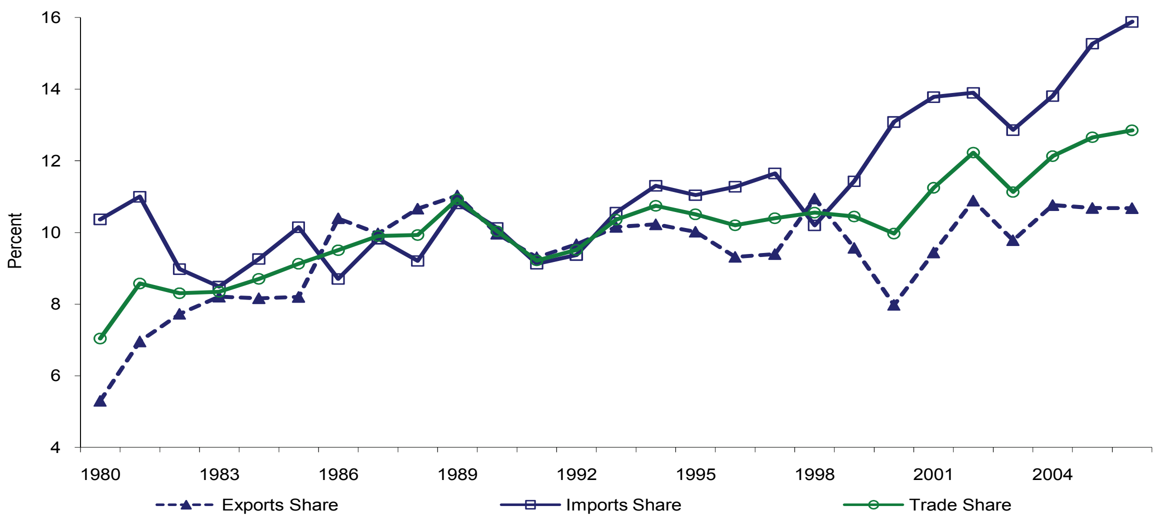
The degree of regional integration through trade in the Middle East and North Africa has been rising fast over the last twenty years. However, in 2006, inter-regional trade share in the Middle East and North Africa 12.9% was much lower than the European Union's share of 67.1% and of 55.2% for the North American Free Trade Agreement. In the same year, the intra-regional export share and import share was 10.7% and 15.9%, respectively (see Table 1 and Figure 1). Foreign direct investment flows in the MENA region have also remained relatively low, as Table 2 illustrates. Regardless of the low level of trade and investment flows in the region, it is interesting to find out reasons why they have remained at low levels.

Table 1: Direction of Trade of MENA Countries, 1980-2006

	Trade Share (%)				Export Share (%)				Import Share (%)			
Group	1980	1990	2000	2006	1980	1990	2000	2006	1980	1990	2000	2006
Industrial Countries	74.3	64.5	56.8	49.0	75.0	62.6	55.1	49.3	72.8	66.9	59.4	48.4
Developing Countries	24.2	30.9	39.1	46.7	23.5	32.2	39.2	43.9	25.5	29.3	38.9	50.6
Africa	1.3	2.1	2.9	2.9	1.4	2.3	3.4	3.6	1.2	1.8	2.0	1.8
Asia	10.6	12.0	22.2	26.1	11.7	13.4	25.8	28.1	8.4	10.3	16.6	23.2
Europe	2.9	5.7	3.5	4.8	1.8	5.5	2.3	2.5	5.1	6.1	5.4	7.9
Middle East	6.5	8.7	8.5	11.3	4.9	8.3	6.2	8.7	9.7	9.2	12.0	14.9
Western Hemisphere	2.9	2.3	2.0	1.7	3.8	2.7	1.4	1.0	1.2	1.9	2.9	2.7
European Union	42.6	38.8	31.5	27.0	39.8	34.1	26.5	22.3	47.8	44.6	39.2	33.6
Non-Oil Developing Countries	19.5	23.8	31.5	38.1	20.4	26.0	34.0	37.8	17.7	21.0	27.7	38.6
Oil Exporting Countries	4.7	7.1	7.6	8.6	3.1	6.1	5.2	6.2	7.7	8.2	11.3	12.0
Middle East and N. Africa	7.0	10.0	10.0	12.9	5.3	10.0	8.0	10.7	10.4	10.1	13.1	15.9
Total	100	100	100	100	100	100	100	100	100	100	100	100

This table shows the direction of trade of the MENA *region during the period 1980-2006. The region still trade mostly with industrial countries, although its trade share has dropped significantly between 1980 and 2006. The figures were taken from the IMF, Direction of Trade Statistics Database.*

Figure 1: Share of Intra-MENA Trade 1980-2006



This figure shows the trends in export share, import share and trade share of the MENA *region during the period 1980-2006. Import share remained consistently above the export and trade shares during this period.*

There is a growing awareness among countries in the MENA region regarding the importance of international trade and foreign direct investment for stimulating growth and integrating into the world economy. The research will attempt to achieve the following objectives: (a) analyze the intra-regional trade and investment flows in the MENA region; (b) identify the major determinants of trade and investment flows in the MENA region using an augmented gravity model applied to panel data; and (c) measure the effect of preferential trading arrangements in the region on members' trade and investment with other MENA countries. Although there are a few studies that analyze the intra-Arab trade, there are no studies to our knowledge that analyze both the trade and investment flows among MENA countries. This study, thus, will contribute the empirical literature on intra- MENA trade and investment.

Table 2: Stock of Intra-MENA Foreign Direct Investment, 1985-2006

Source Country	Host Country										
	Algeria	**Bahrain**	**Egypt**	**Jordan**	**Kuwait**	**Lebanon**	**Saudi Arabia**	**Syria**	**Tunisia**	**UAE**	**Total**
Algeria		20.1	337.1	160.6	26.3	11.2	42.5	30.4	33.1	28.3	751.2
Bahrain	36.0		126.1	4.4	180.8	51.8	1,459.0	0.1	17.6	59.9	1,826.5
Egypt	142.8	56.7		303.1	1,169.2	83.3	2,137.7	133.1	70.2	221.8	5,580.2
Jordan	34.7	16.5	75.6		13.5	153.7	3,011.1	70.4	1.0	11.8	3,956.5
Kuwait	315.5	232.4	242.4	147.2		2,241.8	63.5	482.9	170.1	428.0	4,028.5
Lebanon	4.5	21.5	272.5	6.8	683.6		1,621.6	204.5	0.7	804.4	3,986.0
Libya	19.1	-	22.7	-	-	1.8	64.1	-	19.0	98.7	334.9
Morocco	16.1	65.8	4.2	2.4	60.0	8.7	177.9	26.9	64.4	76.8	469.3
Oman	-	65.2	70.5	12.1	3.4	4.3	35.0	1.2	-	76.5	205.8
Qatar	14.6	2.8	10.8	8.7	107.7	556.3	65.9	2.3	-	78.7	1,044.4
Saudi Arabia	18.3	309.0	551.8	1,139.3	153.3	1,699.4		1,460.0	33.6	868.0	10,063.2
Sudan	2.8	-	69.7	182.1	88.4	65.3	1,586.0	180.2	0.6	209.3	2,564.9
Syria	4.7	21.7	28.9	18.1	338.3	346.1	1,187.5		5.7	375.4	3,154.3
Tunisia	17.5	2.8	5.0	136.0	359.1	8.3	425.3	4.0		144.2	1,599.8
UAE	9.7	50.9	2,963.3	82.7	362.1	1,585.5	35,457.5	640.1	2,328.8		43,575.8
Yemen	-	0.2	50.5	26.6	3.5	12.3	341.2	10.6	1.8	15.5	691.4
Total	637.2	865.6	4,866.7	2,552.3	3,549.4	6,829.9	47,939.3	3,281.5	2,750.3	3,573.2	84,756.5

This table shows the stock of intra-MENA investment flows during the period 1985-2006. There is no clear pattern of the investment flows. The figures were taken from the Inter-Arab Investment Guarantee Corporation, Investment Climate in Arab Countries 2006.

LITERATURE REVIEW

This section summarizes the previous studies that used gravity model to estimate the trade flows. For a more detailed literature review, the reader is directed to any of a number of surveys of various approaches, including Panagariya (1999, 2000), DeRosa (1998), Harrison, Rutherford and Tarr (2003), Robinson and Thierfelder (2002), Evenett and Keller (2002), Feenstra, Markusen, and Rose (2001), Scollay and Gilbert (2000), and Lloyd and MacLaren (2004).

The popularity of the gravity model is relatively recent. It was used during the 1960s and 1970s to estimate trade flows but was criticized because it lacks a strong theoretical foundation. Tinbergen (1962), Poyhonen (1963), and Linneman (1966) provided initial specifications and estimates of the determinants of trade flows. Anderson (1979) provided a rigorous economic justification, deriving a reduced-form gravity equation from a general equilibrium model incorporating the properties of expenditure systems. Bergstrand (1985) and Deardorff (1997) also provided partial theoretical foundations for the gravity equation, although none of the models generated exactly the same equation generally used in empirical work.

Due to a revival of interest among the economists about economics and geography, the gravity model has again become popular. A study by Egger (2008), using a partial equilibrium gravity model of bilateral trade, tests the role of distance on trade. He estimates a gravity equation from a large panel data-set of trade flows comprising all available bilateral export data from the United Nations' Comtrade database in the years 1980, 1990 and 2000. The analysis obtains three implications regarding the empirical specification of trade frictions in gravity models. The study concludes that distance became relatively more important in the two decades after 1980.

Instead of reviewing the recent studies that use the gravity model of trade, we summarize the findings of some recent studies on intra-Arab trade and investment.In a study by Bolbol and Fatheldin (2006) on intra-Arab investment flows, it was concluded that, although direct private flows increased, they were insufficient either to offset the fall in official flows or to boost intra-Arab trade. They highlight the importance of improving the Arab investment environment, particularly in those Arab countries having an investment or resource gap. They also point out that although most capital flows now originate from private sources and are made up of direct investments, the Arab countries continue to be recipients of disproportionately small capital flows relative to their size in the world economy.

Söderling (2005) analyzes export performance in the Middle East and North Africa (MENA) using a gravity model applied to panel data to addresses two questions: (i) are there significant unexploited export markets for the MENA region?; and (ii) have integration efforts with the EU since the mid-1990s yielded positive results? The results of the study suggest that several MENA countries are substantially underexploiting the United States as an export market. Moreover, the impact of integration efforts with the European Union has been moderate overall but significant in individual cases.

Al-Atrash and Yousef (2000) analyze the intra-Arab trade flows using an international dataset on bilateral trade for 65 countries in the 1990s. They estimate a gravity model to address the question of whether intra-Arab trade is too little. Their results suggest that intra-Arab trade and, more broadly, Arab trade with the rest of the world are lower than what would be predicted by the gravity equation, suggesting greater scope for regional integration as well as multilateral integration especially with the European Union. The results also suggest that intra-GCC and intra-Maghreb trade are relatively low while the Mashreq countries exhibit higher level of intra-group trade.

METHODOLOGY AND DATA

Methodology

This study analyzes the trade and investment flows in MENA region. The analytical tool used for this purpose is the standard gravity model of bilateral merchandise trade that has been widely used as the 'workhorse' for empirical analysis of international trade flows. Gravity models were first introduced to economic theory in the 1960s. Linneman's (1966) seminal study applied a gravity model to analyze the factors that explain trade for a sample of 80 countries. The standard gravity model postulates that trade between two countries is a function of their economic size and of the geographic distance between them. Gravity models have been augmented with variables representing factors that could either facilitate of impede trade. We augment this basic structure by adding a number of explanatory variables drawn from the theory of international trade. Since the gravity models have been used extensively to analyze the trade flows, the authors do not plan to discuss the theoretical development of gravity models. An interested reader is directed to any of a number of previous studies that used gravity models, including Panagariya (1999, 2000), DeRosa (1998), Harrison, Rutherford and Tarr (2003), Robinson and Thierfelder (2002), Scollay and Gilbert (2000), and Lloyd and MacLaren (2004).

This paper follows numerous authors and specifies the following gravity equation which controls for the basic determinants of international trade and investment. We also replaced population variable by gross domestic product (GDP), since either one can be used to measure the size of the economy. Our specification of the gravity models are:

$$\begin{aligned}\ln(T_{ij}) = {} & \beta_0 + \beta_1 \ln(PCGDP_i) + \beta_2 \ln(PCGDP_j) + \beta_3 \ln(GDP_i) + \beta_4 \ln(GDP_j) + \beta_5 \ln(Dist_{ij}) + \beta_6 \ln(RER_{ij}) \\ & + \beta_7 Border + \beta_8 Language + \beta_9 Maghreb + \beta_{10} GCC + \beta_{11} Mashreq + \beta_{12} PETRO + u_{ij}\end{aligned} \tag{1}$$

$$\begin{aligned}\ln(FDI_{ij}) = {} & \beta_0 + \beta_1 \ln(PCGDP_i) + \beta_2 \ln(PCGDP_j) + \beta_3 \ln(GDP_i) + \beta_4 \ln(GDP_j) + \beta_5 \ln(Dist_{ij}) + \beta_6 \ln(INF_j) \\ & + \beta_7 \ln(EXP_{ij}) + \beta_8 \ln(RER_{ij}) + \beta_9 Border + \beta_{10} Language + \beta_{11} Maghreb + \beta_{12} GCC + \beta_{13} Mashreq + u_{ij}\end{aligned} \tag{2}$$

where T_{ij} represents the flow of trade from country i to country j; FDI_{ij} represents the flow of foreign direct investment from country i to country j; $PCGDP_i$ is the per capita gross domestic product of country i, $PCGDP_j$ is the per capita gross domestic product of country j; GDP_i is the gross domestic product of country i; GDP_j is the gross domestic product of country j; $Dist_{ij}$ is the geographical or economic distance between the two countries; RER_{ij} is the real exchange rate between the two countries; INF_j is the inflation rate in country j; EXP_{ij} is the exports from country i to country j; *Border* is a dummy variable which takes the value 1 if the two countries share a contiguous border and zero otherwise; *Language* is a dummy variable which takes the value 1 if the two countries share a common language and zero otherwise; *Maghreb* is a dummy variable that equals 1 if the two countries are members of the Maghreb and zero otherwise; *GCC* is a dummy variable that equals 1 if the two countries are members of the Gulf Cooperation Council (GCC) and zero otherwise; *Mashreq* is a dummy variable that equals 1 if the two countries are members of the Mashreq and zero otherwise; *PETRO* is a dummy variables that equals 1 if country i is a petroleum exporting country; and u_{ij} is a normally distributed error term.

According to Frankel (1993), per capita GDP is included to capture the factors associated with the level of economic development. Other authors have also used per capita income to express the level of economic development (see, for example, Carrillo and Li (2002)). Per capita GDP also captures the productive capacity of the exporting country and the purchasing power of the importing country. The coefficients of the per capita GDP variables are expected to be positive.

Gross domestic product variables represent the size of the countries and are expected to have positive signs. According to Krugman (1980), the larger countries are better able to absorb imports than smaller countries and are better able to experience economies of scale and thus develop a comparative advantage in their export industries than are smaller countries. The size of GDP can also be treated as a proxy for market thickness (the economic depth of trading nations) which positively impacts on the location of outsourcing activity (Grossman and Helpman, 2005).

The coefficient of the distance variable ($Dist_{ij}$) is expected to be negative. This is a proxy for transportation costs and time, access to market information, access to markets, and other factors that make it difficult for nations to engage in trade.

Following Pozo (1992), the bilateral real exchange rate, RER, was constructed as,

$$RER_{ij} = \frac{ER_{ij} \times CPI^{j}}{CPI^{i}} \tag{3}$$

where RER_{ij} is the real exchange rate between country i to country j, ER_{ij} is the nominal exchange rate (the home currency price of a unit of foreign currency, for example, the number of Rials per US \$), CPI^i is the consumer price index (2000=100) of origin country i and CPI^j is the consumer price index (2000=100) for a given foreign (destination) country j. The data on nominal exchange rates and CPI were taken from the International Monetary Fund, *International Financial Statistics database*. The coefficient of the RER_{ij} variable is expected to be positive.

The anticipated sign on all dummy variables is positive, reflecting the idea that proximity, common language, historical links, and regional trading agreements are trade creating networks. A common border dummy (*Border*) is included to account for possible additional advantages of proximity that are not captured by the standard distance measure. Common language tends to facilitate trade by enhancing exporters' and importers' understanding of each others' cultures, commercial and legal systems, which have a great deal of influence on trade. Growing empirical literature finds that historical linkages are important determinants of international trade flows (see, for example, Frankel, Stein and Wei (1995), Frankel (1997), and Eichengreen and Inrwin (1998)).

Data Sources

We estimate the models with annual data for 20 MENA countries for the period 1980 to 2006. Algeria, the Kingdom of Bahrain, Djibouti, Egypt, the Islamic Republic of Iran, Jordan, Kuwait, Lebanon, Libya, Mauritania, Morocco, Oman, Qatar, the Kingdom of Saudi Arabia, Somalia, Sudan, Syrian Arab Republic, Tunisia, the Kingdom of United Arab Emirates, and the Republic of Yemen. The dependent variables used in the analysis is exports from country i to country j and foreign direct investment from country i to country j. The data on exports and imports for the study period of 1980-2006 are from the International Monetary Fund, *Direction of Trade Statistics* database. Additional data on exports and imports are from the United Nation's *Commodity Trade Statistics* (Comtrade) database. Data on population are from International Monetary Fund, *International Financial Statistics Yearbook*. Information on per capita gross domestic product is from International Monetary Fund, *World Economic Outlook Database*, April 2008. The distance variable is obtained from the World Bank, *Trade, Production, and Protection* database. The data on foreign direct investment are from the Inter-Arab Investment Guarantee Corporation and from the UNCTAD's *World Investment Report 2007*.

EMPIRICAL RESULTS

Trade Flows

We estimated two sets of regression models to measure the fixed-effects and random-effects. The estimated results of the model analyzing intra- MENA trade flows are presented in the first three columns of Table 3. With twenty countries, where each of them has nineteen country-pairs, our sample has 380 observations per year and (380 observations x 27 years = 10,260) 10,260 observations for the full sample. The model specification test performed using the Hausman test rejected the random effects model specification. As a result, the results are discussed using the estimated results of the fixed-effects model, although the results of the random-effects models are also presented in Table 3. The conventional variables behave very much the same way as the model predicts, and the estimated coefficients are statistically significant. The adjusted R^2 value for the fixed-effects model is 0.515. This value is acceptable for a cross-sectional study and is comparable to those obtained in other studies employing the gravity model to examine intra-regional trade flows.

The coefficients of the GDP variables are positive and highly statistically significant, indicating that size of the economies play an important role in intra- MENA trade flows. However, the coefficient of the exporting country is relatively larger than that of the importing country. The coefficients of the per capita income variables are both negative, though they are expected to be positive. They are also statistically significant at the 1% level of significance. The distance variable also has the expected negative sign and is highly significant.

The real exchange rate variable has the expected positive sign and it is statistically significant at the 1 percent level of significance. This result is comparable to the findings of other studies on the impact of real exchange rate on exports. The Border dummy variable has the expected positive sign and is statistically significant. However, the border effect in the case of MENA trade flows is relatively low. Generally the border effect is estimated by the border dummy coefficient in a regression equation. Since all variables in the model, except the dummy variables, are in logarithm, the border effect is calculated as the anti-log of the border dummy coefficient. In the estimated model, border dummy coefficient is 0.305 for the fixed-effects model. Therefore, the border effect is [exp(0.305) =] 1.36. This value indicates that countries sharing a common border in the MENA region, on average, tend to have 36% more trade compared with countries with no common borders. This result is similar to the finding of the study by Söderling (2005) on trade flows in MENA region. However, Helliwell (1996, 1998) and McCallum (1995) estimate the border effect to be around 20 in Canada-US trade, indicating that there will be 20 times more trade among states/provinces that share a common border.

The language dummy also has the expected positive sign. The common language variable has more effect on trade than the amount of trade when two countries share a common border. Two countries with a common language in MENA region tend to have 4.4 times more than two countries with different languages. Common language in the MENA region tends to facilitate trade by enhancing exporters' and importers' understanding of each others' cultures, commercial and legal systems. Similarly, colonial past also tends to have a positive and statistically significant effect on trade flows in MENA region. Two countries in the region with past common colony appear to have 3 times more trade than two countries with different colonial history.

The dummy variables for membership in a trade preference scheme give mixed results. Membership in GCC tends to have a positive effect on trade flows while memberships of *Maghreb* or *Mashreq* tend to have negative and significant effect on trade flows in the region. This finding is also consistent with the finding of the study by Al-Atrash and Yousef (2000) on intra-Arab trade.

Finally, the dummy variable representing whether or not the origin country is a petroleum exporting country also has a negative and statistically significant effect on intra-MENA trade flows. Petroleum exporting countries in the region tend to trade about 68% less with the countries in the region compared with non-petroleum exporting countries' trade with the countries in the region.

<u>Investment Flows</u>

The estimated results of the model analyzing intra-MENA investment flows are presented in the last two columns of Table 3. Unlike trade statistics, investment statistics are not available for the entire period under study. We were able to find investment data only for the period 1985-2006. Due to the limitation of data, we took the average level of foreign direct investment flows during this period. As a result, we only have 306 observations for the full sample. The model was estimated using the Ordinary Least Squares (OLS) estimation method. The adjusted R^2 value for the estimated model is 0.570, which is slightly higher than that of the trade model. The coefficients of the GDP variables are positive and highly statistically significant, indicating that size of the economies play an important role in intra-MENA investment flows. The coefficients of the per capita income variables are also both positive, though they were both negative in the trade model. However, only one of them is statistically significant at the 10% level of significance. The distance variable has the expected negative sign but it is significant only at the 10% level of significance.

The exports variable has the expected positive sign and it is statistically significant at the 1 percent level of significance. This result indicates that higher levels of exports tend to go with higher levels of foreign direct investment. The rate of inflation in the destination country has a positive sign and it is not statistically significant. The real exchange rate variable has the expected positive sign but it is statistically insignificant. The Border dummy variable also has the expected positive sign but it is statistically insignificant. The border effect in the case of MENA investment flows is relatively low. This value indicates that countries sharing a common border in the MENA region, on average, tend to have

17% more investment compared with countries with no common borders. The language dummy also has the expected positive sign though it is statistically insignificant. Finally, the dummy variables for membership in a preferential trade scheme give mixed and significant results. Membership in *Mashreq* tends to have a positive effect on investment flows while memberships of *Maghreb* or GCC tend to have negative and significant effect on investment flows in the region.

Table 3: Determinants of Trade and Investment Flows in Middle East and North Africa

	Intra-MENA Trade Flows		**Intra-MENA Investment Flows**	
Variable	**Fixed Effects**	**Random Effects**	**Variable**	**OLS Estimates**
Constant	16.948*	15.973*	*Constant*	-7.861
	(29.56)	(31.32)		(-1.48)
$\ln(GDP_i)$	1.479*	1.522*	$\ln(GDP_i)$	1.068*
	(44.15)	(52.23)		(4.34)
$\ln(GDP_j)$	1.151*	1.177*	$\ln(GDP_j)$	1.363*
	(49.21)	(58.74)		(7.10)
$\ln(PCGDP_i)$	-0.153*	-0.043	$\ln(PCGDP_i)$	0.116
	(-3.43)	(-1.26)		(0.30)
$\ln(PCGDP_j)$	-0.432*	-0.439*	$\ln(PCGDP_j)$	0.534**
	(-13.79)	(-12.07)		(2.26)
$\ln(Dist_{ij})$	-2.781*	-2.764*	$\ln(Dist_{ij})$	-0.861***
	(-44.07)	(-56.36)		(-1.87)
$\ln(RER_{ij})$	0.034*	0.041*	$\ln(Expo_{ij})$	0.355*
	(3.74)	(4.15)		(3.76)
Border	0.305*	0.240**	$\ln(INF_j)$	0.041
	(2.89)	(2.35)		(0.26)
Language	1.471*	1.439*	$\ln(RER_{ij})$	0.066
	(12.99)	(14.02)		(1.19)
Colony	1.080*	1.094*	*Border*	0.161
	(18.88)	(18.66)		(0.32)
Maghreb	-0.828*	-0.927*	*Language*	0.504
	(-6.15)	(-9.37)		(0.53)
GCC	0.110	0.308*	*Maghreb*	-1.307***
	(0.81)	(2.62)		(-1.73)
Mahreq	-0.803*	-0.747*	*GCC*	-1.776**
	(-4.82)	(-5.54)		(-2.19)
Petro	-0.961*	-1.210*	*Mahreq*	1.663***
	(-11.68)	(-18.46)		(1.73)
Adjusted R^2	0.515	0.386	*Adjusted* R^2	0.570
Observations	10,260	10,260	*Observations*	306
Hausman Test		196.74*		
Border effect	1.36	1.27	*Border effect*	1.17
language effect	4.35	4.22	*language effect*	1.66
Colony effect	2.95	2.99		

*This table shows the empirical results of the trade and investment models, as given in equations (1) and (2). The first three columns of the table show the empirical results of the trade model while the last two columns show the empirical results of the investment model. *, **, and *** indicate the significance at the 1%, 5%, and 10% level, respectively. Figures in parentheses are t values.*

SUMMARY AND CONCLUSIONS

This paper analyzes the intra-regional trade and investment flows in the Middle East and North Africa (MENA) region using an augmented gravity model applied to panel data. The study uses annual trade and investment data for the period 1980-2006. Employing the gravity model in the analysis of intra-regional trade and investment flows in MENA reveals some interesting observations concerning the Middle Eastern and North African trade and integration arrangements, such as the importance of language and culture as determinants of trade and investment. The findings of this study are, for the most part, are consistent with findings of previous studies on the Middle Eastern and North African trade and investment flows. The coefficients of per capita GDP, population, and distance had expected signs and magnitudes in all models estimated. This confirms the results of other studies. The border effect is relatively smaller in the Middle Eastern and North African region, relative to the regions such as North America and Europe. For example, Helliwell (1996, 1998) and McCallum (1995) estimate the border effect to be around 20 in Canada-US trade, indicating that there will be 20 times more trade among states/provinces that share a common border while this study finds border effects to be only 1.4.

The flow of foreign direct investment to the $MENA$ region continues to be very low, despite the more than two decades of fiscal reforms. The major policy implication of this paper is that $MENA$ countries should be more open in order to attract foreign direct investment. The rapidly evolving economic and political climates in the region provide many opportunities for the investigation of the success of economic integration in the Middle East and North Africa.

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TOP MANAGEMENT COMPENSATION, EARNINGS MANAGEMENT AND DEFAULT RISK: INSIGHTS FROM THE CHINESE STOCK MARKET

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ABSTRACT

China has sustained a rapid rate of economic growth and absorbed a great deal of foreign investment over the past decades. However, the laws pertaining to business in China have not kept up with China's market growth. For this reason, investors in the Chinese stock market must assess associated risks. We set out in this study to examine the relationships that exist between default risk, earnings management, and top management compensation of publicly-listed companies in the Chinese stock market, which is now considered the most important emerging market. The results reveal a greater likelihood of default amongst larger discretionary accruals and lower top management compensation. In addition to studying the relationships which exist in the full sample, we also divide the sample into two sub-groups, based upon the signs of discretionary accruals, to investigate the likelihood of default. We find higher default potential amongst firms only falling into the category of positive discretionary accruals.

JEL: G14; G34; G35; M41

INTRODUCTION

China has sustained a rapid rate of economic growth since the inauguration of its economic reform period. The trials and tribulations of the reform process have been well documented (Cao, Qian, and Weingast, 1999; Gao, 1996; Groves, Hong, McMillan, and Naughton, 1994; Lin and Zhu, 2001) and analyses of the effectiveness of these reforms have begun to appear in the literature (Allen, Qian, and Qian, 2005; Chen, Su, and Zhou, 1998). For the Chinese stock market, one of the most important policy changes was the rapid conversion of its socialist planned economy into a market economy. However, the majority shareholders for most of China's listed state-owned enterprises (SOE's) are state agencies that lack experience in monitoring and controlling public firms. Although government institutions are aimed at improving commercial legal and judicial system; however, China's commercial legal and judicial system are not as transparent as those in developed markets (Davidson, Gelardi, and Li, 2006; Winkle, Huss, and Chen, 1994; and Zhou, 1988). Therefore, the apparent lack of control and monitoring from investors, regulatory agencies, and the lack of transparency in financial disclosure have provided considerable discretion for managing earnings (Aharon, Lee and Wong, 2000).

Accrual accounting is the standard practice for most companies, since it provides an accurate description of the company's current condition. However, in addition to improving firms' financial reporting and disclosure, it can also provide opportunities for managers to use discretionary accruals to manipulate their financial statements. Therefore, many studies use discretionary accruals as a proxy for earnings management and suggested that managers have incentives to manipulate their firms' earnings in order to reduce the effect of negative signals (Sweeney, 1994; Burgstahler and Dichev, 1997; Jaggi and Lee, 2002). Koch (1981) indicated that firms with increased technical default probability had a propensity for making income-increasing accounting changes. Sloan (1996) posited that high-accrual firms would experience lower future earnings performance, while Peltier-Rivet (1999) and Jaggi and Lee (2002) indicated that the larger a firm's debt-equity ratio, the more likely the firm was to adopt income-increasing accounting procedures. However, there are less studies linking earnings management to the probability of default among firms. Hence, we chose to investigate this relationship in China in our study.

After the massive 1997 financial crisis, many Asian countries aimed to strengthen corporate governance, transparency and disclosure levels in order to decrease the managers to manipulate financial statement (Ho and Wong, 2001). Jensen and Meckling (1976) formalized the relationship between agency problems and top management compensation, indicating that higher top management compensation could mitigate agency problems. A large amount of empirical literature followed Jensen and Meckling's study, and found that companies with higher incentive compensation have better performance (Crystal 1991; Jensen and Murphy, 1990; Kaplan 1994; Patton and Baker 1987).

Jensen and Meckling (1976) indicated that good corporate governance procedures not only increase companies' performance but decrease the probability of default among firms. However, studies documenting the association between default risk and top management compensation are much rarer. Only Cyert, Kang and Kumar(2002) have explored the negative association between default risk and CEO compensation. The corporate scandal in the emerging markets is more which have arisen. In China, Sun and Zhang (2006) pointed out that about 20 percent of publicly listed firms have been convicted by China Securities Regulations Committee (CSRC) for serious frauds and scandals since the Chinese stock market was established in the early 1990s. In our study, we extend Cyert et al. (2002) understood how giving top management higher compensation could effectively mitigate the default probability of firms in China. In addition, we also explore the possibility that higher top management compensation could decrease managerial use of discretionary accruals to manipulate financial statements and expand companies' default risk.

Although the practice of using discretionary accruals, either to increase or reduce income, is essentially aimed at managing earnings, there are, nevertheless, differences in the attitudes of executives and the results of their financial statements. Indeed, Epps (2006) notes that income-increasing and income-reducing discretionary accruals have different relationships with corporate governance practices. For the purpose of consistency and in order to determine the impacts of finance-oriented earnings management, we will explore the linkage between discretionary accruals, top management compensation, and default risk, in terms of different discretionary accruals (positive vs. negative).

Our study makes several contributions to the literature in this field. Firstly, we provide evidence of the relationship between discretionary accruals (earnings management), top management compensation, and default risk, thus complementing the findings of prior studies. Secondly, we examine listed firms in the Chinese stock market to explore such relationships. It is common knowledge that investments in unsafe and deficient markets are accompanied by higher risk. Therefore, the relationship between top management compensation, earnings management behavior of firms, and default risk is of considerable interest to all investors. Finally, we also examine whether the relationship between top management compensation, discretionary accruals, and default risk is the same under different categorizations. The empirical results of all of these issues should help investors to make appropriate investment decisions.

We use the random effects panel regression model, as opposed to the ordinary least square (OLS) estimation in this study, since the panel data regression is able to supply more accurate inferences for the parameters and reduce any collinearity that may exist amongst the explanatory variables. Our results show that there is a greater risk of default where firms have greater positive discretionary accruals and less top management compensation. Furthermore, higher CEO compensation could mitigate the positive relation between discretionary accruals and firms' bankruptcy risk. That is to say, companies with higher CEO compensation have less incentive to manipulate financial statements, and thus, less risk of bankruptcy.

The remainder of this paper is organized as follows. Section 2 describes the literature review and develops our hypotheses. Section 3 describes the data sources and empirical methodology adopted for this study. Section 4 provides the descriptive statistics and presents the empirical results and analysis, with the final section summarizing the conclusions drawn from this study.

LITERATURE REVIEW AND HYPOTHESIS DEVELOPMENT

China Economic Environment

China has been transformed into the greatest economic society in the world and, thus, there are many studies which have examined topics relating to mainland China (Allen et al., 2005; Bailey et al., 2003; Kang et al., 2002; and Liu et al., 2002). Following this trend, many international investors have put a lot of capital into the Chinese stock market. In addition, many studies have showed spillover effects, indicating that international financial markets are becoming more affected by changes in the Chinese stock market. Therefore, the influence of the Chinese economy is gaining in importance (Bekaert and Harvey, 2000; Wang et al., 2002). However, the commercial legal/judicial system in China is not yet mature (Hamilton and Biggart, 1988; Whitley, 1994) and it is much easier for managers in Chinese firms to manipulate their earnings.

This unfamiliar legal environment with suspect financial systems gives rise to uncertainty and risk for the foreign investor. China, with its distinctive political, institutional, and cultural characteristics, may have to be analyzed using different methods (Hamilton and Biggart, 1988; Whitley, 1994). A growing body of literature has indicated China's emerging economic might not be analyzed properly in conventional Western terms (Aoki, 1984, 1990; Biggart and Hamilton, 1992; Boisot and Child, 1988; Goto, 1982). This provides evidence that foreign investors may have difficulty becoming familiar with the Chinese economic environment.

As has been seen, the problems of corporate governance in China are quite serious. Hence, we see the firms in China as an appropriate target for research as China may ultimately provide an example for other countries with incomplete legal and financial systems. Moreover, related findings will potentially provide ideas for the design of appropriate corporate governance practices within developing economies.

Default Risk

Many models attempting to forecast the probability of business failure have been developed throughout the years. The earliest and most-cited studies predicting the default probability of firms include the financial ratio analysis model of Beaver (1966) and the Z-score model of Altman (1968). These models are capable of providing accurate predictions of corporate bankruptcy, but have been subjected to numerous revisions. Altman (2000) provides a detailed description of the construction of the second generation credit risk model, and adds several enhancements to the original model.

There are many other approaches to the prediction of bankruptcy risk which attempt to overcome the shortcomings of the earlier models (Männasoo, 2007; Ohlson, 1980; Shumway, 2001; Walker, 2005). However, there has been a general tendency in these studies to use historical data to predict the default risk. This approach may not be capable of adequately reflecting the actual probability of bankruptcy in the changing market (Hillegeist et al., 2004). Furthermore, since a firm's risk of bankruptcy is correlated with the properties of its lines of business, and debt ratios cannot effectively mirror the actual probability of default. Thus, the use of historic ratios as a substitute for default risk would seem to be inappropriate.

Dacorogna, Oderda and Juna (2003) demonstrated that the forecasting ability and time-varying characteristics of the KMV model were superior to those of other models (The KMV model is detailed in Appendix A). Domingues (2004) also suggested that the KMV model seemed to be particularly appropriate for application to publicly-traded companies, since the equity values in this model are determined by the market. Apart from its ability to react appropriately to the condition of the firm, the KMV model can also provide important information for investors. Therefore, in an effort to overcome the shortcomings of the aforementioned models, we use the KMV model in this study to compute the default risk and to provide an optimal description of the probability of bankruptcy.

Earnings Management

Accrual accounting is the standard practice used to mitigate timing and matching problems inherent in cash flows so that earnings more closely reflects firm performance. Dechow (1994) indicated that accrual accounting gave management some discretion over the recognition of accruals. The discretion might be used by management to signal meaningful information that would otherwise not be communicated. Signaling was expected to improve the ability to measure firm performance from earnings since management presumably has superior information about their firm's cash generating ability (Healy and Palepu, 1993; Holthausen, 1990; Holthausen and Leftwich, 1983; Watts and Zimmerman, 1986). Therefore, a credible signal would reduce information asymmetry and result in more efficient contracting.

However, with the advantages of the accrual accounting there was introduced a set of problems, such as earnings management; accordingly, numerous studies use discretionary accruals as a proxy for earnings management. Watts and Zimmerman (1986) found that management used their information advantage to opportunistically manipulate accruals. Healy and Wahlen (1999) showed that managers used accruals to manage earnings and alter firms' reported economic performance to mislead outsiders such as investors, debt holders and government institutions. Thus, outsiders had a less reliable view of firm performance. In addition, many studies reported that financially distressed firms had more incentive to manage their earnings. Jaggi and Lee (2002) showed that managers of financially distressed firms often engaged in earnings management, either to deliver good news or to reduce the effects of bad news on their finances. Peltier-Rivet (1999) also argued that firms with a larger debt-equity ratio have a tendency to adopt income-raising accounting procedures. This suggests, therefore, that there is some relationship between earnings management and the probability of default. However, studies have rarely explored such an association. Accordingly, in this study, we will investigate this relationship in an attempt to provide evidence to supplement these studies. Thus, our first hypothesis is as follows:

> *Hypothesis 1:* Ceteris paribus, the relationship between discretionary accruals and default risk will be positive.

Top Management Compensation

A growing amount of empirical literature indicated that top management compensation was potentially linked to the effectiveness of governance mechanisms (Cheng and Firth; 2006; Cyert et al., 2002; Brunello et al., 2001; Firth et al., 2006; Izan et al., 1998). The indications were that better compensation plans give managers sufficient incentive to maximize shareholder wealth, decrease the motivation to manipulate financial statements, and reduce "agency problems". Furthermore, many studies provided empirical evidence of a positive relationship between top management compensation and company performance (Coughlan and Schmidt, 1985; Murphy, 1985, 1986; Jensen and Murphy, 1990; Abowd, 1990; Leonard, 1990; Kaplan 1994).

Jensen and Meckling (1976) proposed the "agency theory" which suggests conflicts of interest between various contracting parties such as shareholders, company managers, and debt holders. The conflicts of interest include both the maximization of corporate performance and the reduction of default risk of the firms. Many studies have found that conflicts of interest can be reduced by policies such as thorough top management compensation plans (Abowd, 1990; Leonard, 1990; Kaplan 1994). However, these studies have always focused on motivating managers to maximize corporate performance and have ignored the problem of enabling managers to reduce the probability of the default risk. Up to now, only Cyert et al. (2002) have attempted to construct a model concerning the relationship between top management compensation and the probability of default risk. They also provided empirical evidence indicating that top management compensation is negatively related to the firm's bankruptcy risk, to support their theory.

In our study, we follow up Cyert et al.'s (2002) study, investigating whether any relationship exists between top management compensation and default risk in publicly-listed firms in China. Our second hypothesis is therefore presented as follows:

> *Hypothesis 2:* Ceteris paribus, the linkage between top management compensation and default risk will be negative.

Furthermore, we also explore the possibility that higher top management compensation could decrease manager use of discretionary accruals to manipulate financial statements and expand companies' default risk. Our third hypothesis is therefore presented as follows:

> *Hypothesis 3:* Ceteris paribus, the relationship between compensation to encourage top management compared to discretionary accruals together with default risk will be negative.

DATA AND METHODOLOGY

Data Description

Our discussion of the relationship between default risk, discretionary accruals, and shareholder concentration is based upon data obtained from the Chinese Stock Market and Accounting Research (CSMAR) database. This sample is comprised of all publicly-listed enterprises in the Shanghai and Shenzhen Stock Exchange. As the China Securities Regulatory Commission (CSRC) requests all publicly-listed firms in these stock markets to compile their corporate governance data (such as top management compensation) and to compute discretional accruals using prior cash flow since 2001, our sample span covers the five-year period, from 2001 to 2005.

Only those companies which conform to our selection criteria have been used in our analysis. Firstly, we confined ourselves to firms that have their financial year-end in December of each year. This ensures that the information obtained from the financial statements is available each year. Secondly, we chose only those firms on which there is complete data (the book value of total debts and assets, the equity market value, the stock price volatility, and so on) covering the years 2001 to 2005, to fully satisfy the related computation associated with the KMV model. This selection process yields a total of 471 firms, and 2,355 firm-year observations.

Our sample is then sub-divided into groups based upon discretionary accruals (positive vs. negative) to determine whether the effects present in all groups are the same. In this way, we can more clearly understand influences in different situations.

Empirical Models

We employ the Multivariate Random Effects Balanced Panel Regression Method to examine the effects of discretionary accruals, top management compensation, and other financial variables on the default risk for publicly-listed firms in China. We begin by constructing an annual time series model of top management compensation, discretionary accruals, and corporate default risk using the KMV model to assess the firms' probability of default. The KMV model calculates the actual probability of default based upon the option pricing theory of Black and Scholes (1973) and Merton (1974). The computation of 'expected default frequency' (EDF) is based on the company's capital structure, the volatility of its asset returns, and the current asset value. The related process of deriving EDF is expressed in Appendix A.

Guided by related theories drawn from the aforementioned prior studies, the control variables are comprised of the debt ratio, the reciprocal of the current ratio, ROA, and total assets. Calendar year dummy variables are also included to specify the potential time effects on default risk. The empirical model is described as follows:

$$RISK_{it} = \beta_0 + \beta_1 DISACC_{it} + \beta_2 COMPEN_{it} + \beta_3 DISACC_{it} \times COMPEN_{it} + \beta_4 DEBT_{it} + \beta_5 ROA_{it} \\ + \beta_6 ASSET_{it} + \beta_7 DIRECT_{it} + \beta_8 Y02_{it} + \beta_9 Y03_{it} + \beta_{10} 04_{it} + \beta_{11} 05_{it} + \varepsilon_{it} \quad (1)$$

where $RISK_{it}$ is the i^{th} firm's default risk computed from the KMV model in year t; $DISACC_{it}$ represents the i^{th} firm's absolute discretionary accruals estimated from the modified Jones model in year t (the related process of deriving discretionary accruals is expressed in Appendix B); $COMPEN_{it}$ refers to the cash compensation of the three highest paid employees in the i^{th} firm in year t; $DEBT_{it}$ is the i^{th} firm's debt ratio in year t; ROA_{it} indicates the i^{th} firm's return on assets in year t; $ASSET_{it}$ expresses the i^{th} firm's log total assets in year t; $DIRECT_{it}$ indicates to the size of the board of directors in the i^{th} firm in year t; $Y02_{it}$, $Y03_{it}$ and $Y04_{it}$, $Y05_{it}$, are (0,1) dummy variables controlling for the effects of calendar years; if the data is extracted from year 2002, then $Y02_{it}$ is 1, otherwise 0, and so on for the years 2003-2005.

In conducting our study, we were interested in the relationships between discretionary accruals (earnings management), top management compensation, and default risk. The coefficient on the earnings management measure, *DISACC*, captures the connection between firms' manipulation of financial statements (discretionary accruals) and the probability of default risk. There has been widespread use of discretionary accruals as a substitute for earnings management in studies on the relationship between corporate governance and earnings manipulation (Defond and Jiambalvo, 1994; Epps, 2006; Teoh et al., 1998a, b). In view of this, we follow the example of Kothari et al. (2005); calculating the discretionary accruals whilst including a performance-measurement variable, return on assets (ROA), into the modified Jones model. Furthermore, it was argued by both Holthausen and Larcker (1996) and Teoh et al. (1998a) that different industries may have divergent levels of accruals. Our estimate of discretionary accruals therefore also takes into account different industrial affiliations. When the coefficient of *DISACC* is positive, the indication is that firms are more energetically involved in earnings manipulation and have an increased risk of bankruptcy, thus providing support for Hypothesis 1.

The coefficient of top management compensation, *COMPEN*, we use the compensation of the three highest paid employees as a proxy to find the link between top management compensation and the default risk of the firms. When the coefficient of *COMPEN* is negative, the indication is that firms with more compensation to top management have a decreased risk of bankruptcy, which provides support for Hypothesis 2. The coefficient on the interaction between the discretionary accruals and top management compensation (*DISACC*×*COMPEN*) captures the relationship between discretionary accruals, top management compensation, and default risk. If the coefficient of *DISACC*×*COMPEN* is negative, the indication is that firms which give more compensation to top management decrease the manipulation of financial statements and further decrease risk of bankruptcy, which provides support for Hypothesis 3.

EMPIRICAL RESULTS AND ANALYSIS

Summary Statistics

Our sample was comprised of 471 firms, giving a total of 2,355 firm-year observations from 2001 to 2005. Table 1a presents the descriptive statistics for the pooled sample of all firm-year observations. From Table 1a, we find that the mean of firms' probability of default risk (*RISK*) is 0.0060, which means firms' default risk in China is not very high. The mean of *DISACC* is 0.0538, which indicates on average firms' discretionary accruals comprise 5.38 percent of the prior total assets. On average *COMPEN* is 12.76, which is to say that every year firms' top management get 12.76 million cash compensation.

Different signs of discretionary accruals reflect the dissimilar administrative behavior of top management (Healy, 1985). Accordingly, we classify the data into two sub-groups based upon the sign of each firm's five-year average of discretionary accruals. The results in Table 1b show that the firms in the 'positive discretionary accruals' group would manipulate the accruals more than those in the 'negative discretionary accruals' group. In addition, the 'positive discretionary accruals' group has higher asset scale than the 'negative discretionary accruals' groups.

Table 1a: Summary of Descriptive Statistics for All Firms (N=2355)

Variable	Mean	Median	Std. Dev.	Min..	Max.
RISK	0.0060	7.33E-09	0.0708	0.0000	1.0000
DISACC	0.0538	0.0339	0.0715	0.0000	1.1862
COMPEN	12.7592	12.8213	0.8727	10.0750	16.0804
DEBT	0.5900	0.5461	0.9881	0.0273	43.0750
ROA	0.0168	0.0278	0.1152	-1.7491	1.0847
ASSET	14.2694	14.2589	0.9287	10.5763	17.8683
DIRECT	9.6400	9.0000	2.2700	0.0000	19.0000

Table 1b: Summary of Descriptive Statistics for All Firms, by Discretionary Accruals

Variable	Positive Discretionary Accruals (N=1103)			Negative Discretionary Accruals (N=1252)		
	Mean [b]	Std. Dev.	Median	Mean	Std. Dev.	Median
RISK	0.0059	0.0702	1.01E-08	0.0061	0.0713	4.90E-09
DISACC	0.0557 *	0.0816	0.0318	0.0522	0.0613	0.0367
COMPEN	12.7648	0.8718	12.8320	12.7543	0.8738	12.8171
DEBT	0.5783	0.4530	0.5555	0.6004	1.2869	0.5354
ROA	0.0194	0.1160	0.0275	0.0144	0.1145	0.0284
ASSET	14.3111 **	0.9127	14.3082	14.2326	0.9415	14.2062
DIRECT	9.6100	2.3500	9.0000	9.6600	2.1900	9.0000

[a] *$RISK_{it}$ is the i^{th} firm's default risk computed from the KMV model in year t; $DISACC_{it}$ represents the i^{th} firm's absolute discretionary accruals estimated from the modified Jones model in year t; $COMPEN_{it}$ refers to the cash compensation of the three highest paid employees in the i^{th} firm in year t; $DEBT_{it}$ is the i^{th} firm's debt ratio in year t; ROA_{it} indicates the i^{th} firm's return on assets in year t; $ASSET_{it}$ expresses the i^{th} firm's log total assets in year t; $DIRECT_{it}$ indicates to the size of the board of directors in the i^{th} firm in year t; $Y02_{it}$, $Y03_{it}$ and $Y04_{it}$, $Y05_{it}$, are (0,1) dummy variables controlling for the effects of calendar years; if the data is extracted from year 2002, then $Y02_{it}$ is 1, otherwise 0, and so on for the years 2003-2005.*

[b] *The two-tailed t-test was adopted to examine the means according to discretionary accruals (positive vs. negative).*

[c] ** indicates significance at the 10% level; ** indicates significance at the 5% level.*

Empirical Analysis

The empirical results for the total sample are presented in Table 2, which reports the coefficient estimates of the balanced panel multivariate regression model using the full 2,355 firm-year observations. After controlling for all other variables, the estimated coefficient of DISACC is positive and statistically significant, indicating that those firms which engage more vigorously in earnings manipulation have an increased risk of bankruptcy, thereby providing further support for Hypothesis 1. The result of *COMPEN* on default risk is negative and significant, providing support for our Hypothesis 2, and supporting the idea that companies which give higher CEO compensation suffer less default risk, in accordance with Cyert et al. (2002). The interaction term of *DISACC×COMPEN* is also negative and significant, providing support for Hypothesis 3, and indicating that higher CEO compensation could mitigate a positive relationship between discretionary accruals and firms' bankruptcy risk.

Table 2: Regression Results: Full Firms (N=2355)

Variables [a]	*Predicted Sign*	**Coefficient** [b] **(*t*-statistic)**
Constant		0. 0583 (1. 6009)
DISACC	+	0. 6158** (2. 3289)
COMPEN	-	-0. 0041* (-1. 7639)
*DISACC*COMPEN*	-	-0. 0418** (-2. 0068)
DEBT	+	0. 0269*** (20. 0487)
ROA	-	-0. 0322*** (-2. 9669)
ASSET	-	-0. 0019 (-0. 8571)
DIRECT	?	0. 0007 (1. 0203)
Y02	?	-0. 0010 (-0. 2875)
Y03	?	-0. 0026 (-0. 7344)
Y04	?	0. 0036 (0. 9845)
Y05	?	0. 0035 (0. 9484)

[a] *$RISK_{it}$ is the i^{th} firm's default risk computed from the KMV model in year t; $DISACC_{it}$ represents the i^{th} firm's absolute discretionary accruals estimated from the modified Jones model in year t; $COMPEN_{it}$ refers to the cash compensation of the three highest paid employees in the i^{th} firm in year t; $DEBT_{it}$ is the i^{th} firm's debt ratio in year t; ROA_{it} indicates the i^{th} firm's return on assets in year t; $ASSET_{it}$ expresses the i^{th} firm's log total assets in year t; $DIRECT_{it}$ indicates to the size of the board of directors in the i^{th} firm in year t; $Y02_{it}$, $Y03_{it}$ and $Y04_{it}$, $Y05_{it}$, are (0,1) dummy variables controlling for the effects of calendar years; if the data is extracted from year 2002, then $Y02_{it}$ is 1, otherwise 0, and so on for the years 2003-2005.*

[b] ** indicates significance at the 10% level; ** indicates significance at the 5% level; and *** indicates significance at the 1% level.*

The coefficient of *DEBT* is positive and significant, implying that those firms with poor solvency will have a higher probability of default; this is in line with the findings of Opler and Titman (1994). The effect of *ROA* on default risk is negative and significant, indicating that with an increase in the operational performance of firms, there will be a corresponding reduction in default risk. This is consistent with Vasiliou et al. (2003).

Where appropriate, the effects of calendar years are included as dummy intercepts within the panel regression. In general, it would seem that those firms with more aggressive earnings management behavior run a higher risk of default. The firms which gave CEO higher compensation could decrease the risk of default and provide the CEO with less incentive to manipulate earnings. The empirical results showing the relationship between default risk, discretionary accruals, and CEO compensation, for sub-groups based upon the sign of their discretionary accruals (+ or −) are provided in Table 3.

The estimated coefficient of *DISACC* is positive and significant for the positive discretionary accrual group, but not significant for the negative discretionary accruals group. This indicates that when firms use discretionary accruals to raise their apparent performance level, it will also raise the risk of default. This relationship does not hold for firms with negative discretionary accruals. It may be that managers using discretionary accruals to reduce a firm's performance are doing so in accordance with the accounting principal of 'conservatism', and may be regarded as conservative behavior.

Table 3: Panel Regression Estimation Results for the Total Sample, by Discretionary Accruals

Variables [a]	Predicted Sign	Positive Discretionary Accruals (n = 1103)		Negative Discretionary Accruals (n = 1252)	
		Coeff. [b]	t-statistic	Coeff. [b]	t-statistic
Constant		0.0334	0.6886	0.0596	1.2881
DISACC	+	0.7701 **	2.0545	0.3652	1.0123
COMPEN	-	-0.0092 ***	-2.7757	-0.0008	-0.2843
DISACC*COMPEN	-	-0.0520 *	-1.7657	-0.0278	-0.9780
DEBT	+	0.0711 ***	13.8853	0.0185 ***	15.7730
ROA	-	0.0566 ***	3.1640	-0.1051 ***	-8.2682
ASSET	-	0.0032	1.1921	-0.0048 *	-1.7355
DIRECT	?	0.0002	0.1912	0.0009	1.0117
Y02	?	-0.0059	-1.0235	0.0062	1.5662
Y03	?	-0.0095	-1.6179	0.0047	1.1567
Y04	?	-0.0060	-1.0294	0.0140 ***	3.1625
Y05	?	-0.0051	-0.8196	0.0073 *	1.6647

[a] *$RISK_{it}$ is the i^{th} firm's default risk computed from the KMV model in year t; $DISACC_{it}$ represents the i^{th} firm's absolute discretionary accruals estimated from the modified Jones model in year t; $COMPEN_{it}$ refers to the cash compensation of the three highest paid employees in the i^{th} firm in year t; $DEBT_{it}$ is the i^{th} firm's debt ratio in year t; ROA_{it} indicates the i^{th} firm's return on assets in year t; $ASSET_{it}$ expresses the i^{th} firm's log total assets in year t; $DIRECT_{it}$ indicates to the size of the board of directors in the i^{th} firm in year t; $Y02_{it}$, $Y03_{it}$ and $Y04_{it}$, $Y05_{it}$, are (0,1) dummy variables controlling for the effects of calendar years; if the data is extracted from year 2002, then $Y02_{it}$ is 1, otherwise 0, and so on for the years 2003-2005.*

[b] ** indicates significance at the 10% level; ** indicates significance at the 5% level; and *** indicates significance at the 1% level.*

The effects of *COMPEN* and *DISACC×COMPEN* are negative and significant for the positive discretionary accrual group, with the impact being insignificant for firms with negative discretionary accrual. We can infer from this that higher CEO compensation in firms with higher positive discretionary accrual could moderate top managers' manipulation of their earnings statements and lower the level of firms' default risk. Furthermore, for the two sub-groups, the coefficient of *DEBT* is significantly positive, whilst the impact of *ROA* is significantly negative. These results are similar to those presented in Tables 2. The marginal significant and negative influence of *ASSET* on the risk of bankruptcy for the negative discretionary accrual group suggests that when firm size is small, it has a higher probability of default as suggested by Vassalou and Xing (2004). Again, it seems appropriate to have time effects included as dummy intercepts in the regression. Overall, firms using positive discretionary accruals are generally faced with a higher risk of default. However, these conditions would not be apparent in the negative discretionary accruals group.

CONCLUSIONS

The prior literature suggests that if managers wish to deceive investors, they will invariably engage in the management of their earnings reports. In a country with asymmetric information and unsound legal and financial systems, which is the situation currently prevailing in mainland China; there is a particular risk of severe agency problems, whilst corporate financial frauds are also very common. Given that such scandals cause considerable harm to creditors and investors, the appropriate realization of the risk of all participants in Chinese security market is an issue of major concern.

In our study, we explore the relation of firms' earnings management behavior, top management compensation, and the probability of default. The results show firms' earnings management behavior could increase their probability of default. In addition, higher top management compensation not only mitigates firms' probability of default, but also lessens firms' use of discretionary accruals to manipulate financial statements resulting in higher default risk.

Furthermore, in order to investigate whether different earnings management behavior have different impacts on default risk, we separated our sample into two sub-groups based upon positive discretionary accruals (use of discretionary accrual to increase income) vs. negative discretionary accruals (use of discretionary accrual to decrease income). In line with many prior studies, the results reveal that only for firms with positive discretionary accruals, there is a relationship existing among earnings management behavior, top management compensation, and default risk.

Giving the relatively high financial risks involved with investing in China, creditors and investors (or would-be creditors and investors) need to understand firms' characteristics in advance, in order to protect their own interests. Once again, we suggest investing in firms with lower discretionary accruals and higher top management compensation; i.e., investors and creditors would be advised to put their money into firms with better corporate governance structures. If this is done, they can effectively avoid, or reduce, the risk of potential default.

APPENDICES

Appendix A: KMV Model

We use the KMV model – a model developed by the KMV Company in 1993 – to estimate and measure the default risk for the firms used in this study. The KMV model calculates the 'expected default frequency' (EDF) based on the firm's capital structure, the volatility of the asset returns, and the current asset value in accordance with the option pricing model of Black and Scholes (1973) and Merton (1974). This model is best applied to publicly-traded companies for which the value of equity is determined by the market.

There are three steps involved in deriving the actual probability of default. Firstly, we estimate the asset value and the volatility of the asset returns. Financial models usually consider the market value of assets, not the book value, since the latter represents only the historical cost of the physical assets, net of depreciation. Secondly, we calculate the default point. According to the KMV model, default occurs when the asset value reaches a level somewhere between the values of total liabilities and short-term debt. This point, which is referred to as the default point (*DPT*), is considered within the KMV model as the sum of the short-term debt plus half of the long-term debt. Thirdly, we calculate the 'distance to default' (*DD*), an index measure of default risk, which is the number of standard deviations between the mean of the distribution of the asset value and *DPT*. We then scale the *DD* to the actual probability of default using a default database. The estimation procedure is as follows.

$$\frac{dV_A^t}{V_A^t} = ud_t + \sigma_A dZ_t \tag{1A}$$

where V_A^t is the total market value of the assets for the firm at time t for China; u is the expected rate of return; and σ_A is the volatility of the asset returns. Thus, we can state the above equation in accordance with the option pricing model as follows:

$$V_E = V_A N(d_1) - Xe^{-rt} N(d_2) \tag{2A}$$

$$d_1 = \frac{ln(\frac{V_A}{X}) + (r_f + \frac{\sigma_A}{2})t}{\sigma_A \sqrt{t}}, \quad d_2 = d_1 - \sigma_A \sqrt{t} \tag{3A}$$

$$\sigma_E = \frac{V_A}{V_E} N(d_1) \sigma_A \tag{4A}$$

where V_A is the market value of assets for the firm listed in the China Stock Exchange; V_E is the equity market value for the Chinese listed company; σ_E represents the volatility of the equity returns; X is the book value of the total debt on the balance sheet; t represents the time to maturity of the debt; r_f is the one-year risk-free rate in the central bank of China; $N(d_1)$ expresses the hedging ratio with a cumulative probability density function; $N(d_2)$ is the probability that the market value of assets are greater than the liability at maturity t, a cumulative density probability function.

The implied market value and volatility of the asset, V_A and σ_A, can be calculated from Equations (2A) and (4A). We also need to compute the 'distance to default' (*DD*). Given that the total debt is regarded as the default point (*DPT*) for the firm, after being standardized by the standard deviation of asset returns, its *DD* can be expressed as:

$$DD = \frac{ln(V_A) - ln(u - \frac{\sigma_A^2}{2})t}{\sigma_A \sqrt{t}} \tag{5A}$$

The implied default risk for any period t – that is, the probability that the market values of the assets will be lower than those of the liabilities at maturity – is measured in accordance with the risk-neutral method. The procedure is as follows:

$$EDF_t = Pr\left[V_A^t \le X_t \middle| V_A^0 = V_A\right] = Pr\left[ln V_A^t \le ln X_t\right] \tag{6A}$$

After being represented in compliance with the Ito Process, the market values of the assets can be expressed, in logarithmic form, as follows:

$$ln V_A^t = ln V_A^0 + \left(u - \frac{\sigma_A^2}{2}\right)t + \sigma\sqrt{T}\varepsilon \tag{7A}$$

where ε denotes a random factor of asset returns.

We replace Equation (8A) into Equation (7A) after hypothesizing that the asset returns follow normal distribution. After arranging the related term, we obtain the default probability EDF_t, as follows:

$$\begin{aligned} EDF_t &= \Pr\left[V_A^t \le X_t \middle| V_A^0 = V_A\right] = \Pr\left[\ln V_A^0 + \left(u - \frac{\sigma_A^2}{2}\right)t + \sigma\sqrt{T}Z_t \le X_t\right] \\ &= \Pr\left[Z_t \le -\frac{\ln\left[\frac{V_A^0}{X_t}\right] + \left[r - \frac{\sigma_A^2}{2}\right]t}{\sigma\sqrt{t}}\right] = N(-d_2) \end{aligned} \tag{8}$$

Appendix B: Model to Predict Discretionary Accruals

In the study, we use the below model to predict discretionary accruals, which is stated as follows.

$$\frac{ACC_{it}}{TA_{it-1}} = \alpha_0\left(\frac{1}{TA_{it-1}}\right) + \alpha_0\left[\frac{(\Delta REV_{it} - \Delta REC_{it})}{TA_{it-1}}\right] + \alpha_2\left(\frac{PPE_{it}}{TA_{it-1}}\right) + \alpha_3 ROA_{it} + \alpha_4 Y02_{it} + \alpha_5 Y03_{it} + \alpha_6 Y04_{it} + \alpha_7 Y05_{it} + \varepsilon_{it} \tag{1B}$$

where ACC_{it} is the i^{th} firm's accruals in year t, which is defined as the earnings minus the cash flow from operations, both are drawn from the cash flow statement; REV_{it} denotes the i^{th} firm's net revenue in year t; $\triangle REV_{it}$ represents the i^{th} firm's change in revenue in year t; that is, it is equal to $REV_{it} - REV_{it-1}$; REC_{it} refers to the i^{th} firm's net revenue in year t; $\triangle REC_{it}$ stands for the change in sales in year t, namely, it is identical to $REC_{it} - REC_{it-1}$; PPE_{it} indicates the i^{th} firm's gross property, plant, and equipment in year t; ROA_{it} expresses the i^{th} firm's return on assets in year t, which is defined by the net income scaled by lagged total assets; TA_{it} is the i^{th} firm's total assets in year t; $Y02_{it}$, $Y03_{it}$, $Y04_{it}$ and $Y05_{it}$ are the i^{th} firm's (0,1) dummy variables controlling for the effect of calendar years; ε_{it} is the i^{th} firm's error term of model (1B) in year t.

Although previous papers have used the time-series approach for each firm (Jones, 1991) or the cross-section (DeFond and Subramanyam, 1998; Cohen and Lays, 2006) models to estimate the accruals, both approaches have their limitations. The time series approach assumes the state of the parameter follows temporal stationary whereas the cross-sectional method supposes the homogeneity is being across firms. Our study uses random effects panel regression model to estimate the accruals model, which contain both cross-sectional and time series dimensions.

The estimated parameters obtained from Equation (1B) are used to evaluate expected nondiscretionary accruals ($NDACC_{it}$). The related equation is set as follows.

$$NDACC_{it} = \hat{\alpha}_0\left(\frac{1}{TA_{it-1}}\right) + \hat{\alpha}_1\left[\frac{(\Delta REV_{it} - \Delta REC_{it})}{TA_{it-1}}\right] + \hat{\alpha}_2\left(\frac{PPE_{it}}{TA_{it-1}}\right) + \hat{\alpha}_3 ROA_{it} + \hat{\alpha}_4 Y02_{it} + \hat{\alpha}_5 Y03_{it} + \hat{\alpha}_6 Y04_{it} + \hat{\alpha}_7 Y05_{it} \tag{2B}$$

where the definition of the independent variables are identical to those in equation (2B). Then the discretionary accruals ($DACC_{it}$) equals to the actual accruals minus the predicted accruals ((ACC_{it} / TA_{it-1}) - $DACC_{it}$).

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DO DIVIDEND CLIENTELES EXPLAIN PRICE REACTIONS TO DIVIDEND CHANGES?

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ABSTRACT

Previous studies find that stock price reactions to dividend announcements are positively related to dividend yield, consistent with the dividend-clientele hypothesis. In this paper, we argue that this yield-related clientele effect can be attributed to estimation biases in using preannouncement dividends as a proxy for market's anticipated dividends. Based on our samples constructed to mitigate the dividend estimation biases, we find that dividend yield has no additional power beyond the standardized dividend change in explaining the announcement-period excess returns. Our results are consistent with the information/signaling hypothesis, but inconsistent with the dividend-clientele hypothesis. In addition, we find that firm size remains negatively related to the price reactions to dividend changes.

JEL: G14, G35

INTRODUCTION

This study reexamines the dividend-clientele hypothesis presented in previous studies. Bajaj and Vijh (1990) suggest that the existence of dividend clienteles may partially explain price reactions to dividend change announcements. They argue that if marginal investors in different stocks value dividends differently, anticipated dividend yield should be associated with the price reactions to dividend change announcements.

We argue that this hypothesis has no solid theoretic base and is an empirical issue. If there is sufficient adjustment of investors in response to dividend yield changes, with some leaving a clientele replaced by others entering it, dividend yield should play little role in explaining the stock price reaction to dividend changes. As shown in this paper, their founding of a positive relationship between stock price reactions to dividend changes and dividend yield may be attributed to improper estimation of anticipated dividends in previous studies that use preannouncement dividends as a proxy for market's anticipated dividends. If markets partially anticipate a dividend increase, for example, the abnormal stock price reaction will be relatively small and reflect only the unexpected portion in the dividend change. Using preannouncement dividends as the proxy for the market's expected dividends may cause two kinds of estimation biases. First, it underestimates anticipated dividend yield by markets and therefore associates low dividend yields with small excess returns. Second, it exaggerates the unexpected dividend increases (or dividend surprises) and thus underestimates the information effect of dividend changes, making the above spurious dividend yield effect more conspicuous. The dividend expectation is less a problem in the case of dividend decreases. Since firms seldom cut dividends, dividend decreases should contain a greater unexpected component.

In this paper, we construct two subsamples of dividend increase announcements which largely mitigate the dividend estimation problem. The first subsample is obtained by excluding those announcements with no or negative stock price reactions. We believe these dividend increases are largely anticipated by the market. If the realized dividend increase is less than what the market expects, the dividend increase announcement in fact represents a negative dividend surprise and causes stock price to decrease. Our second subsample consists of only the announcements with unusually large dividend increases, say, at least 50 percent. We conjecture that these large dividend changes contain greater unexpected components than small dividend increases. This idea is similar to the view of Asquith and Mullins (1983) that unusual dividend policy changes such as dividend initiations are more likely to be unexpected.

Our sample closely resembles those in the previous studies. There are 7715 dividend increase and 849 dividend decrease announcements over the period 1970-2001. We find that announcement period excess returns are positively related to the magnitude of standardized dividend changes and to dividend yield for the dividend increase sample. But the excess returns are unrelated to dividend yield in the dividend decrease sample. This evidence supports our dividend expectation argument and is inconsistent with the dividend-clientele hypothesis.

Our main evidence against the dividend-clientele hypothesis comes from the two subsamples. Within each subsample, when there is no control for dividend change, the higher the yield, the greater the announcement period excess return. However, no such pattern is observed after controlling for dividend change. The results from cross-sectional regressions provide further support for our argument. Although the coefficient of dividend yield is highly significant and positive in univariate regressions, it becomes insignificant and positive in the multivariate regressions with the inclusion of standardized dividend change. In addition, consistent with prior studies, firm size is found to be negatively related to excess returns. Stock price, however, is no longer associated with excess returns after controlling for the dividend change.

The rest of the paper is organized as follows. In the following section, we discuss the relevant literature. Next, we describe the data and methodology and provide summary statistics of the samples. The empirical results are reported and discussed in the following section. In addition, findings of previous studies are also replicated for comparison. The paper closes with some concluding comments.

LITERATURE REVIEW

In their seminal work, Miller and Modigliani (1961) demonstrate that, absent imperfections, a firm's dividend policy does not affect its value. Since then, challenges to this dividend irrelevance proposition have focused on imperfections. In particular, subsequent research has extensively explored the effects of tax-induced clienteles on capital asset prices and the stock price reactions to dividend announcements.

Before the implementation of the 1986 Tax Reform Act, the dividend income was taxed at a higher rate than capital gains, and this suggests a negative price impact of dividends. This idea is supported by the CAPM-based studies including Litzenberger and Ramaswamy (1979, 1980), Rosenberg and Marathe (1979), and Blume (1980). Using capital asset pricing models incorporating taxes, these studies find that, if risk is held constant, before-tax returns are an increasing function of dividend yield.

Black and Scholes (1974) and Miller and Scholes (1978) demonstrate that dividend irrelevance may hold even if there is differential taxation of dividends and capital gains. Miller and Scholes (1978) argue that the dividend receipts can be made tax exempt by laundering them with personal borrowing. Black and Scholes (1974) extend the concept of investor clienteles proposed by Miller and Modigliani (1961), i.e., low (high) yielding stocks being held by investors in high (low) marginal tax brackets if tax rates vary across investors. They emphasize the ability of firms to adjust dividends to appeal to tax-induced investor clienteles and argue that this supply effect may account for their finding of no significant relationship between dividend yields and stock returns.

An alternative viewpoint, set forth by Litzenberger and Ramaswamy (1980) suggests that firms may make incomplete supply adjustments and individuals' portfolios may be limited by shortsale and margin restrictions. In equilibrium, therefore, the relative prices of dividends and capital gains will reflect the tax situation of the marginal investor in the stock. Miller and Scholes (1982) criticize that the dividend yield effect found in Litzenberger and Ramaswamy (1979, 1980) may be attributed to the information biases. In response, Litzenberger and Ramaswamy (1982) use an "information free" sample and still find the yield coefficient to be positive and significant in their after-tax CAPM model.

As debate goes on, Kalay and Michaely (1983) argue that while the after-tax CAPM predicts cross-sectional return variation as a function of dividend yield, the Litzenberger and Ramaswamy test is inadvertently designed to discover whether the ex-dividend period offers unusually large risk adjusted returns. Separating the time series from the cross-sectional return variation, Kalay and Michaely cannot detect any return variation across stocks with different yields. Chen, Grundy, and Stambaugh (1990) show that the positive association between yields and returns can be explained by a time-varying risk premium correlated with yield. When they allow the risk measures to vary, the yield coefficient was found positive but insignificant.

While the evidence about the tax-induced dividend yield effect is far from conclusive, there seems to be an overall agreement that the market perceives that dividend changes convey new information about the value of firm. In their original article, Miller and Modigliani (1961) suggest that dividends may provide a vehicle for communicating management's superior information concerning their assessment of the firm's prospect. This view of "information content of dividends" is supported by the empirical evidence in numerous studies examining the price reactions to dividend changes. For example, Pettit (1972) shows that announcements of dividend increases are followed by a significant price increase and announcements of dividend decreases are followed by a significant price drop. Aharony and Swary (1980) find that these relationships hold even after controlling for contemporaneous earnings announcements. Focusing on extreme changes in dividend policy, Asquith and Mullins (1983), Healy and Palepu (1988), and Michaely, Thaler, and Womack (1995)) show that the market reacts quite severely to dividend initiations or omissions announcements.

Recently, Nissim and Ziv (2001) find that dividend changes provide information about the level of profitability in subsequent years, incremental to market and accounting data. They also find that dividend changes are positively related to earnings changes in each of the two years after the dividend change. Koch and Sun (2004) present results suggesting that changes in dividends cause investors to revise their expectations about the persistence of past earnings changes. Docking and Koch (2005) document that dividend change announcements elicit a greater change in stock price when the nature of the news (good or bad) goes against the grain of the recent market direction during volatile times

Extending the previous studies, as noted earlier, Bajaj and Vijh (1990) suggest that the existence of dividend clienteles may partially explain price reactions to dividend change announcements. They argue that if marginal investors in different stocks value dividends differently, anticipated dividend yield should be associated with the price reactions to dividend change announcements. For an investor with a relatively high aversion to dividends, for example, the positive information in a dividend increase is accompanied by the negative effect of higher-than-anticipated yield. In contrast, the two effects act in the same direction for an investor with a preference for dividends. If investors with preference for dividends are marginal investors in high-yield stocks, the price reaction to dividend change should be larger, the higher the anticipated yield of the stock. This dividend-clientele hypothesis is supported by the evidence in Bajaj and Vijh (1990) and Denis, Denis, and Sarin (1994). Both studies use preannouncement dividend yield as a proxy for anticipated yield and find that the magnitude of stock price reaction to a dividend change announcement is positively related to dividend yield. In addition, Bajaj and Vijh (1990) find that stock price changes are negatively related to firm size and stock price.

Several approaches have been proposed in the literature to capture the unexpected component of dividend changes, including Lintner (1956) model, the Box-Jenkins model, and the Value Line dividend forecasts. However, according to a study by Bar-Yosef and Sarig (1992), the measures of dividend surprises based on these methods are not significantly correlated with the market reactions to dividend change announcements in a sample of large firms traded in the NYSE.

DATA AND METHODOLOGY

In our sample, information on dividend declarations is obtained from the CRSP NYSE/AMEX Monthly Master Files and daily rates of returns are from the CRSP Daily Master File. The time period covers from July 1970 to December 2001, excluding the latter half of 1987 to avoid the period affected by the market crash in the October of that year. In addition, we use the following criteria to select firms:

(1) Absolute changes in consecutive regular quarterly dividends per share are greater than 10%. We require that no other type of distribution is made over the period between the two quarterly dividend announcements. Thus, firms that pay stock dividends or special dividends and firms that split their shares during the quarter in question are not included in the sample.

(2) The dividend initiations and omissions are excluded from the sample. This is due to the difficulty involved in calculating the anticipated yield and much larger price responses to these events than those to regular dividend changes (See Asquith and Mullins (1983), Healy and Palepu (1988), and Michael, Thaler, and Womack (1995)).

(3) We choose only dividend announcements for which the announcement date precedes the ex-dividend day by at least eight trading days. The eight-day window is chosen because the ex-day effect is observed up to five days before the ex-day, as documented by Eades, Hess, and Kim (1984)).

Out sample consists of 8664 dividend announcements that satisfy the above criteria. Of these, 7715 are dividend increases and 849 are dividend decreases. Most of the reduction in the sample (45.2% of the cases) occurs as a result of criterion that the announcement day and the ex-dividend day must be separated by at least 8 trading days. The sample is representative of the CRSP population. In particular, there is no significant difference between the in-sample and out-of-sample firms in the concentration of announcements in any calendar year, calendar month, or particular industry group.

We categorize stocks into low-, medium-, and high-yield groups of equal size. Dividend yield, used as proxy for anticipated yield in previous studies, is measured as the most recent ordinary cash dividend preceding the sample announcement divided by the stock price of the firm as of two days prior to the sample announcement. Bajaj and Vijh (1990) measure anticipated yield by using the dividends over the prior 12-month period and the stock price at the beginning of the period. They also rank yields by calendar quarters, rather than over the aggregate data for the entire-sample period. We checked that our results are not sensitive to their yield measure and ranking procedure.

Table 1 shows the distribution of annualized dividend yield for each of the dividend yield groups in dividend increase and decrease samples. Overall, the mean annual yield is 3.09% for the dividend increase sample and 8.30% for the dividend decrease sample, with corresponding standard deviations of 1.71% and 5.68%, respectively. For the dividend increase sample, the low-yield group has a sample mean of 1.42% with a standard deviation of 0.51%. The medium- and high-yield groups have sample means of 2.85% and 4.99%, with corresponding standard deviations of 0.39% and 1.38%, respectively. These statistics are also reported for the dividend decrease sample.

The standard event study methodology is employed to examine the stock price reactions to dividend change announcements. We define the event period affected by the dividend announcement as the day before to the day after the CRSP announcement date. Daily abnormal returns are measured as unadjusted returns subtracting the returns on the CRSP value-weighted index. We also used the equally-weighted index returns and the market model to measure abnormal returns. Since similar results are obtained, we report only those from using value-weighted index adjusted returns.

Table 1: Summary Statistics of Dividend Yields

		Dividend Increases		Dividend Decreases	
Yield Category	**Variables**	**Mean**	**Std. dev.**	**Mean**	**Std. dev.**
Low	YLD	1.42	0.51	3.97	.27
	Obs.	2573		282	
Medium	YLD	2.85	0.39	7.15	0.86
	Obs.	2569		284	
High	YLD	4.99	1.38	13.70	6.69
	Obs.	2573		283	
Total	YLD	3.09	1.71	8.30	5.68
	Obs.	7715		849	

The sample of 8664 dividend announcements consists of all NYSE/AMEX stocks from CRSP tape that satisfy the following criteria: (1) Absolute changes in consecutive regular quarterly dividends are greater than 10%; (2) The announcement does not represent dividend initiation or omission; and (3) The announcement date precedes the ex-dividend day by at least 8 trading days. The sample is partitioned on the basis of preannouncement dividend yield (YLD) into low-yield, medium-yield and high-yield groups for both dividend increase and dividend decrease samples. YLD is measured as the firm's most recent preannouncement dividends divided by the firm's stock price 2 days prior to the announcement.

RESULTS

We begin by replicating the basic results of previous empirical studies using our sample. Panel A of Table 2 presents excess returns during the 3-day announcement period for dividend increase sample. Overall, announcements of dividend increases are associated with an average excess return of 1.46%, significant at 0.01 level. In addition, we divide the sample on the basis of the absolute value of the standardized dividend change. Consistent with the information/signaling hypothesis, each row indicates that the magnitude of the stock price reaction increases with size of dividend changes.

Each column in Panel A of Table 2 shows that the magnitude of the average abnormal stock price response to dividend increase announcement increases with dividend yield, with or without controlling for the dividend change. Panel B of Table 2 reports the results of cross sectional regression tests of excess returns surrounding the dividend change announcement. The coefficient of the yield is very significant and positive despite the inclusion of the proxies for dividend change. These results replicate the findings of Bajaj and Vijh (1990) and are consistent with the dividend-clientele hypothesis.

The results obtained from the dividend decrease sample, as shown in Table 3, are also consistent with the information/signaling hypothesis, however, they do not support the clientele hypothesis. The magnitude of stock price reaction increases with the absolute value of the size of the dividend change. But there is no such monotonic relationship between the stock price reactions and dividend yield, with or without controlling for dividend change.

In the cross-sectional regressions, presented in Panel B of Table 3, although the coefficient of dividend yield is significant and negative in the univariate regression, it becomes positive and insignificant after the inclusion of the absolute value of dividend change. The similar result from cross-sectional regressions is found in Denis, Denis, and Sarin (1994). However, this evidence is ignored in their study.

Table 2: Excess Returns and Cross-Sectional Regressions for Dividend Increases

Panel A: Announcement Period Excess Returns (Numbers in parentheses are standard errors of the mean)

Yield Category	Dividend Change Category			
	Low	**Medium**	**High**	**All**
Low	0.46	0.96	1.70	0.71
	(0.09)	(0.19)	(0.33)	(0.08)
Medium	0.78	1.01	2.28	1.30
	(0.12)	(0.11)	(0.17)	(0.08)
High	0.76	1.29	3.14	2.37
	(0.41)	(0.11)	(0.16)	(0.08)
All	0.57	1.11	2.72	1.46
	(0.07)	(0.07)	(0.09)	(0.05)

Panel B: Estimated Coefficients of Cross-Sectional Regressions (numbers in parentheses are t-statistics)

	INTCP	**CHGN**	**YLD**	**MVAL**	**PRC**	**R^2-adjusted**
1	0.43	160				0.0584
	(6.5)***	(21.9)***				
2	0.03		46.3			0.0363
	(0.33)		(17.1)***			
3	5.77			-0.36		0.0234
	(18.01)***			(-13.6) ***		
4	1.91				-0.01	0.0067
	(24.6)***				(-7.30) ***	
5	-0.19	129	26.4			0.0680
	(-2.02)**	(16.2)***	(8.98)***			
6	3.01	120	26.1	-0.28	-0.01	0.0782
	(8.38)***	(14.8)***	(8.79) ***	(-9.29) ***	(-3.87) ***	

*The sample of 7715 dividend increase announcements consist of all NYSE/AMEX stocks from CRSP tape that satisfy the following criteria: (1) Changes in consecutive regular quarterly dividends are greater than 10%; (2) The announcement does not represent dividend initiation; and (3) The announcement date proceeds the ex-dividend day by at least 8 trading days. The sample is partitioned on the basis of the preannouncement dividend yield (YLD) and the value of the standardized dividend change (CHNG). Three-day announcement period excess returns (CARs) are calculated as unadjusted returns minus the returns on the CRSP value-weighted index. The dependent variable in the regressions is CAR. MVAL is the natural log of the market value of the firm's equity at the end of year prior to the announcement; PRC is the stock price as of two days prior to the announcement. ***, **, and * indicate significance at 1, 5, 10 percent levels respectively.*

Table 3: Excess Returns and Cross-Sectional Regressions for Dividend Decreases

Panel A: Announcement Period Excess Returns (numbers in parentheses are standard deviations of the mean)

Yield Category	Dividend Change Category			
	Low	**Medium**	**High**	**All**
Low	-1.89	-5.37	-8.97	-3.14
	(0.43)	(0.77)	(3.12)	(0.40)
Medium	-3.58	-6.38	-7.18	-5.89
	(0.69)	(0.64)	(0.97)	(0.45)
High	-1.89	-2.94	-5.43	-4.61
	(0.90)	(0.97)	(0.65)	(0.51)
All	-2.32	-5.34	-5.99	-4.55
	(0.35)	(0.44)	(0.53)	(0.26)

Panel B: Estimated Coefficients of Cross-Sectional Regressions (numbers in parentheses are t-statistics)

	INTCP	CHGN	YLD	MVAL	PRC	R^2-adjusted
1	-3.28	-33.7				0.0268
	(-8.88) ***	(-4.92) ***				
2	-3.12		-17.5			0.0157
	(-6.72) ***		(-3.80) ***			
3	-5.59			0.09		-0.0008
	(-3.11) ***			(0.57)		
4	-4.52				-0.00	-0.0012
	(-9.93) ***				(-0.15)	
5	-3.53	-42.9	7.14			0.0263
	(-7.37) ***	(-3.20) ***	(0.80)			
6	-5.30	-43.4	4.70	0.24	-0.04	0.0282
	(-2.91) ***	(-3.25) ***	(0.52)	(1.43) *	(-1.72) **	

*The sample of 849 dividend decrease announcements consist of all NYSE/AMEX stocks from CRSP tape that satisfy the following criteria: (1) Absolute changes in consecutive regular quarterly dividends are greater than 10%; (2) The announcement does not represent dividend omission; and (3) The announcement date proceeds the ex-dividend day by at least 8 trading days. The sample is partitioned on the basis of the preannouncement dividend yield (YLD) and the absolute value of the standardized dividend change (CHNG). Three-day announcement period excess return (CAR) is calculated as unadjusted returns minus the returns on the CRSP value-weighted index. The dependent variable in the regressions is CAR. MVAL is the natural log of the market value of the firm's equity at the end of year prior to the announcement; PRC is the stock price as of two days before the announcement. ***, **, and * indicate significance at 1, 5, 10 percent levels respectively.*

Main Empirical Results

As we argued in Section I, using preannouncement dividends as a proxy for market's anticipated dividends may cause a spurious dividend yield effect in stock price reactions to dividend increase announcements. Since firms seldom reduce the level of dividends, dividend decreases more likely represent dividend surprises than dividend increases. Therefore, the dividend estimation is less a problem in the case of dividend decreases. In the following, we utilize the two subsamples of dividend increases which we believe largely reduce the dividend estimation problem.

We construct our first subsample by selecting stocks which have positive price reactions to dividend increase announcements, i.e., by choosing the firms with CAR>0 in the announcement period. Those firms with CAR≤0 are excluded from the sample because their dividend increase announcements may contain relatively small or even negative dividend surprises. If realized dividend increases are below the market expectation, the dividend increase announcements in effect are negative dividend surprises, and therefore may cause stock prices to decline. The firms selected with CAR>0 should contain greater unexpected component of dividend increases in the announcement.

The second subsample is selected in a relatively straightforward way. Specifically, we construct our second subsample by choosing the firms with dividend increase of at least 50%. We conjecture that unusually large dividend increases contain greater unexpected component than small dividend increases, and that smaller dividend increases are more likely anticipated by the market. This idea is consistent with the evidence that price change is positively associated with the magnitude of dividend change.

Our primary concern in building these two subsamples is that they may be concentrated in a certain yield range, thus possibly causing selection biases. Table 4 summarizes the statistics of the two subsamples, along with the total sample of dividend increase for comparison. The two subsamples consist of 4934 and 643 observations respectively.

The first subsample CAR>0 eliminates more firms in the low dividend yield range than in high dividend yield range. This is not surprising because there is positive correlation between standardized dividend change and dividend yield. By imposing the restriction of CAR>0, we are eliminating more firms that have small standardized dividend changes, and hence also firms that have small dividend yields. Table 4 reports summary statistics for all three yield categories. For each of the three categories, the standardized dividend change and dividend yield are larger in the subsample CAR>0 than in the total sample. But the differences are small. The average dividend yields are 3.09% for the total dividend increase sample and 3.29% for the subsample CAR>0. The second subsample (ΔD/D≥50%) has relatively large standardized dividend changes by construction. Since firms in this subsample on average has lower dividend level before the dividend increase, the average yield for each yield category is smaller in this subsample than in the total sample.

Table 4: Summary Statistics on the Subsamples of Dividend Increases

Yield Category		Dividend Increases	CAR>0	ΔD/D≥50%
Low	CHNG	0.40	0.43	0.74
	YLD	1.42	1.52	0.80
	PRC	38	38	37
	MVAL	12.4	12.4	12.0
	N	2573	1644	213
Medium	CHNG	0.63	0.69	1.33
	YLD	2.85	3.06	1.85
	PRC	31	32	22
	MVAL	12.1	12.0	11.2
	N	2569	1645	216
High	CHNG	0.91	0.99	2.64
	YLD	4.99	5.27	3.79
	PRC	25	24	17
	MVAL	11.7	11.6	10.7
	N	2573	1645	214
All	CHNG	0.65	0.70	1.57
	YLD	3.09	3.29	2.15
	PRC	31	31	25
	MVAL	12.1	12.0	11.3
	N	7715	4934	643

Of the sample of 7715 dividend increase announcements, two subsamples are created. The first subsample CAR>0 consists of only the firms with positive announcement period abnormal returns. The subsample ΔD/D≥50% consists of only the firms with at least 50% increase in dividends. The two samples are partitioned on the basis of the preannouncement dividend yield (YLD). CHNG is the standardized dividend change, calculated as the dividend increase divided by the stock price two days prior to the announcement. MVAL is the natural log of the market value of the firm's equity at the end of year prior to the announcement; PRC is the stock price as of two days before the announcement; N is the number of observations in each subcategory.

Panel A of Table 5 presents excess returns for the first subsample consisting of only firms that have positive excess returns. Within each yield category, the magnitude of average excess return increases monotonically with standardized dividend changes. This is consistent with the information/signaling hypothesis. However, controlling standardized dividend change, the announcement period return does not show an increasing pattern as dividend yield increases.

Panel B reports same statistics for the second subsample, composed of firms that increase their dividends by at least 50%. Controlling standardized dividend changes, the excess returns do not show an increasing pattern from low-dividend yield category to high-dividend category. Instead, the figures indicate a U-shaped pattern between excess return and dividend yield. The evidence in Table 5 is consistent with the signaling hypothesis but does not support the dividend-clientele hypothesis.

Table 5: Announcement Period Excess Returns for the Subsamples of Dividend Increase

Panel A: Subsample CAR>0

Yield Category	Dividend Change Category			
	Low	**Medium**	**High**	**All**
Low	2.92	3.58	4.79	3.27
	(0.08)	(0.19)	(0.34)	(0.08)
Medium	2.73	3.02	4.46	3.36
	(0.12)	(0.10)	(0.19)	(0.08)
High	2.51	3.16	4.73	4.07
	(0.28)	(0.11)	(0.13)	(0.09)
All	2.85	3.19	4.66	3.57
	(0.07)	(0.07)	(0.10)	(0.05)

Panel B: Subsample ΔD/D≥50%

Yield Category	Dividend Change Category			
	Low	**Medium**	**High**	**All**
Low	0.90	2.60	7.44	1.48
	(0.36)	(0.76)	(3.48)	(0.36)
Medium	0.73	1.70	3.32	1.81
	(0.75)	(0.54)	(0.74)	(0.39)
High	N/A	3.16	5.27	4.77
		(0.86)	(0.49)	(0.43)
All	0.86	2.19	4.99	2.69
	(0.33)	(0.41)	(0.43)	(0.24)

Of the sample of 7715 dividend increase announcements, two subsamples are created. The first subsample CAR>0 consists of only the firms with positive announcement period abnormal returns. The subsample ΔD/D≥50% consists of only the firms with at least 50% increase in dividends. The two samples are partitioned on the basis of the preannouncement dividend yield (YLD). Three-day announcement period excess return (CAR) is calculated as unadjusted returns minus the returns on the CRSP value-weighted index. CHNG is the standardized dividend change, calculated as the dividend increase divided by the stock price two days prior to the announcement. MVAL is the natural log of the market value of the firm's equity at the end of year prior to the announcement; PRC is the stock price as of two days before the announcement; N is the number of observations in each subcategory. Numbers in Parentheses are standard errors of the mean.

The regression results in Table 6 reinforce the evidence in Table 5. Panel A of Table 6 summarizes the regression results for the subsample CAR>0. The dividend yield is significant only in the univariate model not controlling the information content of the announcement. When the standardized dividend change is included in the regression, the "yield effect" disappears. Panel B shows similar results for the second subsample.

Table 6: Estimated Coefficients of Cross-Sectional Regressions

Panel A: Subsample CAR>0

	INTCP	CHNG	YLD	MVAL	PRC	R^2-adjusted
1	2.64	131				0.0710
	(39.6) ***	(19.4) ***				
2	2.77		23.99			0.0159
	(27.6) ***		(8.97) ***			
3	9.29			-0.48		0.0637
	(29.4) ***			(-18.3) ***		
4	4.29				-0.02	0.0289
	(55.9) ***				(-12.2) ***	
5	2.55	127	3.57			0.0711
	(25.9) ***	(17.1) ***	(1.25)			
6	7.28	107	0.00	-0.37	-0.00	0.1088
	(20.3) ***	(14.5) ***	(0.00)	(-12.1) ***	(-0.78)	

Panel B: Subsample ΔD/D≥50%

	INTCP	CHNG	YLD	MVAL	PRC	R^2-adjusted
1	1.21	94.5				0.0597
	(3.72) ***	(6.46) ***				
2	0.98		79.8			0.0438
	(2.53) ***		(5.51) ***			
3	10.30			-0.67		0.0362
	(6.70) ***			(-5.01) ***		
4	3.19				-0.02	0.0063
	(9.87) ***				(-2.25) **	
5	0.91	73.9	328.7			0.0612
	(2.38) ***	(3.59) ***	(1.42) *			
6	8.35	83.5	13.0	-0.67	-0.01	0.0882
	(5.00) ***	(4.09) ***	(0.63)	(-4.53) ***	(-1.29) *	

*Of the sample of 7715 dividend increase announcements, two subsamples are created. The first subsample CAR>0 consists of only the firms with positive announcement period abnormal returns. The subsample ΔD/D≥50% consists of only the firms with at least 50% increase in dividends. The two samples are partitioned on the basis of the preannouncement dividend yield (YLD). Three-day announcement period excess return (CAR) is calculated as unadjusted returns minus the returns on the CRSP value-weighted index. CHNG is the standardized dividend change, calculated as the dividend increase divided by the stock price two days prior to the announcement. MVAL is the natural log of the market value of the firm's equity at the end of year prior to the announcement; PRC is the stock price as of two days before the announcement; N is the number of observations in each subcategory. Numbers in parentheses are t-statistics. ***, **, and * indicate significance at 1, 5, 10 percent levels respectively.*

CONCLUSIONS

Bajaj and Vijh (1990) propose that investors' preference for dividends should be reflected in the stock price reaction to dividend change announcements. Their study and Denis, Denis, and Sarin (1994) find that the magnitude of price change in response to dividend announcements is positively related to dividend yield, supporting this dividend-clientele hypothesis. In this study, we argue that the dividend yield effect found in the previous studies may result from the estimation biases in using preannouncement dividends as a proxy for market's anticipated dividends. Based on our samples which we believe effectively mitigate the estimation biases, we find that dividend yield effect is insignificant and dominated by the information effect. The evidence presented in this study raises serious doubts about the existence of yield-related dividend clientele effect in the price reactions to dividend change announcements. We are inclined to believe that the yield-related clientele effect does not show in the stock price reactions to dividend change announcements. However, the paper does not address the estimation biases in dividend yield-related effect in capital assets pricing. Further research should explore this important issue.

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ON THE OPTIMAL PACKAGE FORMAT FOR ASSET SELLERS

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ABSTRACT

A seller who owns two common-value assets can choose to either sell them as a bundle or separately. In this paper, we present a theoretical model to select the optimal selling option when there is asymmetric information between the seller and the buyers. Our main finding is that separate selling makes the seller fall into a bilateral monopoly environment, in which the assets are sold through bargaining, while bundled selling leads to a competitive bidding environment. When the seller's bargaining ability is given, the difference between the two assets' values increases, so the seller's incentive to sell as a bundle decreases. On the other hand, given the values of both assets, when the seller's bargaining power increases, the incentive to sell as a bundle decreases.

JEL : D44; D82

INTRODUCTION

Consider a large corporation with several divisions or branches being sold. The seller can either sell those branches as a bundle or take them apart and sell them separately. The seller may have more information about the assets being sold than potential buyers. Some of the potential buyers may have specialized knowledge of particular branches being sold, but may be unfamiliar with other branches. This situation is not only common in the sale of corporations but is frequently seen in the sale of financial assets. Of interest is to determine the optimal decision for the seller—to sell these assets as a bundle, or to sell them separately.

Earlier studies, such as Stigler (1968), Adams and Yellen (1976), and McAfee et al. (1989) among others, have already dealt with bundling problems. However, most of these studies focused on cases in which the value of goods varies with buyers' preferences or purposes. This is not consistent with the common-value characteristic of assets. To fill this gap, this paper examines the bundling problem in regard to the sale of assets.

We provide a model to consider the optimal package format for the seller of assets. The common-value characteristic of assets and asymmetric information between the seller and the buyers are considered in this model. The seller owns two different but symmetric assets that are for sale, and has complete information about these assets. On the other hand, each buyer has information regarding only one of the assets, and each buyer's information concerns a different asset. When the assets are sold separately, an individual buyer will not participate in the transaction for the object where he has inferior information. Thus, separate selling makes the seller fall into a bilateral monopoly environment, in which the object is sold through bargaining.

When the assets are sold in a bundle, where every buyer possesses some information about the object, and bidding in accordance with this information enables them to generate positive revenue, bundled selling leads to a competitive bidding environment. We find that given the seller's bargaining ability, the smaller that the difference between the two assets' values is, the stronger the tendency is for the seller to sell the assets as a bundle. On the other hand, given both assets' values, the seller's increasing bargaining power weakens his/her incentive to sell the assets in a bundle. The rest of the paper is organized as follows. In section 2, we discuss the relevant literature. Section 3 provides the introduction to the model. Section 4 explores how the seller decides the optimal selling format. Section 5 is the conclusion.

RELATED LITERATURE

Some studies that focus on asset-backed securities provide various explanations for the reasons why people often adopt a bundling strategy for selling mortgages. Boot and Thakor (1993) concluded that bundling improves the precision of the information of informed traders, increases their return on information, and encourages higher informed demand. This makes the seller of a high-quality firm better off. Others, such as Riddiough (1997) and Glaeser and Kallal (1997), suggested that packaging mortgages is a strategy that can soften the problem of a lemon-related subordinated effect, and therefore increase liquidation proceeds. Since bundling effectively weakens the problem that results from information asymmetry, most of these studies regard bundling as a strategy that dominates unbundled sales. However, as we can observe, unbundled sales are not uncommon in practice. In our model, we explore an environment where both formats are possible. In this environment, the relative value of assets determines the seller's choice on package format.

Studies focusing on auctions of multiple products note that a bundling strategy should be adopted conditionally. Palfrey (1983) suggested that a bundling strategy is optimal when there are few bidders, and that a seller tends to prefer a separate auction when there are more bidders. Avery and Hendershott (2000) noted that the optimal auction format favors allocations that bundle the product to the bidders who are interested in both products, and the form of bundling can be different in an auction. Probabilistic bundling – where the bidder's chance of winning the product increases with her value of the other product – is usually adopted in an auction. Armstrong (2000) suggested that the optimal auction takes one of two formats: the objects are sold at independent auctions or with a degree of bundling, and are in a sense similar to probability bundling. The optimal format of an auction depends on the number of bidders, as reflected by Palfrey's results (1983). In our paper, we only consider a two assets case, so the problem of choosing package format is simpler compared to these papers. However, we emphasize the impact of information asymmetry between the seller and buyers. This situation is not considered by earlier authors, but is quite general in the case of financial or several real assets.

THE MODEL

Consider a simple economy with one seller who issues two different assets, *asset 1* and *asset 2*, that will generate nonnegative random cash flows of $\tilde{v}_1$ and $\tilde{v}_2$ in the future, respectively. There are also two different kinds of buyers in the economy; named by *buyer 1* and *buyer 2*. We assume that $\tilde{v}_1$ and $\tilde{v}_2$ are identically and independently distributed between 0 and $\bar{v}$ according to a density function $f(v)$, which is common knowledge to all agents in the economy. We further assume that the seller has complete information about the realized cash flows generated by the two assets, denoted by v_1 and v_2, respectively. By contrast, the buyers have less information regarding the assets. Buyer 1 is informed about the realized value of asset 1 only, whereas buyer 2 is informed about the realized value of asset 2. The buyers have no information about the other asset except its density function $f(v)$. For all agents in our economy, the values of assets satisfy the additive condition, that is, the total value for an agent who owns both assets is the sum of the two asset values.

Market interest is normalized to be zero. Therefore, in the absence of market imperfections, the total reservation price of the two assets for the seller is $v_1 + v_2$. However, we assume that the market has some imperfections. For example, most banks and institutions in the financial service industry may face credit constraints or minimum-capital constraints. As a result, the seller is willing to sell these assets at a discount in order to raise cash. In particular, we assume that the seller discounts future cash flows at a rate of $\delta \in [0,1]$. Consequently, if the seller retains the assets, they have a private value of $\delta(v_1 + v_2)$. In this model, assets are sold at an auction. The standard sealed-bid second-price auction is considered in this preliminary model.

Moreover, since the seller is more informed than the buyers, the latter may make inferences from the seller's selection of package format to get more specific information about the realized values of the two assets. Hence, the interaction and the information delivery between the seller and buyers should be considered when solving this model.

EQUILIBRIUM AND OPTIMAL SELLING FORMAT

In this section, we compare bundled selling with separate selling. We find that when the seller decides to sell the two assets separately, the buyer without information regarding the target will not participate in the transaction of the given asset, which will lead to a bilateral environment where the seller has to bargain with the informed buyer. On the other hand, if the seller sells both assets as a bundle, bidding competitively is profitable for the buyers. Finally, we will show the conditions that determine the seller's choice of package format. Since the seller is more informed about the assets than the two buyers, the choice of package formats can be regarded as a tactic to take advantage of the buyers' inferior information.

Selling the Assets Separately

In this subsection, we consider the situation in which the seller sells the assets separately. Because of the symmetry between the two assets, we start with an analysis of the sale of asset 1. When the second-price auction is held, it is easy to check that truth-bidding is a weak dominant strategy for buyer 1, which drives buyer 2, who has no information about asset 1, out of the auction. From the viewpoint specifically of buyer 2, he knows that buyer 1 will submit the true value v_1 as his bid. If he wins in the auction, buyer 2 will pay v_1 and collect zero profit. If he loses in the auction, the profit will also be zero. Since the best result for buyer 2 in the auction is zero, he has no incentive to put in his bid for asset 1. By the same token, buyer 1 does not participate in the auction in which asset 2 is on sale.

Having got the point when the seller decides to sell the assets separately, only the informed buyer would like to participate in the transaction. We know that the seller and the informed buyer are dwelling in a bilateral monopoly scenario. In this market, the asset is traded through bargaining between the informed buyer and the seller. Since the value of asset i for the seller and informed bidder i is δv_i and v_i, respectively, where $i = 1, 2$, the final deal price in the transaction will be somewhere between the two values, depending on both parties' bargaining power. When the seller's bargaining power is stronger, the price will be closer to v_i. In the opposite case, the price will be closer to δv_i. As a result, the final deal price can be the weighted average between δv_i and v_i, and we denote it as

$$p_i = \eta v_i + (1-\eta)\delta v_i \tag{1}$$

where the weights η and $1-\eta$ represent the bargaining power of the seller and the informed buyer, respectively. Note that $\eta \in [0,1]$, and that the closer η is near to 1, the stronger the seller's bargaining power will be.

We are now in a position to calculate the total revenues that the seller will receive when he sells the assets separately. By imposing the additive assumption, the total revenues will be

$$R_s = p_1 + p_2 = \eta(v_1 + v_2) + (1-\eta)\delta(v_1 + v_2) \quad (2)$$

As we can see in Equation (2), the total revenues collected from the sales of the assets increase with the seller's bargaining power. Equation (2) seems to imply that the issuer has the same bargaining power for each asset. Actually, it is allowed to let the issuer have different bargaining power when selling different assets to different buyers. We can regard η as the average bargaining power of the issuer for each asset. To be specific, suppose that η_1 and η_2 represent the bargaining power of the issuer for each asset, the total revenues from the sale of the two assets are $R_{ds} = \eta_1 v_1 + (1-\eta_1)\delta v_1 + \eta_2 v_2 + (1-\eta_2)\delta v_2$. If we replace η by $[\eta_1 v_1 + \eta_2 v_2]/[v_1 + v_2]$ in Equation (2), we get the same result as R_{ds}, where η is the average of η_1 and η_2 weighted by the ratio of each asset's value relative to the aggregate value of total assets.

Intuitively, it is unfavorable for the seller to sell the assets separately when his bargaining power is weak. At this time, bundled selling may be a wise decision for the seller to redeem himself from his weak bargaining ability because it stimulates buyers to compete. From now on, for the sake of simplicity, we focus on the case where the seller has zero bargaining ability, i.e. $\eta = 0$. The case where $\eta > 0$ will be discussed at the end of this paper.

Bundling the Assets

In this section we consider an alternative case where the seller bundles the assets and sells them as a composite by holding a second-price auction. When the assets are sold as a bundle, since each buyer has partial information about the true value of the whole bundle and offers his bid according to this information with profitable revenue, bundled selling makes both buyers bid competitively in the auction.

The bundling environment can be modeled as a affiliated common value auction, where the realized value of the object for each buyer is $v = v_1 + v_2$. To take the true value of the object as the average or summation of the signals of bidders is usually adopted in the common value auction literatures, such as Bikhchandani and Riley (1991), Albers and Harstad (1991), Krishna and Morgan (1997), Klemperer (1998), Bulow and Klemperer (1996), and Georee and Offerman (2003).

Since buyer 1 and buyer 2 are symmetric, we can just analyze the bidding behavior of bidder 1 and then extend the result to the behavior of bidder 2. In the affiliated model, $v(x, y) = E[v \mid v_1 = x, y_1 = y]$, which refers to the expected value function of the object conditional on bidder 1's own signal x and the highest signal of the remaining bidder's signal y, plays an important role in featuring the bidders' bidding strategies. In our model, the expected value function can be expressed as $v(x, y) = x + y$.

Following Milgrom and Weber (1982), the optimal bidding strategy for a bidder who receives the signal x is $b_s = v(x, x)$ in the second-price auction. This means that a bidder, say buyer 1, with signal x is asked to bid an amount such that if he were to just win the auction with that bid he would just break even. Thus, in our model a bidder with signal v_i submits a bid equal to

$$b_s(v_i) = 2v_i \quad (3)$$

However, since the seller is more informed, his decision for the package format of assets may reveal extra information to the buyers, and the revealed information may affect the bidders' bidding strategy. In Proposition 1 we have the result that bidders will still follow the bidding strategy as in Equation (3) even if the effect of the revealed information is considered. Furthermore, it is easy to check that the expected revenue for each buyer is positive (The proof in detail can be found in Milgrom and Weber (1982)).

Proposition 1: Each buyer's optimal bidding strategy will follow Equation (3), even if the information revealed by the seller's selling decision is considered.

Proof of Preposition 1: When bundling is held by the seller, buyer 1 infers that the revenue from auction $2v_2$ is at least as large as that from negotiation, δv, where $v = v_1 + v_2$. Thus, he knows that $2v_2 \geq \delta(v_1 + v_2)$ if he wins. The relationship between the two values can be further rearranged as $\frac{\delta}{2-\delta} v_1 \leq v_2$, meaning that bundling signals to him that the lower bound of bidder 2's signal is rising from zero to $\frac{\delta}{2-\delta} v_1$. Hence, buyer 1's revenue can be expressed as

$$R_1 = \int_L^{\hat{b}^{-1}(b_1)} (v_1 + v_2 - b(v_2)) f(v_2) dv_2 \,, \quad \text{where } L = \frac{\delta}{2-\delta} v_1 .$$

To find the optimal bidding strategy, the first-order condition should be zero when $\hat{b}(v_1) = b_1$. After rearrangement, we have the following equation

$$(v_1 + \hat{b}^{-1}(\hat{b}(v_1)) - \hat{b}(\hat{b}^{-1}(\hat{b}(v_1))) f(\hat{b}^{-1}(\hat{b}(v_1))) = (2v_1 - \hat{b}(v_1)) f(v_1) = 0 .$$

It is easy to verify that the solution is $\hat{b}(v_1) = 2v_1$, which is the same as the strategy derived without considering the information revealed by the seller's selling decision.

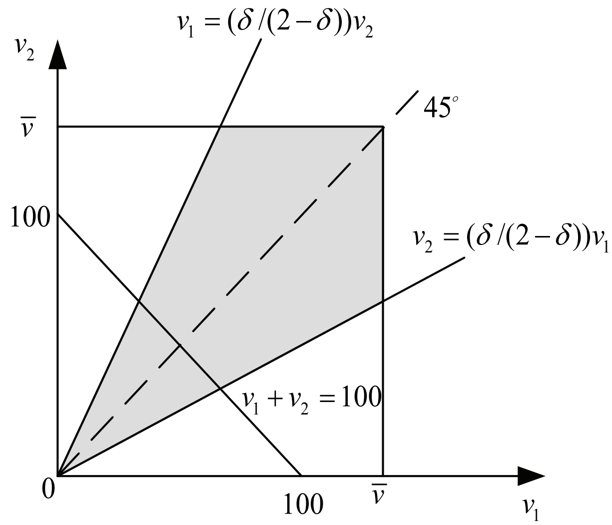
The buyers' bidding strategy is common knowledge and the seller has precise information about the buyers' signals, so that the revenue generated from the sale of the bundled assets is a certainty for the seller, which is represented as

$$R_b = \min\{2v_1, 2v_2\} \tag{4}$$

<u>The Package Decision: Bundling vs. Separating</u>

In what follows, we compare the revenue from bundling with that from separating, and try to find out how the seller picks the optimal selling option. Proposition 2 states the criteria that help the seller make the decision.

Figure 1: Relation between the Values of Assets and the Package Formats



This figure relates the optimal selling format to the values of the assets. The seller selects bundling strategy when both assets' values fall into the shaded area. This implies that the seller prefers bundling when the difference of assets' values is small.

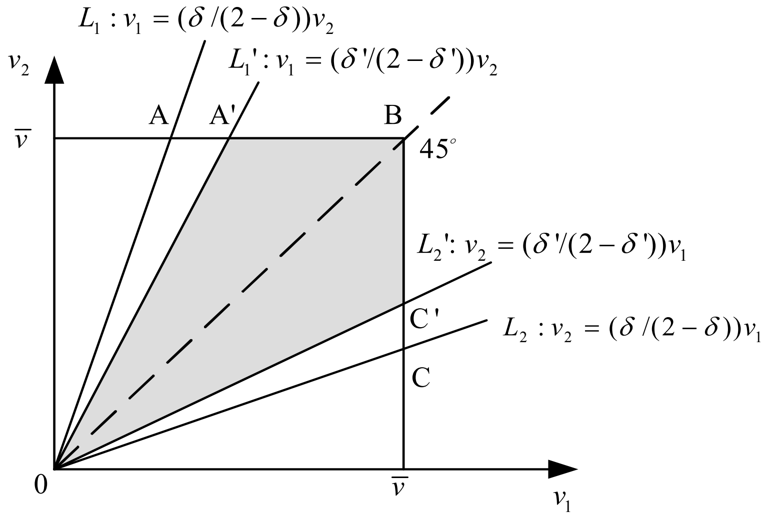
Proposition 2: Let $v_m = \min\{v_1, v_2\}$ and $v_M = \max\{v_1, v_2\}$, where v_1 and v_2 are realized values of $\tilde{v}_1$ and $\tilde{v}_2$. For all $0 < \delta < 1$, when $v_m > \delta v_M /(2-\delta)$, the seller can be better off by selling the assets in a bundle. By contrast, when $v_m < \delta v_M /(2-\delta)$, the seller will prefer to sell the assets separately. In the case with $v_m = \delta v_M /(2-\delta)$, the seller will be indifferent between choosing the two selling formats.

Proof of Proposition 2: Although bundling reveals information to the bidders, the bidding strategy of the bidder does not change, as shown by the results in Proposition 1. Thus, the revenue from selling assets in a bundle is $2v_m$, and that from selling assets separately is $\delta v_M + \delta v_m$. Since the seller sells assets in a bundle when bundling generates higher revenue than separating, we have the relationship denoted by $\delta v_M + \delta v_m < 2v_M$. Thus, the seller would rather sell the assets in a bundle when $v_m > \delta v_M /(2-\delta)$. By the same token, separating is held by the seller while $v_M > (2-\delta)v_m / \delta$.

The optimal selling format involves a tradeoff between a bargaining cost and a bidding cost for the seller. When selling separately, the seller suffers a bargaining cost, which results from the fact that the seller has to sell the assets at a discounted price of δv. The cost decreases when δ increases. On the other hand, when selling as a bundle, the seller faces a bidding cost, which results from the fact that the high-valued asset is sold at a price equal to the low-valued asset. This cost can be estimated by the relative value of both assets. The bigger the difference between the values of both assets, the higher the bidding cost will be. Because the optimal format is the one that costs the seller less, the seller's incentive to use bundled (separate) selling increases when δ decreases (increases) or the difference between the values of both assets decreases (increases).

Figure 1 relates the optimal selling format to the values of the assets. In this figure, the seller selects bundling when both assets' values fall into the shaded area. It is verified that the decision for the selling format depends on the relative value of both assets. Moreover, we find that the total value of assets does not affect the seller's decision. To be clearer, we provide a numerical example for illustration. Given that $\delta = 0.5$, we compare the following two cases. We let $v_1 = 90$ and $v_2 = 10$ in case (1), while $v_1 = 50$ and $v_2 = 50$ in case (2). In case (1), the revenue is 50 from the separate selling, while it is only 20 from the bundled selling. Thus, separate selling is better in case (1). In case (2), the revenue is also 50 from separate selling. However, the seller receives 100 from the bundled selling. Thus, in case (2) the seller prefers to sell the assets as a bundle. Therefore, the seller may choose different selling formats even when the aggregate value of the assets is the same. The result can be shown in Figure 1, and it is obvious that the line $v_1 + v_2 = 100$ pass through both areas.

Figure 2: The Relation between the Seller's Bargaining Ability and the Package Formats



This figure relates the optimal selling format to the seller's bargaining ability. The increase in the seller's bargaining ability reduces the bundling area from OABC to OA'BC'. This implies that the increase in the seller's bargaining power weakens his tendency to adopt a bundling format.

Now, we go back to the situation where the seller's bargaining ability η is not zero. After rearranging Equation (1) and replacing $\eta + (1-\eta)\delta$ with δ', we have

$$R_s = p_1 + p_2 = \eta(v_1 + v_2) + (1-\eta)\delta(v_1 + v_2) = (\eta + (1-\eta)\delta)(v_1 + v_2) = \delta'(v_1 + v_2) \tag{5}$$

which implies that the increase in the seller's bargaining ability can be regarded as the increase in the discounted factor value from δ to δ'. In Figure 2, we can see that the increase in the seller's bargaining ability reduces the bundling area from OABC to OA'BC'. Consequently, we can conclude that the increase in the seller's bargaining power weakens his tendency to adopt a bundling format.

CONCLUSION

In this paper, we consider how a seller, who owns multiple assets, increases his advantage through packaging strategies. We set up an environment where the seller has two assets and wants to sell them because of some financial constraints. The seller is precisely informed all of the values of his assets while each buyer has only partial information concerning the assets. In this environment, selling separately drives out the buyer who has no information regarding the item being sold, which leads to a bilateral environment where the seller has to bargain with the informed buyer. If the seller's bargaining ability is weak, separate selling may not be a wise decision for him. On the other hand, if the seller sells both assets as a bundle, since each buyer possesses some information regarding the object, and bidding according to this information enables the buyers to generate positive revenue, bundled selling leads to a competitive bidding environment.

We establish a theoretical model in this study and show that package formats selected by the seller can be regarded as tactics for him to take his information advantage when a monopolistic seller is better informed than the buyers. Our study provides two primary results. First, by giving the seller's bargaining ability, we show that the relative value of assets determines the selling format. In specific, the smaller the difference between the two assets' values, the greater the tendency will be for the seller to sell the assets as a bundle. Second, given the assets' values, the increase in the seller's bargaining power weakens the seller's incentive to use bundled selling.

We consider only the sealed-bid second-price auction to obtain the preliminary results. To arrive at a clearer picture in the scenario we set, it is necessary to discuss the effects of introducing other auction formats, e.g., the first-price auction and English auction. This will be left to our future research.

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IS THERE A SYNCHRONICITY BETWEEN THE PHILIPPINE STOCK EXCHANGE AND NEW YORK STOCK EXCHANGE?

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ABSTRACT

This study examines the impact of macroeconomic variables such as real Gross Domestic Product, inflation rate, savings interest rate, foreign exchange rate, oil price and economic disturbances including the 9/11 incident and Internet bubble burst on the Philippines Stock Exchange composite price index. Using multiple linear regression analysis, monthly data from 1996 to 2006, including 119 observations were analyzed. The results indicate that rates of inflation, savings interest rate, foreign exchange rate and oil price significantly affected the Philippine Stock Exchange composite price index. The lag one first difference in the unit root test revealed stability of Philippine Stock Exchange (PSE) market and New York Stock Exchange (NYSE) market. Both were also found to be significantly affected by the two economic disturbances. Likewise, we find synchronicity between PSE and NYSE markets using Granger-causality test. Specifically, causality runs one-way from NYSE stock prices to PSE stock prices.

JEL: E44, G15

INTRODUCTION

The Philippine history reveals a deep relation to USA. Consequently, there are similarities in some of their economic and political policies. Among indicators of a country's economy is the performance of the Stock Market. This measure is often referred to as a barometer of the present condition of a country. The development of the stock market is, therefore, a function of economic growth. When economic fundamentals are bullish, traders will buy more stocks. If it is bearish, they will sell stocks leading to a stock price decline, ceteris paribus. The New York Stock Exchange (NYSE), is the largest stock exchange market in the world by dollar volume and the second largest by number of companies listed. Any fluctuation in this large market could potentially influence smaller stock markets such as the Philippine Stock Exchange (PSE). Globalization has increased the interconnection of stock markets. In addition, because of this phenomenon, stock markets have easy access to each other, with some stock issues trading on multiple markets.

In theory, increasing GDP and depreciating foreign exchange rate will result in a positive impact to the economy of an exporting country if its stock market is bullish. On the other hand, the inflation rate has an adverse effect on the stock market because of its negative correlation to future output. Increasing savings interest rates also have a negative effect on the stock market. If the interest rates increase, the higher expected returns will entice investors to deposit their money instead of investing in the stock market. Lastly, increasing oil prices have a negative impact on the stock market because it slows down the investments as it makes the business operations costly.

The examination of factors affecting the stock market is an interesting subject. However, little is known about granger-causality analysis and the effect of economic disturbances in the Philippine context. This paper fills this gap in the literature. There are three objectives of this article. First is to examine the impact of selected macroeconomic factors and economic disturbances on the PSE. The macroeconomic factors examined are real Gross Domestic Product (GDP), inflation rate, foreign exchange rate, oil price, and savings interest rate. The economic disturbances examined are the Internet bubble burst in 2000 and the 9/11 incident in 2001. The second objective is to determine the direction of the synchronicity, if there is, between the PSE and NYSE. Lastly is to project the future fluctuations in PSE. The remainder of this article is organized as follows: The following section reviews the past literature. The next two sections present the methodology and test results. The paper closes with some concluding comments.

LITERATURE REVIEW

Studies have produced conflicting theories and empirical results regarding the relationship between stock market prices and economic variables. The study of Eun and Shim (1989) on non-economic factors carefully examines international stock market movements in recent years and suggests that there is an existing substantial degree of interdependence among international stock markets. Choi (1991) found that Gross National Product (GNP) has a positive effect on stock prices while increasing inflation and interest rates, and depreciating foreign exchange rate have negative impacts.

According to Pane (1995), certain advantages were realized as a result of the unification of the Manila and Makati Stock Exchanges in the Philippines. The combination resulted in a larger and more organized market. He found a negative relation between stock price and interest rates, inflation rate and stock transaction tax. However, he finds that there is some positive relation between the inflation rate and taxes. All variables with the exception of interest rate are strongly correlated with the stock price.

Alcaly (2003) noted in his book that neither overly inflated share prices nor seriously undervalued ones can persist. Accordingly, the economy and stock price forecasts change in a fast and unstable manner over the end of the period analyzed. Internet stocks that once rose by more than 200% fell dramatically because of the 9/11 incident. It recovered in the next three months before it fluctuated in the spring and summer of 2002. The working paper of Gupta et al. (2000) in Jakarta, Indonesia found an irregular unidirectional causality between closing stock prices and interest rates and vice versa. And, there is weak unidirectional causality from exchange rates to stock prices. However, they failed to establish a consistent causal relationships between any of the economic variables.

Ofek and Richardson (2003) point out two reasons for the decline in stock prices. First, the optimistic behavior of investors and analysts was a factor. Second, deception by business firms in an attempt to maintain their ratings and stock prices impacted stock prices. Kawamoto (2005) argued that because of the aftermath of the internet bubble burst, technology has taken a step back as investors were disheartened by the event.

Two separate studies in stock market integration have been conducted in India. Ahmad (2005) found that there is no long-term relationship between the Indian equity market and the US and Japanese equity markets. Using daily closing data from January 1999-August 2004, he found that these stock markets have no tendency to move together in the long-term but a strong causal relationship was established in 1999-2001. In the case of Greece, Floros (2008) found out that there is a long-run relation between open interest and future prices in the Greek Stock Index Future Market. Mishra (2004), on the other hand, established that both interest rates and exchange rates affect stock returns in India.

Zahid (2000) created an econometric model based on the relationship between money supply, GNP, exchange rate, inflation, interest rate, stock transaction taxes, GDP, consumer price index, growth rate and unemployment and the composite stock index in emerging Association of Southeast Asian Countries (ASEAN) countries. He categorized the fluctuation in the stock index of any emerging ASEAN countries into economic variables, non-economic variables and financial factors. Some of the non-economic factors cited are government policy, political stability, disclosure of corporate and governmental information, and psychology of the investor. The major significant variables that positively influenced the composite stock index are foreign exchange rate, consumer price index, and GNP. Only the lending rate is found to negatively influence the composite stock index but to a less significant degree. Mishra (2004), found that the exchange rate return affects the demand for money and stock returns and that interest rate changes cause exchange rates to change.

DATA AND METHODOLOGY

Monthly observations from the period 1996 to 2006 were used in this study. Quarterly real GDP data were transformed into monthly average data using 1985 as a base year. This period was selected to capture the effects of economic disturbances that occurred during this time period. The PSE composite price index in Philippine peso are obtained from PSE. The NYSE composite price indices under the 500 base values in dollars were gathered from its website. The savings deposit interest rate, inflation rate and foreign exchange rate are in monthly weighted average format. The world's price of oil is in dollar per barrel.

The data are analyzed using multiple linear regression with the aid of Eviews version 4 and PhStat software. In determining the extent of the impact of the selected macroeconomic factors on the PSE, regression model 1 was used. The impact of the 9/11 incident and Internet bubble burst was measured in regression models 2 and 3, respectively. Both economic disturbances are dummy variables. The variables used here are PSE= Phil. Stock Exchange composite price index, GDP= gross domestic product, IR= inflation rate, Int= savings interest rate, FER= foreign exchange rate, OP= oil price, NYSE= New York Stock Exchange price index, INC= 9/11 incident, and IBB= Internet Bubble burst.

$$PSE = \beta_0 + \beta_1 GDP - \beta_2 IR - \beta_3 Int + \beta_4 FER - \beta_5 OP + \mu \quad (1)$$
$$PSE = \beta_0 + \beta_1 NYSE - \beta_2 INC + \mu \quad (2)$$
$$PSE = \beta_0 + \beta_1 NYSE - \beta_2 IBB + \mu \quad (3)$$

An analysis of time series data requires a preliminary examination of their stationarity status. All variables should be integrated of order zero, *I(0)*. The unit root test ensures that the probability structure of the series is stable over time and it makes statistical sense to combine them in a regression equation. To test for stationarity, we use the Dickey-Fuller Test which is a special test of the augmented Dickey-fuller test. The former becomes the latter when the number of the lagged first difference terms in the test regression is chosen to be zero (Danao, 2005). The existence of a unit root as indicated by the Durbin-Watson *d* value justifies using the augmented Dickey-Fuller test (ADF) where its null hypothesis is that Y_t is a non-stationary series. The asymptotic distribution of the ADF test statistic qualifies for the use of MacKinnon critical values which are readily provided by Eviews at 1%, 5%, and 10% significance levels. When the ADF test statistic is greater than the MacKinnon critical value, the null hypothesis is rejected. Its formula is the *t* statistic on ζ (zeta) as seen below. The difference operator is represented by Δ. ε_t represents the stationary random error, and Y_t refers to the dependent and independent variables

$$\Delta Y_t = \beta_1 + \beta_2 t + \zeta Y_{t-1} + \alpha \sum \Delta Y_{t-1} + \varepsilon_t \quad (4)$$

To determine the extent of synchronicity between the stock price index of NYSE and PSE, the Granger-causality test was applied. The Granger-causality test measures precedence and information content. Consequently, Y is said to be Granger-caused by X if the coefficients of X are statistically significant (Granger, 1981). The Granger-causality test is then the Wald test with the null hypothesis that Y_t does not granger-cause X_t. The bivariate regression that is run by Eviews for this study are in these forms:

$$PSE_t = \sum_{i=1}^{k} \alpha_1 PSE_{t-1} + \sum_{j=1}^{k} \beta_j NYSE_{t-j} \quad (5)$$

$$NYSE_t = \sum_{i=1}^{k} \alpha_1 \, NYE_{t-1} + \sum_{j=1}^{k} \beta_j \, PSE_{t-j} \tag{6}$$

The PSE and NYSE are two stationary series with *i* and *j* as lag lengths. The expected outcome would be one of the following (Danao, 2005):

> *X Granger-causes Y* (the estimated βs are jointly significantly different from 0) but *Y does not Granger-cause X* (the estimated δs are jointly not significantly different from 0).
>
> *Y Granger-causes X* (the estimated δs are jointly significantly different from 0) but *X does not Granger-cause Y* (the estimated βs are jointly not significantly different from 0).
>
> *X Granger-causes Y* (the estimated βs are jointly significantly different from 0) and Y Granger-causes X (the estimated δs are jointly significantly different from 0).
>
> *X does not Granger-cause Y* (the estimated βs are jointly not significantly different from 0) and *Y does not Granger-cause X* (the estimated δs are jointly not significantly different from 0).

Finally, after finding the significant macroeconomic factors and economic disturbances affecting PSE stock price, we predict the future fluctuations of PSE using regression model 7.

$$PSE = \beta_0 + \beta_1 NYSE + \beta_2 GDP - \beta_3 IR - \beta_4 Int + \beta_5 FER - \beta_6 OP - \beta_7 AFC - \beta_8 IBB + \beta_9 INC + \mu \tag{7}$$

RESULTS AND DISCUSSION

We organize our results based on five propositions that are examined. The first proposition is that higher savings interest rate (IR), foreign exchange rate (FER), inflation rate (IR), oil price (OP) and real GDP influence the PSE stock price. The second proposition is that there is synchronicity between Philippine Stock Exchange and New York Stock Exchange. The third proposition is that the PSE will be affected by the internet bubble and 9/11 incident. The fourth proposition is that the presence of NYSE composite price index and Internet bubble burst will lead to more changes in PSE composite price index. Finally, notable changes in significant macroeconomic factors and economic disturbances will lead to more changes in PSE composite price index.

Proposition 1- Higher Savings Interest Rate (IR), Foreign Exchange Rate (FER), Inflation Rate (IR), Oil Price (OP) and Real GDP Influence the PSE Stock Price: In order to measure the contribution of each macroeconomic variable to PSE stock price, multiple regression analysis was used. The results show that the majority of these variables exhibited a significant impact on the PSE stock price except for real GDP. Hence, we tested a new econometric model without GDP. Table 1 presents the summary of the final regression result of all significant variables affecting the PSE. The findings show that 96% of the changes in PSE stock price can be explained by the savings interest rate, inflation rate, foreign exchange rate, and oil price. Furthermore, 95% in total variation of PSE stock price index can be explained by the delineated factors mentioned. The *p* value for F statistics is zero indicating that the regression as a whole is significant at 5% level. The results show that there is an estimated average of Php 3,914.96 in the PSE stock price index when there are no effects of savings interest rate, inflation rate, foreign exchange rate, and oil price presence. There is an estimated decrease of Php 36.43 in the average value of PSE stock price index for every one percent increase in savings interest rate, *ceteris paribus*. An estimated increase of Php 41.98 in the average value of the PSE stock price index is expected for every one percent increase in the inflation rate. An estimated decrease of Php 54.64 is expected for every one percent appreciation of the foreign exchange rate. Lastly, an estimated increase of Php 14.64 is expected for every one dollar increase in oil price per barrel.

The positive effect of inflation may be due to the notion that if inflation is increasing, people tend to buy more stocks rather than save money as a hedge against inflation. The finding is analogous to the studies made by Choi (1991), Pane (1995) and Gupta et al. (2000) but contrary to Zahid's (2000) studies on the relationship between interest rates and stock price.

Table 1: Multiple Regression Result

Delineated Factors (n = 119)	Coefficient Value	t
Intercept	3,914.958	
savings interest rate	-36.4281	
inflation rate	41.98441	10.19016*
foreign exchange rate	-54.6389	-2.72578*
oil price	14.63994	3.15978*
R^2	0.959226	-6.82426*
Adjusted R^2	0.952400	4.717079*
Durbin Watson	2.023062	
F	139.8820	

** denotes significance at 5% level*

The foreign exchange rate is negatively related to stock prices. The Philippine peso continuously depreciated during the time period covered. This could partly be explained by the imbalance of trade. As oil prices increase, the general price level will increase resulting in higher inflation. Oil shocks that cannot be anticipated by nature are also an explanatory factor because a sudden increase in oil prices in the short run will lead to a lower profit among business companies. However after they have adjusted, in the long-run, they could increase their prices and recover from their losses in the previous period.

The findings in Table 2 show that oil prices and foreign exchange rates are the two factors that could influence most the PSE stock price. Oil is a necessity to countries that utilize motor related things in order to generate power (Yee, 1995). The service sector, which consumes a large portion of the total electricity consumption in the Philippines, is a substantial contributor to gross domestic product. Consequently, performance in the stock market is affected whenever there is an oil price increase. Moreover, because of the sophistication of financial markets, an investor could compare the expected returns in different currencies, leading to a separate foreign exchange effect.

Table 2: Most Influential Variables

Delineated Factors (n = 119)	Coefficient Value	p value
oil price	14.6399	0.0000*
savings interest rate	-36.4281	0.0076*
inflation rate	41.98441	0.0021*
foreign exchange rate	-54.6384	0.0000*

** denotes significance at 5% level*

Proposition 2-There Is Evidence of Synchronicity between Philippine Stock Exchange and New York Stock Exchange: Our analysis continues by examining the synchronicity between the Philippine and New York Stock exchanges. Table 3 shows the results for testing unit root at the level variables of PSE and NYSE stock prices. The researchers conducted a test of stationary since the stock prices are always in random motion making it hard to calculate a specific prediction of their movement. The initial test shows that both variables have unit root correlation so the ADF was applied. The third column shows the unit root test after differencing the data once. In this case, the ADF statistic in absolute term is now greater than the critical values of MacKinnon at 1%, 5% and 10% level of significance indicating that the PSE and NYSE stock prices are now stationary. Hence, variables can now be tested using granger-causality test.

Table 3: Unit Root Test

Delineated Factors	Level Variables	First Difference
PSE	-1.681800	-6.576948*
NYSE	-0.385780	-7.684460*

The critical values for ADF are -3.4819, -2.8838,and -2.5785 for 1%, 5%** and 10%*** level of significance, respectively.*

The results of the Granger-causality test are presented in Table 4. Using 128 observations we find that there is synchronicity between PSE and NYSE. The *F* statistic (2.72512) reveals that NYSE granger-causes PSE and it appears that granger-causality runs one-way only from NYSE to PSE stock price. The availability of information technology makes the possibility of market interconnections. In the case of the Philippines and USA, the influence of NYSE to PSE stock price can be ascribed to the political influence of the latter.

Table 4: Granger-causality Test

Propositions (n = 128)	F-statistic	P value
NYSE does not granger-cause PSE	2.72512	0.10129*
PSE does not granger-cause NYSE	0.21046	0.64720

** denotes significance at 10% level*

Proposition 3-The PSE Will Be Affected by the Internet Bubble and 9/11 Incident: Next, we examine how the PSE is affected by shocks to the system. Table 5 reveals that the NYSE and 9/11 incident explain 96% of the changes in PSE stock price index. The Durbin Watson *d* statistic of (2.047) indicates that there is no autocorrelation among the data. There is an estimated decrease of Php 460.43 in the PSE stock price index when the NYSE stock price index and the 9/11 incident are non-existent. Moreover, there is an estimated average increase of Php 2.99 in the PSE stock price index for every one dollar increase in the NYSE stock price index, *ceteris paribus.* The PSE stock price index after the 9/11 incident is higher by Php 238.08. The possible explanation for this is a higher demand from foreign investors as they transfer their investments to other countries like the Philippines after this incident.

Table 5: Regression Result with 9/11 Incident

Delineated Factors (n = 119)	Coefficient Value	t
intercept	-460.4317	-1.06322
New York Stock Exchange	2.99106	6.136635*
9/11 incident	238.077	2.104448*
R^2	0.95951	
Adjusted R^2	0.95411	
Durbin Watson	2.04705	
F	177.7373	

** denotes significance at 5% level*

Proposition 4-The Presence of NYSE Composite Price Index and Internet Bubble Burst Will Lead to More Changes in PSE Composite Price Index: Next we examine the extent to which the PSE is explained by other economic factors. The results are presented in Table 6. The regression results indicate that the decrease in the PSE stock price by Php 557.14 can be explained by other factors not related to the NYSE stock price and Internet bubble burst. The R^2 tells that 96% of the changes in PSE stock price can be explained by these two significant factors. The *p* value for the F statistic of zero means that the regression model as a whole is significant at the 5% level. The value of the Durbin Watson tells that data are free from autocorrelation. The PSE stock price after the Internet bubble burst increased by Php 259.89. We attribute this change to a higher inflation rate and a greater demand for PSE than NYSE stocks because of their lower vulnerability to external shocks. Likewise, there is an estimated average increase of Php 3.06 in the PSE stock price for every one dollar increase in NYSE stock price.

Table 6: Regression Result with Internet Bubble Burst

Delineated Factors (n = 119)	Coefficient Value	t
intercept	-557.1376	-1.231702
NYSE	3.062487	6.075735*
IBB	259.8871	2.130792*
R^2	0.962119	
Adjusted R^2	0.955227	
Durbin Watson	1.991837	
F	286.9991	

** denotes significance at 5% level*

Proposition 5-Notable Changes in Significant Macroeconomic Factors and Economic Disturbances Will Lead to More Changes in PSE Composite Price Index: Finally, we examine how changes in significant macroeconomic factors and economic disturbances affect the PSE composite price indexes. Table 7 shows the final regression result between the PSE stock price index movement and the significant variables namely: NYSE stock price index, savings interest rate, inflation rate, foreign exchange rate, oil price and the 9/11 incident. The regression included 119 observations. The Durbin Watson *d* statistic of (2.05) indicates that there is no autocorrelation. The adjusted R^2 indicates that 96% of the total average variation of PSE stock price index can be explained by these factors. Likewise, findings show that 97% of the changes in PSE stock price are attributable to them. Since all significant variables have computed t values that are greater, in absolute value, to the tabular value of *t* that is 1.96, it is safe to say that they are all significant in explaining PSE stock price index movement prediction.

No multicollinearity was detected in the process. Also, F statistic indicates that the regression as a whole is significant at the 5% level. There is an estimated Php 2,718.62 of PSE stock price index that is not attributable to the movement of NYSE stock price index, savings interest rate, inflation rate, foreign exchange rate, oil price and the 9/11 incident. Moreover, there is an estimated average increase of Php 2.11 in the PSE stock price index for every one dollar increase in the NYSE stock price index, *ceteris* paribus. An estimated increase of Php 27.46 is also expected to take place for every one percent increase in the inflation rate. Furthermore, for every one percent increase in the savings interest rate, the PSE stock price is expected to decline by Php 29.9. PSE stock price is expected to increase by Php 6.83 for every one dollar increase in oil price per barrel. The average value of the PSE stock price index decreases by Php 54.30 for every one percent appreciation of foreign exchange rate. In addition, the PSE stock price is higher by Php 178.45 after the 9/11 incident.

Table 7: Multiple Regression Result

Delineated Factors (n = 119)	Coefficient Value	t
intercept	2,718.623	5.2749*
New York Stock Exchange	2.110957	5.5087*
savings interest rate	-29.8982	-2.3517*
inflation rate	27.4641	2.2478*
foreign exchange rate	-54.2995	-6.2940*
oil price	6.82521	2.3689*
9/11 incident	178.4454	2.0044*
R^2	0.969466	
Adjusted R^2	0.963606	
Durbin Watson	2.059496	
F	165.4342	

** denotes significance at 5% level*

CONCLUDING REMARKS

While there are few attempts to study the possible connection of a stock market from a developing country to a more sophisticated stock market, this study is different because it considered the effects of economic disturbances that are unpredictable and sometimes difficult to quantify. This paper provides more and new evidence to aid in the understanding of the movements in a stock market of a developing country like the Philippines. It is noteworthy that this study is a sample of the general concept as to what happens in a smaller stock market in relation to a larger stock market.

We find that there is synchronicity between the PSE and NYSE stock markets. Specifically, the NYSE price Granger-causes the PSE price. Future research could explore the synchronicity of the PSE to other stock markets in Asian developing countries. In this way, two-way direction causality might be established. The time period covered might also be extended to capture the political aspects of the country. Finally, while the NYSE composite price index granger-causes the PSE composite price index, it is noteworthy that other significant factors can influence the movement of the latter. Hence, it is very important that the Philippine government maintain healthy domestic policies in order to maintain the stability of savings interest rate, inflation rate, foreign exchange rate, and oil price.

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THE DETERMINATION OF THE COSTA RICA COLON/USD EXCHANGE RATE

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ABSTRACT

The purpose of this paper is to compare four major exchange rate models for the Costa Rica Colon. We examine exchange rate data for the Costa Rica/U.S. dollar relationship from 1981-2007 and find that monetary models have a higher explanatory ability whereas the Mundell-Fleming model performs better in forecasting exchange rates than other models. The coefficient of the interest rate differential in the uncovered interest parity model has a wrong sign.

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INTRODUCTION

After many years of adopting the crawling peg exchange rate regime, the Costa Rican authorities have moved to a crawling band system and modified monetary policy and other procedures to foster exchange rate flexibility. The expected inflation rate has been used to determine the change in crawl of the exchange rate. A narrowed spread for foreign exchange transactions with the central bank has been opened to promote interbank market development. The central bank has established an electronic mechanism to handle transactions in foreign exchange market. New rules have been issued to raise the limit on changes in foreign exchange positions. The overnight rate has replaced the 30-day deposit rate to become the policy rate.

The Costa Rica Colon/USD exchange rate has depreciated in the long run from 8.57 in 1980 to 308.19 in 2000 and 517.9 in 2006. One possible cause for a weaker Costa Rica Colon (CRC) may be due to declining interest rates in order to stimulate consumption and investment spending. For example, the deposit rate in Costa Rica reached a high of 27.32% in 1991 and then declined to 13.38% in 2000 and 9.77% in 2006. The uncovered interest parity model suggests that a declining domestic interest rate relative to the world interest rate would cause the Colon to depreciate, holding other factors constant. Other possible reasons for a weaker Costa Rica Colon are a relatively high inflation rate, more money supply, and high output growth. During 2000-2006, the inflation rate, M2 money, and real GDP increased at an annual rate of 14.89%, 38.03%, and 5.31%, respectively.

This paper attempts to examine the behavior of the CRC/USD exchange rate and has several focuses. First, four different models are considered. They include the purchasing power parity model (Taylor and Taylor, 2004; Taylor 2006; Breitung and Candelon, 2005; Yotopoulos and Sawada, 2006; Alba and Papell, 2007), the uncovered interest parity model (Dekle, Hsiao, and Wang, 2002; Chinn and Meredith, 2004), the monetary models (Meese and Rogoff, 1983; Chinn, 1999, 2000; Cheung, Chinn, and Pascual, 2005), and the Mundell-Fleming model (Romer, 2001; Hsing, 2005, 2007). Second, in the monetary models, four different versions proposed by Dornbusch (1976), Frenkel (1976), Bilson (1978), and Frankel (1979) are compared. Third, in the Mundell-Fleming model, comparative-static analysis is applied to determine the impact of a change in an exogenous variable on the equilibrium exchange rate. Fourth, the Newey-West (1987) method is applied in order to address the issue of both autocorrelation and heteroskedasticity when their forms are unknown.

The remainder of the paper is organized as follows. In the following section we discuss the literature related to exchange rates. This section is followed by a discussion of the models that are tested in the paper. Two sections follow that discuss the data used for the empirical tests along with the results of the empirical tests. The paper closes with some concluding comments.

LITERATURE REVIEW

This section reviews several recent articles of exchange rate determination and related subjects. Jalbert, Stewart and Jalbert (2006) examine the efficiency and rate spread of the Costa Rica certificate of deposit (CD) market. Contrary to conventional wisdom, U.S. banks are found to pay higher rates than Costa Rican banks on dollar denominated CDs. Empirical tests reveal uncovered interest rate arbitrage opportunities. Breitung and Candelon (2005) show that before 1997, PPP holds for Asian countries but not for Latin American countries and that long-run PPP holds for Asian countries owing to flexible exchange rate systems and breaks down for South American countries due to long-time pegging to the dollar. Yotopoulos and Sawada (2006) reveal that PPP holds for 132 countries in a 20-year time period and for 105 countries in a 10-year time period. Applying the KSS (Kapetanios, Shin, and Snell, 2003) unit root test and based on a sample of 13 countries including Costa Rica, Francis and Iyare (2006) find that real exchange rates in most countries are nonlinear and stationary and that nominal exchange rates and relative prices are cointegrated.

Using a sample of 30 LDCs including Costa Rica, Holmes (2006) shows that 16 out of 30 countries exhibit nonlinearity in the real exchange rate. Based on a sample of 88 LDCs including Costa Rica and a new unit root test (KSS, 2003), Bahmani-Oskooee, Kutan and Zhou (2008) reveal that the number of countries that PPP holds are doubled, that there is nonlinear adjustment toward PPP in LDCs, and that PPP is more likely to hold for countries with relatively high exchange rate flexibility and high inflation. Alba and Papell (2007) indicate that PPP is valid for Latin American and European panel data, but not for Asian and African panel data. They also found stronger evidence of PPP for countries with more openness, lower inflation rates, moderate volatility of exchange rates, similar rates of economic growth as the U.S., and less distance from the U.S. Taylor and Taylor (2004) and Taylor (2006) review major previous works, present issues and challenges in verifying PPP, and maintain that long-run PPP has gained more support as the gap between theory and data and the deviation of exchange rates from PPP have narrowed.

Chinn (1999) reveals that the five Asian currencies under study are consistent with the specifications of some types of monetary models, that exchange rates do most of the adjustments toward equilibrium except for the New Taiwan dollar and the Thai baht, and that out-of-sample forecasts work well for the Korean won, the New Taiwan dollar, and the Singapore dollar. In another study, Chinn (2000) uses different models to evaluate currency overvaluation for several Asian currencies. As of May 1997, the PPP model shows that the Malaysian ringgit, the Thai baht, the Hong Kong dollar, and the Philippine peso were overvalued. A monetary model reveals that the Indonesian rupiah and the Thai baht are overvalued whereas the New Taiwan dollar, the Korean won, and the Singapore dollar are undervalued.

Applying an extended Mundell-Fleming model, Hsing (2005) finds that the real exchange rate in Slovakia is positively influenced by deficit spending/GDP ratio and the stock price index and negatively associated with real M2, the US Treasury bill rate, country risk, and the expected inflation rate. The error variance can be characterized by the GARCH process. Hsing (2007) shows that the US dollar/kuna exchange rate for Croatia is negatively associated with real M1, the US T-bond rate, the euro interest rate, the expected inflation rate, and the relative price and positively influenced by the expected exchange rate. Deficit spending does not affect the exchange rate. Most of the variation in exchange rates can be explained by the open economy model and uncovered interest-rate parity.

THE MODEL

This section presents four exchange rate models, namely, the purchasing power parity model, the uncovered interest parity model, the monetary models, and the Mundell-Fleming model.

The Purchasing Power Parity Model

In the purchasing power parity (PPP) model, the nominal exchange rate is a function of the relative price:

$$E = F(P / P^*) \tag{1}$$

where E, P, and P^* denote the CRC/USD exchange rate, the price level in Costa Rica, and the price level in the U.S. The sign of the relative price in equation (1) is expected to be positive, suggesting that a higher relative price would cause the CRC/USD exchange rate to rise or the Costa Rica Colon to depreciate against the U.S. dollar.

The Uncovered Interest Parity Model

In the uncovered interest parity (UIP) model, under the assumption of perfect capital mobility, the interest rate differential can be offset by the exchange rate depreciation or appreciation. If the domestic interest rate is greater than the foreign interest rate, then the domestic currency is expected to depreciate by the same magnitude. If the domestic interest rate is less than the foreign interest rate, then the domestic currency is expected to appreciate by the same magnitude. The UIP model can be expressed as

$$R = R^* + (E^e - E) / E \tag{2}$$

where R, R^*, and E^e stand for the interest rate in Costa Rica, the interest rate in the U.S., and the expected exchange rate. Expanding the second term on the right-hand side and moving E to the left-hand side and other terms to the right-hand side in equation (2), in general form, the nominal exchange rate is a function of the interest rate differential and the expected exchange rate:

$$E = H(R - R^*, E^e) \tag{3}$$

The sign of the interest rate differential is expected to be negative, and the sign of the expected exchange rate is expected to be positive, suggesting that when the interest rate differential rises, the Costa Rica Colon would appreciate against the U.S. dollar.

The Monetary Models

Several versions of the monetary models include:

$$E = V(M - M^*, Y - Y^*, R - R^*) \tag{4}$$

$$E = V(M - M^*, Y - Y^*, \pi^e - \pi^{e^*}) \tag{5}$$

$$E = V(M - M^*, Y - Y^*, R - R^*, \pi^e - \pi^{e^*}) \tag{6}$$

where M, Y, π^e, M*, Y*, and π^{e^*} denote money supply in Costa Rica, real GDP in Costa Rica, the expected inflation rate in Costa Rica, money supply in the U.S., real GDP in the U.S., and the expected inflation rate in the U.S.

Equation (4) describes the Dornbusch model and the Bilson model. The sign of the relative interest rate is negative in the Dornbusch model and positive in the Bilson model. Equation (5) illustrates the Frenkel model. The sign of the expected inflation rate is positive. In the Frankel model in equation (6), the nominal exchange rate is expected to have a positive relationship with the relative money supply and the relative expected inflation rate and a negative relationship with the relative output and the relative interest rate.

The Mundell-Fleming Model

Extending Romer (2001), we can express the equilibrium in the goods market and the money market as:

$$Y = Z(Y, R - \pi^e, G, T, \varepsilon) \tag{7}$$

$$M / P = L(Y, R, R^*, \varepsilon) \tag{8}$$

where ε, G, T, L, and R^* are the real exchange rate, real government spending, real government taxes, the demand for money, and the world interest rate. Solving for Y and ε, we have the equilibrium real exchange rate as:

$$\bar{\varepsilon} = f(M / P, G, T, R, R^*, \pi^e) \tag{9}$$

The respective impacts of a change in real money supply, real government deficit spending, the domestic interest rate, and the world interest rate on the equilibrium real exchange rate can be written by:

$$\partial\bar{\varepsilon} / \partial(M / P) = -(1 - Z_Y) / |J| > 0, \tag{10}$$

$$\partial\bar{\varepsilon} / \partial(G - T) = (Z_G + Z_T) L_Y / |J| < 0, \tag{11}$$

$$\partial\bar{\varepsilon} / \partial R = [L_R(1 - Z_Y) + L_Y Z_R] / |J| > 0, \tag{12}$$

$$\partial\bar{\varepsilon} / \partial R^* = L_{R^*}(1 - Z_Y) / |J| < 0 \; if \; L_{R^*} > 0 \; or > 0 \; if \; L_{R^*} < 0, \tag{13}$$

where $|J|$ is the endogenous-variable Jacobian with a negative value, assuming that L_ε is positive. Thus, the equilibrium real exchange rate is expected to have a positive relationship with real money supply and the domestic interest rate and a negative relationship with real government deficit spending.

THE DATA

The data were collected from the *International Financial Statistics* published by the International Monetary fund. The nominal exchange rate is measured as Costa Rica Colon per U.S. dollar. In estimating the PPP model, the relative consumer price index (CPI) and the relative produce price index (PPI) are both considered. In estimating the UIP model, the deposit rates in Costa Rica and the U.S. are used to measure the interest rate differential because the money market rate or the Treasury bill rate for Costa Rica is not available. The lagged exchange rate is chosen to represent the expected exchange rate. In estimating the monetary models, M2 money, real GDP, the deposit rate, and the lagged inflation rate for both Costa Rica and the U.S. are used. In estimating the Mundell-Fleming model, the real exchange rate, real M2, the domestic deposit rate, the U.S. deposit rate, and the lagged inflation rate are used.

Government spending and tax revenues are not included due to lack of complete data. The consumer price index is used to derive real M2. Nominal M2 and real M2 are measured in billion colons for Costa Rica and billion dollars for the U.S. Real GDP is measured in million colons for Costa Rica and billion dollars for the U.S. The log scale is used except for variables with negative values.

Monthly data are used for the PPP and UIP models where quarterly data are used for the monetary and Mundell-Fleming models because the data for real GDP are available on a quarterly or yearly basis. The sample ranges from 1981.M1 to 2007.M9 for the PPP model, 1982.M1 to 2007.M8 for the UPI model, 2000.Q1 or 2000.Q2 to 2007.Q2 for the monetary models, and 2000.Q2 to 2007.Q2 for the Mundell-Fleming model. Different periods and data frequencies are used in order to increase the sample size. The monthly data have 321 observations for the PPP model and 308 observations for the UIP model. If quarterly data during 2000.Q2 – 2007.Q2 were used to test the PPP or the UIP, there would be only 29 observations in the sample.

EMPIRICAL RESULTS

Unit root tests in Table 1 show that all the variables are stationary in the first difference form. The cointegration test reveals that the variables in each of the four models are cointegrated and have a stable long-term relationship.

Table 1: Elliott-Rothenberg-Stock Unit Root Test

Variable	Test Statistic in Level Form	Test Statistic in First Difference Form
Log E	3.088	16.496
Log CPI/CPI*	1862.072	5.253
Log PPI/PPI*	447.058	2.715
Log R-Log R*	5.038	3.296
Log E*	12.682	13.522
Log M – Log M*	337.447	1.969
Log Y-Log Y*	597.498	70.911
$\pi^e - \pi^{e*}$	3.50	4.860
Log ε	11.708	3.760
Log M/P	225.781	2.156
*Log R**	5.144	4.073
π^e	4.174	11.554

Critical values: 1.87, 2.97, and 3.91 at the 1%, 5%, and 10% level respectively. This table shows the results of tests for the unit root for each of the variables. Values of the test statistic in the level and first difference forms are compared with the critical values at different significance levels.

Estimated regressions and related statistics are presented in Tables 2. Figures in the parenthesis are t-statistics. The Newey-West method is applied in empirical work to correct for both autocorrelation and heteroskedasticity when their forms are unknown. In the PPP model, both regressions have relatively high explanatory power, and the coefficient of the relative CPI or PPI is significant at the 1% level. The Wald test shows that the null hypothesis that the coefficient of the relative price measured either by the CPI or the PPI is equal to one cannot be rejected at the 5% level. The relative CPI seems to perform better in forecasting as the mean absolute percent error (MAPE) is calculated to be 3.927 compared with 4.589 when the relative PPI is used.

In the UIP model, 99.9% of the behavior of the exchange rate can be explained by the two right-hand side variables. Both of the coefficients are highly significant. The positive significant sign of the interest rate differential is opposite to the expected negative sign because a larger interest rate differential would cause the Costa Rica Colon to appreciate. The results may be due to a high degree of collinearity or the use of the lagged dependent variable as an explanatory variable. If the expected exchange rate is deleted from the regression, the coefficient of the interest rate differential is still positive and significant at the 1% level, and the value of R^2 declines to 17.6%. In the monetary models, the nominal exchange rate has a negative relationship with the relative money supply and the relative interest rate and is not affected by the relative real output and the relative inflation rate. The values of adjusted R^2 are relatively high. Empirical results suggest that the behavior of the exchange rate can be characterized by the Bilson model.

Table 2: Estimated Regressions for the Colon/USD Exchange Rate

Panel A: Purchasing Power Parity Model (Sample size = 321: 1981.M1-2007.M9)						
log E	**Intercept**	**log CPI/CPI***				
	5.725***	0.991***				Adj. R^2 = 0.995
	(789.651)	(73.669)				MAPE = 3.927
	Intercept	log PPI/PPI*				
	5.727***	1.015***				Adj. R^2 = 0.994
	(706.635)	(67.730)				MAPE = 4.589
Panel B: Uncovered Interest Parity Model (Sample size = 308: 1982M.1-2007.M8)						
log E	**Intercept**	**log R-log R***	**log E^e**			
	0.017***	0.004***	0.997***			Adj. R^2 = 0.999
	(3.863)	(3.017)	(1168.194)			MAPE = 5.427
Panel C: Monetary Models (Sample size = 30 for Version A: 2000.Q1-2007.Q2; and 29 for Versions B and C: 2000.Q2-2007.Q2)						
log E	**Intercept**	**log M – log M***	**log Y – log Y***	**log R – log R***	**π^e - π^{e*}**	
Eq. (4)	6.524***	0.579***	0.149	0.048***		Adj. R^2 = 0.988
	(136.143)	(17.846)	(0.760)	(6.601)		MAPE = 1.494
Eq. (5)	6.456***	0.554***	0.092		0.005	Adj. R^2 = 0.962
	(119.658)	(10.915)	(0.319)		(0.880)	MAPE = 2.549
Eq. (6)	6.511***	0.570***	0.181	0.047***	0.002	Adj. R^2 = 0.987
	(155.378)	(18.247)	(1.084)	(7.024)	(0.514)	MAPE = 1.509
Panel D: Mundell-Feming Model (Sample size = 29: 2000.Q2-2007.Q2)						
log ε	**Intercept**	**log M/P**	**log R**	**log R***	**π^e**	
	4.670***	0.115***	0.092***	-0.027***	0.005	Adj. R^2 = 0.622
	(29.755)	(6.734)	(5.005)	(-4.903)	(1.570)	MAPE = 0.994

This table shows the estimated regressions for the purchasing power parity model in equation (1), the uncovered interest parity model in equation (3), the monetary models in equations (4), (5) and (6), and the Mundell-Fleming model in equation (9). ***, ** and * indicate significance at the 1, 5 and 10 percent levels, respectively.
E = the nominal exchange rate (colon per U.S. dollar),
CPI = the consumer price index in Costa Rica,
CPI^* = the consumer price index in the U.S.,
PPI = the producer price index in Costa Rica, and
PPI^* = the producer price index in the U.S.
R = the interest rate in Costa Rica,
R^* = the interest rate in the U.S., and
E^e = the expected exchange rate.
log E = log of the nominal exchange rate (colon per U.S. dollar),
log M – log M^* = log of nominal money supply in Costa Rica – log of nominal money supply in the U.S.,
log R – log R^* = log of the interest rate in Costa Rica – log of the interest rate in the U.S.,
log Y – log Y^* = log of real GDP in Costa Rica – log of real GDP in the U.S., and
ε= the real exchange rate,
M/P = real money supply in Costa Rica,
π^e = the expected inflation rate in Costa Rica.
π^e - π^{e*} = the expected inflation rate in Costa Rica – the expected inflation rate in the U.S.

In the Mundell-Fleming model, the value of adjusted R^2 is 62.2%. The real exchange rate has a positive relationship with real M2 and the domestic interest rate and a negative relationship with the world interest rate. These suggest that more real money supply or a higher domestic interest rate would cause the Costa Rica Colon to depreciate and that a higher world interest rate would cause the Costa Rica Colon to appreciate. Table 3 reestimates the regressions based on a common sample period of 2000.Q2-2007.Q2 with a total of 29 observations. Although the sample size is much smaller for the PPP model and the UIP model, the MAPE improves in these two models. The values of adjusted R^2 are relatively high. The estimated slope coefficients of the PPP model are slightly larger than those in Table 1. The estimated coefficient of the variable log R – log R^* in the UIP model is also larger than that in Table 1. Estimated results in Panels C and D are either very similar or identical because of the use of the same or similar sample size.

Table 3: Estimated Regressions for the Colon/USD Exchange Rate Based on the Same Sample Period of 2002.Q2-2007.Q2

Panel A: Purchasing Power Parity Model (Sample size = 29)

log E	**Intercept**	**log CPI/CPI***				
	5.735***	1.056***				Adj. R^2 = 0.988
	(550.037)	(29.862)				MAPE = 1.499
	Intercept	log PPI/PPI*				
	5.720***	1.137***				Adj. R^2 = 0.970
	(353.753)	(19.179)				MAPE = 2.362

Panel B: Uncovered Interest Parity Model (Sample size = 29)

log E	**Intercept**	**log R-log R***	**log E^e**			
	0.036	0.009***	0.995***			Adj. R^2 = 0.999
	(1.090)	(5.615)	(172.065)			MAPE = 0.638

Panel C: Monetary Models (Sample size = 29)

log E	**Intercept**	**log M – log M^***	**log Y – log Y^***	**log R – log R^***	**π^e - π^{e*}**	
Eq. (4)	6.515***	0.570***	0.186	0.047***		Adj. R^2 = 0.987
	(124.357)	(15.406)	(0.878)	(6.290)		MAPE = 1.520
Eq. (5)	6.456***	0.554***	0.092		0.005	Adj. R^2 = 0.962
	(119.658)	(10.915)	(0.319)		(0.880)	MAPE = 2.549
Eq. (6)	6.511***	0.570***	0.181	0.047***	0.002	Adj. R^2 = 0.987
	(155.378)	(18.247)	(1.084)	(7.024)	(0.514)	MAPE = 1.509

Panel D: Mundell-Feming Model (Sample size = 29)

log ε	**Intercept**	**log M/P**	**log R**	**log R^***	**π^e**	
	4.670***	0.115***	0.092***	-0.027***	0.005	Adj. R^2 = 0.622
	(29.755)	(6.734)	(5.005)	(-4.903)	(1.570)	MAPE = 0.994

This table shows the estimated regressions for the purchasing power parity model in equation (1), the uncovered interest parity model in equation (3), the monetary models in equations (4), (5) and (6), and the Mundell-Fleming model in equation (9). ***, ** and * indicate significance at the 1, 5 and 10 percent levels, respectively.
E = the nominal exchange rate (colon per U.S. dollar),
CPI = the consumer price index in Costa Rica,
CPI^* = the consumer price index in the U.S.,
PPI = the producer price index in Costa Rica,
PPI^* = the producer price index in the U.S.
R = the interest rate in Costa Rica,
R^* = the interest rate in the U.S., and
E^e = the expected exchange rate.
log E = log of the nominal exchange rate (colon per U.S. dollar),
log M – log M^* = log of nominal money supply in Costa Rica – log of nominal money supply in the U.S.,
log R – log R^* = log of the interest rate in Costa Rica – log of the interest rate in the U.S.,
log Y – log Y^* = log of real GDP in Costa Rica – log of real GDP in the U.S., and
ε= the real exchange rate,
M/P = real money supply in Costa Rica,
π^e = the expected inflation rate in Costa Rica and π^e - π^{e*} = the expected inflation rate in Costa Rica – the expected inflation rate in the U.S.

SUMMARY AND CONCLUSIONS

This paper has examined the behavior of the Costa Rica Colon exchange rate against the U.S. dollar. Four different models are considered in empirical work. The coefficient of the interest rate differential has a wrong sign in the uncovered interest parity model. Higher relative prices, higher interest rate differentials, and more money supply are expected to cause a weaker Colon against the U.S. dollar. Excluding the UIP model, the PPP model and monetary models have higher explanatory power than the Mundell-Fleming model. However, the Mundell-Fleming model performs the best in forecasting, followed by the Bilson model, the Frankel model, the Frenkel model, the PPP model with the relative CPI, and the PPP model with the relative PPI.

There may be areas for future research. The unexpected positive sign of the interest rate differential in the UIP model may suggest that more work needs to be done in the study of exchange rate movements for Costa Rica. The expected exchange rate plays a significant role in the determination of the exchange rate and may need to be constructed with more advanced methodologies.

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BIOGRAPHY

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MARKET STATES AND THE PROFITABILITY OF MOMENTUM STRATEGIES: EVIDENCE FROM THE TAIWAN STOCK EXCHANGE

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ABSTRACT

This paper examines the impact of market states on the profitability of momentum strategies using weekly data from the Taiwan Stock exchange over the 10-year period 1997-2006. Market states refer to the states of market such as up or down markets. In this paper, the formation period is defined as in an up (down) state if the market return over the six-month period prior to the holding period is nonnegative (negative). The results indicate that market states in the formation period are positively associated with the profitability of the momentum strategies. The results are consistent with the overreaction theory developed in Daniel et al. (2004). Moreover, the empirical results indicate that market states in the holding period are negatively associated with the profitability of the momentum strategies. The holding period is defined as in an up (down) state if the market return in the six-month period following the formation period is nonnegative (negative). The momentum profits appear to be higher in a bearish holding period and lower for a bullish holding period. Thus, the market states in the holding period also provide information regarding the profitability of the momentum strategies.

JEL: G11, G14

INTRODUCTION

Previous research has documented the short-term cross-sectional momentum in stock returns (i.e., Jagadeesh and Titman, 1993) and long-term cross-sectional reversals in stock returns (i.e., De Bondt and Thaler, 1985, 1987). Several behavioral theories have been developed to explain the anomalous price behavior. Daniel, Hisrhleifer and Subrahmanyam (1998), for example, propose a theoretic model in that investors' overconfidence about their private information leads to the short-term price momentum and the long-term price reversal. Specifically, investors tend to react differently to new information due to a self-attribution bias. If the upcoming information is consistent with these investors' prior belief, they tend to attribute the confirming news to their own skill. In contrast, if the upcoming information is inconsistent with their prior belief, they tend to attribute such disconfirming information to external noise. The self-attribution bias reinforces the overconfidence following the arrival of the confirming news, which generates the pattern of short-term price momentum. In the long-run, however, the arrival of information regarding the fair firm value leads to a price correction and thus long-term price reversals.

Cooper, Gutierrez and Hameed (2004) examine the overreaction theories by examining the impact of market states on momentum profits. In their research, the stock market is defined as in an up (down) market if the market return in the portfolio formation period is nonnegative (negative). Specifically, they test the theory of Daniel et al. (1998) which predicts greater aggregate overconfidence following market gains. Since investors in aggregate hold a long position in equity securities, their overconfidence tend to be greater following market gains due to the reinforcement of the self-attribution bias. If so, short-term price momentum should be greater following up markets. Using CRSP monthly data from 1929-1995, Cooper et al. (2004) examine the profitability of momentum strategies that take a long position in the prior winner portfolio and a short position in the prior loser portfolio. Moreover, they examine whether the state of markets affects the profitability of momentum strategies. The state of markets is divided into up and down markets based on market returns in the portfolio formation period. They find that the average monthly momentum profit following up markets is significantly positive at 0.93%. In contrast, the average monthly momentum profit following down markets is negative at -0.37%. Moreover, the up-market momentum profit reverses in the long-run. The results of asymmetrical momentum profits in the up and down markets are consistent with the prediction of the overreaction model in Daniel et al. (1998). While the empirical analysis in Cooper et al. (2004) examines the impact of formation-period market states on momentum profits, the overreaction theory in Daniel et al. (1998) can extended to examine the impact of market states in both the formation and the holding period.

Following Cooper et al. (2004), we examine the impact of formation-period market states on the profitability of momentum strategies using data from the Taiwan Stock Exchange over the 10-year period 1997-2006. The formation period is defined as in an up (down) state if the market return over the six-month period *prior to* the holding period is nonnegative (negative). The market return is based on the value-weighted market index compiled by the Taiwan Stock Exchange. Aside from the formation-period market states, we also examine the impact of holding-period market states on the profitability of the momentum strategies. The holding period is defined as in an up (down) state if the market return in the six-month period *following* the formation period is nonnegative (negative).

Consistent with the finding in Cooper et al. (2004), our results indicate that the formation-period market states affect the momentum profits. Momentum strategies generate significantly positive returns following market gains in the formation periods. Moreover, our empirical result indicates that the holding-period market states also affect momentum. Momentum strategies perform better for holding periods in down markets as opposed to up markets. The plan of this paper is as follow. Section 2 provides a brief literature review. Section 3 introduces the institutional background of the Taiwan Stock Exchange. Section 4 describes the data and methodology employed in this paper. Section 5 presents empirical results and Section 6 concludes.

LITERATURE REVIEW

De Bondt and Thaler (1985) note that, based on research in experimental psychology (i.e., Kahneman and Tversky, 1982), most people tend to overreact to unexpected and dramatic news events. If investors overreact to unexpected good news and bad news in the prior period and correct their overreaction in the subsequent period, we would expect price reversals in the successive periods. De Bondt and Thaler (1985) test the overreaction hypothesis by examining the stock price behavior for stocks contained in the CRSP monthly data file over the period 1926 through 1982. Their results indicate that the return data are consistent with the prediction of the overreaction hypothesis. For stocks ranked in the past 3-year formation period, losers outperform winners in the subsequent 3-year holding period. The average cumulative abnormal return in the holding period is significantly positive at 19.6% for the losers, but only -5% for the winners. Thus, the contrarian strategies of buying prior losers and selling prior winners would yield abnormal returns over a 3-year holding period.

Similarly, other research provides evidence of shorter-term price reversals (i.e., Jagadeesh, 1990; Lehmann, 1990). These papers indicate that contrarian strategies based on returns in prior weeks as well as prior months generate significant abnormal returns. However, since shorter-term strategies involve intensive transactions, abnormal returns generated from these strategies may be sensitive microstructure issues such as bid-ask spreads. Lo and MacKinlay (1990) point out that a large part of the short-term contrarian profits can be attributed to a delayed price reaction to common factors rather than to overreaction.

Although the profitability of contrarian strategies has attracted much attention in academic literature, early literature on trading strategies focused on momentum strategies that buy prior winners and sell prior losers. Levy (1967), for example, documents empirical evidence that momentum strategies based on the relative strength of stocks in the past 27 weeks produce significant abnormal returns. Similarly, Jegadeesh and Titman (1993) report significant abnormal returns in the holding period of 3-12 months for momentum strategies that buy prior winners and sell prior losers for stocks ranked in the formation period of 3-12 months. They note that the momentum profits cannot be attributed to systematic risk or to delayed stock price reactions to common factors.

Several explanations have been offered regarding the abnormal returns of the momentum strategies. Jegadeesh and Titman (1995) suggest that investors' underreaction to firm-specific information is a major cause of momentum abnormal returns. Conrad and Kaul (1998) indicate that momentum profits result from cross-sectional differences in expected returns. Jegadeesh and Titman (2001) suggest that momentum profits are due to delayed overreactions that eventually reverse. Bhojraj and Swaminathan (2006) propose that investors initially under-react to information which results in undervalued stock prices. Then, stock prices move to their fundamental values gradually so that price continuation exists in the short run. The stock prices may rise even above and far from their fundamental values. Finally, the stock prices may revert to their fundamental values eventually in the long-run.

Moreover, recent literature provides empirical evidence that the state of market is an important factor that may affect momentum profits (see, for example, Griffin, Ji and Martin, 2003; Cooper et al., 2004; Antonios and Patricia, 2006; Huang, 2006). Griffin et al. (2003) indicate that momentum profits are pronounced in both good and bad economic states, and this phenomenon reverses in the holding periods of 1- to 5- years. Cooper et al. (2004) classify market returns into up and down states based on lagged three-year market returns. Their empirical results indicate that short-run momentum profits exist following up markets and that the mean monthly profit in the up markets is higher than that in the down markets. This phenomenon is robust when one- or two-year market returns are used to classified market states. Finally, they find significant prices reversals following both up and down markets in the long run. They consider that the short-term momentum and long-term price reversals are consistent with the overreaction hypothesis.

Antonios and Patricia (2006) examine the profitability of momentum strategies following bull and bear markets utilizing data from the London Stock Exchange. They define bull and bear markets based on market returns over different periods. Their empirical findings indicate that momentum profits are more pronounced following bear markets. In addition, the longer the bear market periods, the more pronounced the momentum returns. Moreover, they find that momentum profits become negative following stronger bull markets. One possible explanation for the momentum profits is that investors who realize their losses (gains) in the past tend to underreact (overreact) to present information. Their findings are consistent with the behavioral model proposed by Daniel et al. (1998) and Hong and Stein (1999).

Institutional Background

Established in 1962, the Taiwan Stock Exchange is the major stock market in Taiwan. The number of listed stocks varies in the sample period 1997 through 2006 ranging from 404 in 1997 to 688 in 2006. Individual investors contribute a large share in trading volume although institutional investors play an increasing important role in recent years. In particular, trading volume contributed by individual investors declines from 90.7% in 1997 to 72.8% in 2006 with the remaining trading volume coming from institutional investors. Moreover, trading activity is heavy with turnover ratios ranging from 407.32% in 1997 to 142.2% in 2006. As such, investors appear to hold their stocks in relatively shorter periods than investors in other more mature markets (i.e., Securities and Futures Bureau, Taiwan, R.O.C., 2008).

The Taiwan Stock Exchange is an order-driven market without the aid of market makers. Investors submit buy and sell orders to their brokers; these orders are then matched by the computer system in the stock exchange. The matching process is based on the price and time priority. Thus, all buy orders with bidding prices above the transaction prices are filled with priority. Similarly, all sell orders with asking prices below the transaction prices are filled with priority.

The Taiwan Stock Exchange opens from 9:00 a.m. to 1:30 p.m. Mondays through Fridays. During the pre-trade period 8:30-9:00 a.m., investors submit their orders which are accumulated to determine the opening price at 9:00 a.m. through a call auction method. Following opening, the transaction prices are determined through the same call method every 30 to 45 seconds. Finally, the closing price is determined by the same call method for orders accumulated over the last 5-minute interval 1:25 to 1:30 p.m. preceding the closing of the trading session

Daily price limits and a sliding schedule of tick sizes are utilized in the Taiwan Stock Exchange. In an attempt to control excessive price volatility, the Taiwan Stock Exchange imposes a daily price limit of 7%. The prevailing tick sizes or minimum price variations are NT$0.01, 0.05, 0.10, 0.50, 1.00, 5.00 respectively for trading prices in the range of NT$0-10, 10-50, 50-100, 100-500, 500-1000, and 1000-above respectively.

The major transaction costs involve commission fees and transaction taxes. For a buy transaction, investors pay a commission fee of 0.1425% to brokers and a transaction tax of 0.3% to the government. For a sell transaction, investors pay only the brokerage fee of 0.1425% although brokers may provide discounts to investors in order to promote business. For a typical round-trip transaction without discounts, the transaction cost involve a 0.585% ((0.1425%)×(2)+0.3%=0.585%) of the trading value.

SAMPLE AND METHODOLOGY

The sample involves all stocks listed on the Taiwan Stock Exchange over the 10-year period 1997 to 2006. Weekly return data are obtained from the Taiwan Economic Journal Database. Sample firms suffering from financial distress are excluded. The screening process yields a total of 294 to 597 sample firms in the sample period.

To examine the profitability of momentum strategies, sample firms are first sorted into quintiles based on market-adjusted abnormal returns in the *formation period*. The quintile with the highest average abnormal return is the *winner* portfolio while the quintile with the lowest average abnormal return is the *loser* portfolio. The stocks in the winner or loser portfolio are equally weighted to yield the portfolio returns. The momentum strategies involve a long position in the winner portfolio and a short position in the loser portfolio. The profitability of the momentum strategies is assessed by the abnormal return of the momentum strategies in the subsequent *holding period.*

The formation periods involve five time intervals ranging from 1, 2, 4, 12 to 26 weeks while the subsequent holding periods involve six time intervals ranging from 1, 2, 4, 12, 26 to 52 weeks. Thus, the profitability for a total of 5 × 6 =30 momentum strategies is examined. The time windows for the formation and the holding periods roll over the whole sample period. The momentum strategy with the formation periods of 1 week and the holding period of 1 week, or strategy (1, 1), starts from the formation period of January 4, 1997 to January 11, 1997 and the corresponding holding period from January 11, 1997 to January 18, 1997. This process continues until the end of the sample period.

The overreaction theory in Daniel et al. (1998) predicts greater short-term momentum profits following market gains. To examine this theory, we follow Cooper et al. (2004) by identifying the market states in the holding period into either in an up or down market. The formation period is considered in an up (down) state if the market return over the six-month period *prior to* the holding period is nonnegative (negative). The value-weighted market index compiled by the Taiwan Stock Exchange is used to derive the market return. To examine the impact of market states on the profitability of momentum strategies, momentum profits following an up-market formation period are compared to those following a down-market formation period. According to the overreaction theory, we would expect the momentum profits following an up market to perform better than those following a down market.

For each momentum strategy (J, K) with a formation period J and a holding period K, the momentum profit in an up or down market is estimated by averaging abnormal returns over holding periods across securities and market states. First, market-adjusted abnormal returns for each stock i in the holding period is estimated. The market-adjusted abnormal returns for firm i in holding period t following a market state c, $AR_{i,t,c}$, is estimated as the return on stock i minus the corresponding market return as follow (the subscripts (j, k) for the associated strategy (J, K) are omitted in $AR_{i,t,c}$ for brevity):

$$AR_{i,t,c} = (R_{i,t,c} - R_{m,t,c}) \tag{1}$$

The stock returns, Ri,t,c, and market returns, $R_{m,t,c}$, are evaluated as the price relatives: $R_{i,t,c} = (P_{i,t,c} - P_{i,t-1,c})/P_{i,t-1,c}$, $R_{m,t,c} = (P_{m,t,c} - P_{m,t-1,c})/P_{m,t-1,c}$, where $R_{i,t,c}$ denotes the return on stock i in either an up or down state, $R_{m,t,c}$ denotes the market return derived from the Taiwan Weighted Stock Index, and [t-1, t] is the holding period over which returns are estimated. The average abnormal returns (AARs) for each strategy (J, K) in the up or down states are evaluated by averaging the abnormal returns $AR_{i,t,c}$ across the total sample observations for the winner and the loser portfolios respectively as follows:

$$AAR^{L}_{j,k,c} = \sum_{s=1}^{OBS} AR^{L}_{j,k,c,s} / OBS^{L}_{j,k,c} \tag{2a}$$

$$AAR^{W}_{j,k,c} = \sum_{s=1}^{OBS} AR^{W}_{j,k,c,s} / OBS^{W}_{j,k,c} \tag{2b}$$

where $OBS^{L}_{j,k,c}$ and $OBS^{W}_{j,k,c}$ are the total numbers of return observations for the momentum strategy (J, K) following market state c, whereas the superscripts L and W denote the loser portfolio and the winner portfolio respectively and c denotes the market state as either an up or a down state in the formation periods.

Since the momentum strategy consist of taking a long position in the winner portfolio and a short position in the loser portfolio. The average abnormal returns of the strategy (J, K) in a market state c are evaluated as the differences between the average abnormal returns on the winner quintiles and those on loser quintiles in each state of the market, or $AAR_{j,k,c} = (AAR^{W}_{j,k,c} - AAR^{L}_{j,k,c})$.

Aside from the consideration of market states in the formation periods, we also examine how momentum strategies perform when the holding periods turn out to be in either an up or down market. The state of the market for the holding period is determined similar to that for the formation periods. Specifically, a holding period is in an up (down) state if the market return in the six-month period *following* the formation period is nonnegative (negative). While previous research documents the impact of formation-period market states on momentum profits, a further analysis on how market states in holding periods affect momentum profits should enhance our understanding of stock price behavior.

EMPIRICAL RESULTS

Market States in the Formation Periods and Momentum Profits

Table 1 reports momentum profits following up-market formation periods. The results indicate that momentum strategies generate significantly positive abnormal returns for 24 out of the 30 momentum strategies. The average abnormal return across the 30 momentum strategies is 2.17% with winner and loser portfolios earning average abnormal returns at 3.05% and 0.88% respectively. Thus, the results indicate price continuation for the winner portfolios and slight price reversals for the loser portfolios.

Moreover, Table 1 indicates that the momentum profits become larger as formation periods increase up to 12 weeks. For example, with holding periods of one year (K = 52 weeks), momentum profits (W-L) with long positions in winners and short positions in losers increase from 3.81% for the formation period of one week (J = 1 week) to 9.80% for formation period of 3 months (J = 12 weeks). Similarly, the momentum profits become larger as holding periods become longer. For formation period of 4 weeks (J = 4 weeks), for example, the momentum profits increase from 0.38% for holding period of one week (K = 1 week) to 6.49% for holding period of 52 weeks (K = 52 weeks).

Table 1: Average Abnormal Returns for Strategies (J, K) Following Up-Market Formation Periods

Formation Period (J)	Holding Period (K)	Following Up-market Formation Periods					
		Winner	t_W	Loser	t_L	W-L	t_{W-L}
1	1	0.0023	4.93 *	0.0022	5.52 *	0.0002	0.25
	2	0.0056	8.53 *	0.0021	3.90 *	0.0035	4.07 *
	4	0.0093	9.95 *	0.0020	2.58 *	0.0073	5.93 *
	12	0.0168	10.65 *	0.0093	6.43 *	0.0076	3.54 *
	26	0.0377	15.95 *	0.0180	8.35 *	0.0197	6.15 *
	52	0.0698	18.19 *	0.0317	8.80 *	0.0381	7.25 *
2	1	0.0028	4.29 *	0.0012	2.10 *	0.0017	1.94
	2	0.0076	8.44 *	0.0007	0.91	0.0069	5.93 *
	4	0.0102	8.01 *	-0.0013	-1.24	0.0115	7.00 *
	12	0.0211	9.52 *	0.0062	3.27 *	0.0149	5.12 *
	26	0.0406	12.43 *	0.0073	2.50	0.0333	7.60 *
	52	0.0734	13.86 *	0.0199	4.08 *	0.0535	7.43 *
4	1	0.0035	3.54 *	-0.0003	-0.38	0.0038	3.05 *
	2	0.0092	6.89 *	0.0013	1.32	0.0079	4.71 *
	4	0.0126	6.59 *	0.0024	1.66	0.0102	4.20 *
	12	0.0262	8.10 *	0.0072	2.68 *	0.0191	4.54 *
	26	0.0498	10.45 *	0.0100	2.37 *	0.0398	6.25 *
	52	0.0853	11.32 *	0.0204	2.96 *	0.0649	6.35 *
12	1	0.0049	2.80 *	-0.0024	-1.75	0.0073	3.28 *
	2	0.0100	4.63 *	0.0028	1.50	0.0072	2.54 *
	4	0.0146	4.56 *	-0.0040	-1.50	0.0185	4.47 *
	12	0.0313	5.33 *	0.0004	0.09	0.0309	4.14 *
	26	0.0520	6.66 *	0.0058	0.73	0.0461	4.12 *
	52	0.0779	5.86 *	-0.0201	-1.65	0.0980	5.42 *
26	1	0.0052	2.56 *	0.0110	5.07 *	-0.0058	-1.93
	2	0.0014	0.48	0.0157	5.53 *	-0.0143	-3.50 *
	4	-0.0034	-0.82	0.0082	2.15 *	-0.0116	-2.06
	12	0.0370	5.09 *	0.0087	1.43	0.0284	3.00 *
	26	0.0746	6.72 *	0.0090	0.77	0.0656	4.08 *
	52	0.1253	6.56 *	0.0883	4.11 *	0.0370	1.29
Mean		0.0305		0.0088		0.0217	

*An asterisk, *, indicates significance at 5% of one-tail test.*

Table 2 reports momentum profits following down-market formation periods. The results indicate that momentum strategies generate significantly negative abnormal returns for 17 out of the 30 momentum strategies. In contrast, only 4 strategies generate significantly positive abnormal returns. The average abnormal return across the 30 momentum strategies is -1.18% with the winner portfolios earning average abnormal returns of 0.29% and the loser portfolios 1.47%. Thus, the results indicate stronger price reversals for the loser portfolios but only minor price continuation for the winner portfolios following down-market formation periods.

Moreover, Table 2 indicates that the momentum profits become more negative as formation periods increase up to 4 weeks. For example, with the formation periods of 4 weeks (J = 4 weeks), momentum profits are all significantly negative ranging from -0.34% with the holding period of one week (K = 1 week) to -5.52% for the holding period of 26 weeks (K = 26 weeks).

Thus, the results in Tables 1 and 2 are consistent with the prediction of the overreaction theory in Daniel et al. (1998) in that the state of markets provides additional information regarding the profitability of momentum strategies. Following market gains in the formation period, investors tend to be overconfident. Moreover, investors may attribute trading gains to their own selection skill more than they should. Thus, the overconfidence appears to be stronger for the winner portfolio than for the loser portfolio. The short-run price continuation for the winner portfolio may reflect this self-attribution bias. In contrast, following market losses in the formation period, investors tend to attribute trading losses to external noise more than they should. Price reversals are more evident especially for the loser portfolio.

<u>Market States in the Holding Periods and Momentum Profits</u>

Tables 3 and 4 report momentum profits classified by market states in the holding periods. Table 3 report momentum profits following market gains in the formation period. As expected, Table 3 indicates that the momentum profits are generally positive following market gains. However, momentum profits are higher for holding periods in a down market as opposed to holding periods in an up market. The average abnormal return for the 30 momentum strategies is 2.81% for the down-market holding period as opposed to the 1.33% for the up-market holding market. Moreover, for the holding periods in down markets, 25 momentum strategies experience significantly positive abnormal returns. In contrast, for the holding periods in a down market, only 14 momentum strategies earn significantly positive abnormal returns.

The higher momentum profits for holding periods in down markets appear to be driven by the different reaction between the winner and the loser portfolio in the holding period. Specifically, loser portfolios appear to be more sensitive to market states in the holding period than do winner portfolios. That is, the loser portfolio performs better than the market in the up-market holding period, but worse than the market in the down-market holding period. For the loser portfolio, the average abnormal return across the 30 strategies is 2.16% in an up-market holding period but only -0.03% in a down-market holding period. The higher sensitivity of the loser portfolio to holding-period market states is more evident for longer holding periods. In contrast, winner portfolios are less sensitive to market states in the holding period. For the winner portfolio, the average abnormal return across the 30 strategies is 3.49% in an up-market holding period and still 2.78% in a down-market holding period. The higher market sensitivity of the loser portfolio results in higher momentum profits in the down state of the holding period.

Table 2: Average Abnormal Returns for Strategies (J, K) Following Down-Market Formation Periods

Formation Period (J)	Holding Period (K)	Following Down-market Formation Periods					
		Winner	t_W	Loser	t_L	W-L	t_{W-L}
1	1	-0.0020	-4.16 *	-0.0002	-0.43	-0.0018	-2.50 *
	2	0.0000	0.05	-0.0017	-2.42 *	0.0018	1.78
	4	0.0009	0.91	0.0001	0.13	0.0008	0.55
	12	0.0056	2.92 *	0.0139	7.02 *	-0.0083	-3.00 *
	26	0.0061	2.28 *	0.0338	11.51 *	-0.0277	-6.94 *
	52	0.0143	3.67 *	0.0364	8.69 *	-0.0221	-3.87 *
2	1	-0.0011	-1.51	-0.0015	-2.01 *	0.0004	0.42
	2	0.0003	0.25	-0.0024	-2.23 *	0.0026	1.78
	4	0.0024	1.57	0.0011	0.75	0.0013	0.62
	12	0.0080	2.77 *	0.0119	4.05 *	-0.0039	-0.95
	26	0.0059	1.46	0.0407	9.20 *	-0.0347	-5.77 *
	52	0.0085	1.50	0.0355	5.59 *	-0.0269	-3.16 *
4	1	-0.0045	-4.26 *	-0.0011	-0.99	-0.0034	-2.19 *
	2	-0.0064	-4.40 *	0.0001	0.04	-0.0065	-3.13 *
	4	-0.0057	-2.61 *	0.0117	5.35 *	-0.0173	-5.63 *
	12	0.0017	0.41	0.0217	5.28 *	-0.0200	-3.43 *
	26	-0.0137	-2.57 *	0.0415	6.45 *	-0.0552	-6.61 *
	52	-0.0058	-0.74	0.0395	4.51 *	-0.0453	-3.86 *
12	1	0.0005	0.30	-0.0056	-2.73 *	0.0061	2.28 *
	2	0.0021	0.96	-0.0004	-0.14	0.0025	0.74
	4	-0.0165	-5.09 *	0.0139	3.69 *	-0.0304	-6.11 *
	12	-0.0213	-3.90 *	0.0240	3.68 *	-0.0453	-5.32 *
	26	-0.0135	-1.54	0.0370	3.87 *	-0.0505	-3.89 *
	52	0.0292	2.07 *	0.0111	0.85	0.0181	0.95
26	1	0.0103	3.76 *	-0.0068	-2.58 *	0.0170	4.50 *
	2	0.0146	3.94 *	-0.0070	-1.86	0.0217	4.08 *
	4	0.0011	0.21	0.0252	4.63 *	-0.0241	-3.14 *
	12	0.0066	0.80	0.0892	7.56 *	-0.0826	-5.74 *
	26	0.0150	1.04	-0.0055	-0.40	0.0205	1.03
	52	0.0441	2.14 *	-0.0152	-0.72	0.0592	2.02 *
		0.0029		0.0147		-0.0118	

*An asterisk, *, indicates significance at 5% of one-tail test.*

Table 3: Average Abnormal Returns for Strategies (J, K) Following Up-Market Formation Periods

Formation Period (J)	Holding Period (K)	Holding Periods in Up Markets				Holding Periods in Down Markets			
		Winner	Loser	W-L	t_{W-L}	Winner	Loser	W-L	t_{W-L}
1	1	0.0004	0.0016	-0.0012	-1.55	0.0043	0.0027	0.0016	1.72
	2	0.0014	0.0000	0.0014	1.22	0.0101	0.0044	0.0057	4.35 *
	4	0.0049	-0.0022	0.0071	4.42 *	0.0139	0.0065	0.0074	3.99 *
	12	0.0055	0.0016	0.0039	1.34	0.0292	0.0176	0.0116	3.67 *
	26	0.0417	0.0253	0.0164	3.33 *	0.0335	0.0103	0.0232	5.77 *
	52	0.0788	0.0535	0.0254	3.07 *	0.0602	0.0085	0.0517	8.11 *
2	1	0.0022	0.0001	0.0021	1.88	0.0035	0.0022	0.0012	0.94
	2	0.0030	-0.0024	0.0054	3.57 *	0.0126	0.0039	0.0086	4.77 *
	4	0.0054	-0.0052	0.0107	4.87 *	0.0153	0.0030	0.0124	5.04 *
	12	0.0107	-0.0003	0.0109	2.81 *	0.0324	0.0131	0.0192	1.74
	26	0.0482	0.0077	0.0405	6.15 *	0.0327	0.0069	0.0258	4.50 *
	52	0.0920	0.0443	0.0477	4.22 *	0.0536	-0.0060	0.0596	6.85 *
4	1	-0.0003	-0.0013	0.0011	0.65	0.0068	0.0006	0.0061	3.38 *
	2	0.0020	-0.0002	0.0022	0.96	0.0157	0.0027	0.0130	5.35 *
	4	0.0068	-0.0006	0.0073	2.18 *	0.0176	0.0050	0.0126	3.67 *
	12	0.0125	0.0037	0.0088	1.53	0.0385	0.0103	0.0282	4.67 *
	26	0.0569	0.0093	0.0476	4.73 *	0.0436	0.0106	0.0329	4.09 *
	52	0.0950	0.0377	0.0573	3.41 *	0.0774	0.0064	0.0710	5.69 *
12	1	-0.0088	-0.0075	-0.0013	-0.34	0.0114	0.0000	0.0114	4.16 *
	2	-0.0026	-0.0037	0.0012	0.25	0.0160	0.0059	0.0101	2.83 *
	4	0.0037	-0.0045	0.0082	1.17	0.0198	-0.0037	0.0234	4.57 *
	12	0.0278	-0.0003	0.0281	2.27 *	0.0330	0.0008	0.0323	3.46 *
	26	0.0910	0.0309	0.0601	2.86 *	0.0334	-0.0061	0.0395	3.01 *
	52	0.1252	0.0939	0.0313	0.76	0.0562	-0.0723	0.1285	7.04 *
26	1	0.0024	0.0212	-0.0188	-4.33 *	0.0081	0.0006	0.0075	1.87
	2	-0.0055	0.0253	-0.0308	-4.93 *	0.0084	0.0058	0.0026	0.50
	4	-0.0080	0.0293	-0.0373	-4.39 *	0.0013	-0.0132	0.0145	2.02 *
	12	0.0290	0.0207	0.0083	0.71	0.0453	-0.0036	0.0488	3.26 *
	26	0.1014	0.0131	0.0883	4.02 *	0.0474	0.0048	0.0426	1.99 *
	52	0.2232	0.2574	-0.0342	-0.67	0.0529	-0.0369	0.0897	2.88 *
		0.0349	0.0216	0.0133		0.0278	-0.0003	0.0281	

*An asterisk, *, indicates significance at 5% of one-tail test.*

Table 4 reports momentum profits following market losses in the formation period. Table 4 indicates that most momentum strategies yield negative abnormal returns following market losses. Moreover, the momentum profits appear to be higher in a down-market holding period as opposed to those in an up-market holding period. Of the 30 momentum strategies, 25 strategies perform better in down markets as opposed to in up markets. The average abnormal return for the 30 momentum strategies is 0.42% in the down-market holding period as opposed to the -1.79% in the up-market holding market. Moreover, only 9 momentum strategies experience significantly negative abnormal returns in a down market. In contrast, 23 momentum strategies experience significantly negative abnormal returns in an up market.

Again, the poor performance of the momentum strategies following market losses appears to be driven by different market sensitivity between winner and loser portfolios. That is, the loser portfolios are more sensitive to market states than the winner portfolios. The average abnormal returns across the 30 strategies indicate that the loser portfolio perform better than the winner portfolio in the up market, but worse than the winner portfolio in the down market. For up markets in the holding period, the average abnormal return across the 30 strategies is 1.18% for the loser portfolio, which is higher than the -0.61% for the winner portfolio. In contrast, for down markets in the holding period, the average abnormal return across the 30 strategies is 1.61% for the loser portfolio, which is lower than the 2.03% for the winner portfolio.

Table 4: Average Abnormal Returns for Strategies (J, K) Following Down-Market Formation Periods

Formation Period (J)	Holding Period (K)	Holding Periods in Up Markets				Holding Periods in Down Markets			
		Winner	Loser	W-L	t_{W-L}	Winner	Loser	W-L	t_{W-L}
1	1	-0.0050	-0.0022	-0.0028	-3.25 *	0.0030	0.0032	-0.0001	-0.09
	2	-0.0052	-0.0021	-0.0031	-2.55 *	0.0090	-0.0011	0.0101	5.87 *
	4	-0.0068	-0.0030	-0.0038	-2.21 *	0.0143	0.0055	0.0088	3.56 *
	12	-0.0033	0.0125	-0.0159	-4.32 *	0.0205	0.0163	0.0042	1.03
	26	-0.0032	0.0305	-0.0337	-6.03 *	0.0221	0.0395	-0.0173	-3.44 *
	52	0.0190	0.0495	-0.0306	-3.88 *	0.0061	0.0135	-0.0073	-0.97
2	1	-0.0055	0.0000	-0.0055	-4.27 *	0.0066	-0.0042	0.0108	5.92 *
	2	-0.0072	-0.0017	-0.0055	-3.00 *	0.0134	-0.0036	0.0170	6.74 *
	4	-0.0069	-0.0011	-0.0057	-2.24 *	0.0190	0.0050	0.0140	3.86 *
	12	-0.0010	0.0141	-0.0151	-2.72 *	0.0232	0.0082	0.0150	2.55 *
	26	-0.0027	0.0392	-0.0420	-4.94 *	0.0210	0.0431	-0.0221	-3.03 *
	52	0.0105	0.0504	-0.0399	-3.37 *	0.0051	0.0095	-0.0044	-0.40
4	1	-0.0089	0.0003	-0.0091	-4.86 *	0.0022	-0.0032	0.0054	2.07 *
	2	-0.0126	-0.0038	-0.0088	-3.54 *	0.0038	0.0065	-0.0027	-0.73
	4	-0.0148	0.0034	-0.0182	-4.82 *	0.0105	0.0263	-0.0158	-3.02 *
	12	-0.0093	0.0124	-0.0217	-2.76 *	0.0180	0.0354	-0.0174	-2.05 *
	26	-0.0214	0.0288	-0.0502	-4.18 *	-0.0016	0.0616	-0.0632	-6.22 *
	52	-0.0125	0.0314	-0.0438	-2.74 *	0.0046	0.0522	-0.0476	-2.51 *
12	1	0.0012	0.0037	-0.0025	-0.71	-0.0002	-0.0161	0.0159	3.92 *
	2	-0.0046	0.0009	-0.0055	-1.28	0.0097	-0.0018	0.0115	2.27 *
	4	-0.0342	0.0122	-0.0464	-7.17 *	0.0100	0.0166	-0.0066	-0.86
	12	-0.0429	-0.0029	-0.0399	-3.37 *	0.0033	0.0548	-0.0516	-4.27 *
	26	-0.0364	0.0072	-0.0436	-2.19 *	0.0125	0.0710	-0.0585	-3.68 *
	52	0.0522	-0.0271	0.0793	2.61 *	0.0026	0.0553	-0.0527	-2.28 *
26	1	-0.0026	-0.0115	0.0090	2.15 *	0.0349	0.0023	0.0326	4.48 *
	2	-0.0023	-0.0069	0.0046	0.81	0.0526	-0.0075	0.0600	5.37 *
	4	-0.0122	0.0409	-0.0531	-6.89 *	0.0626	-0.0472	0.1098	5.02 *
	12	-0.0187	0.1152	-0.1339	-7.07 *	0.0632	0.0312	0.0320	1.77 *
	26	-0.0087	-0.0038	-0.0049	-0.19	0.0681	-0.0093	0.0774	2.88 *
	52	0.0241	-0.0314	0.0555	1.49	0.0885	0.0210	0.0675	1.46
		-0.0061	0.0118	-0.0179		0.0203	0.0161	0.0042	

*An asterisk, *, indicates significance at 5% of one-tail test.*

Regression of Momentum Profits Against Market States

To further examine the relationship between the profitability of momentum strategies and the market states in both the formation period and the holding period, the following regression analysis is performed:

$$R = \alpha + \beta_1 \times \text{Market}_f + \varepsilon, \tag{3a}$$

$$R = \alpha + \beta_1 \times Market_f + \beta_2 \times Market_h + \varepsilon, \quad (3b)$$

where R denotes the momentum profits for winner, loser, and winner less loser, respectively. For simplicity, the average abnormal returns across the 30 strategies for winner, loser, and winner less loser, are selected to indicate the momentum profits.

$Market_f$ is a dummy variable for the formation-period market state that assumes a value 1 if the market is in an up state, and zero otherwise, and

$Market_h$ is a dummy variable for the holding-period market state that assumes a value 1 if the market is in an up state, and zero otherwise.

Table 5 reports the regression results between the momentum profits and the market states. Panel A of Table 5 indicates a positive association between the momentum profits and the formation-period market states. The estimated coefficients for the holding-period market states, β_1, are significantly positive in both regressions. The estimated coefficient is 2.83 with a t-value of 23.96 for the first regression. This positive association suggests that the momentum profits tend to be positive following market gains and negative following market losses. The results are consistent with those documented in Cooper et al. (2004)

However, the second regression in Panel A of Table 5 indicates that the momentum profits are negatively related to market states in the holding period. The estimated coefficient for the holding-period market state, $\beta 2$, is significantly negative. The estimated coefficient for the holding-period market state is -1.04% with a t-value of -8.78. The results are consistent with the finding in Tables 3 and 4 that momentum strategies perform better in the down state of the holding period.

Panel B of Table 5 reports the sensitivity of the winner portfolios to market states. The results indicate that the average abnormal returns for the winner portfolio are positively related to the formation-period market states. The estimated coefficients for the formation-period market states, β_1, are significantly positive in both regressions. However, the second regression indicates that the sensitivity of the winners' abnormal returns to market states in the holding period is negative. The estimated coefficient for the holding-period market state is -0.77% with a t-value of -8.88.

Panel C of Table 5 indicates that the average abnormal returns for the loser portfolio are negatively related to formation-period market states. However, the abnormal returns for the loser portfolios are positively related to the market states in the holding period. The estimated coefficients for the holding-period market states, β_2, is significantly positive. The estimated coefficient is 0.28 with a t-value of 3.25. This positive association suggests that the loser portfolios are more sensitive to market states than the winner portfolios in the holding period. That is, the loser portfolio tends to perform better in an up-state holding period, but worse in a down-state holding period.

Table 5: Regression of Momentum Profits against Market State in Formation and Holding Periods.
$R = A + \beta_1 \times Market_f + \varepsilon$ and $R = \alpha + \beta_1 \times Market_f + \beta_2 \times Market_h + \varepsilon$

	α	β_1	β_2	F-value	Pr > F	R^2	adj-R^2
Panel A: Panel A: Momentum Profits (W-L) Against Market States							
Equation1	- 0.0166	0.0283		573.96	<.0001	0.0014	0.0014
Equation2	-0.0114	0.0270	-0.0104	325.60	<.0001	0.0016	0.0015
Panel B: Winner Profits Against Market States							
Equation1	-0.0248	0.0221		659.24	<.0001	0.0016	0.0016
Equation2	-0.0209	0.0211	-0.0077	369.14	<.0001	0.0018	0.0018
Panel C: Loser Profits Against Market States							
Equation1	-0.0082	-0.0063		54.92	<.0001	0.0001	0.0001
Equation2	-0.0096	-0.0059	0.0028	32.75	<.0001	0.0002	0.0002

*In Panel A, the equation 1 is: W-L = α + β 1 × Market f + ε where W-L denotes the profits of the momentum strategies. The equation 2 is: W-L = α +β 1 × Market f + β 2 × Market f, h + ε. In Panel B and C, the equation 1 is: R = α + β 1 × Market f + ε, where R denotes the profits of the winners and losers respectively, Market f denotes the market state of the formation periods. The numbers in the parentheses are t-values. An asterisk, *, indicates significance at 5% of one-tail test. The equation 2 is: R = α +β 1 × Market f + β 2 × Market h + ε, where Market h denotes the market state of the holding periods while following up or down markets.*

Discussion

Two interpretations are possible regarding the higher market sensitivity of the loser portfolio in the holding period. First, since the loser portfolio is more sensitive to market returns in the holding period, the higher sensitivity risk for the loser portfolio may be undervalued in estimating the momentum profits. If so, the positive momentum profits in a down-state holding period can be due to the underestimated risk premium for the loser portfolio. While we cannot rule out this possibility, the fact that the loser portfolio performs worse in the formation period suggests that the systematic risk for the loser portfolio probably is not high at least in the up-state formation period.

Alternatively, the higher market sensitivity for the loser portfolio in the holding period may reflect the asymmetrical reaction of investors regarding prior winner and the loser portfolios. If the market turns out to be bullish in the holding period, price reversals may be more likely for the loser portfolio. Investors may adjust their belief and consider these stocks undervalued so that price adjustment is warranted. This is especially true if the up state in the holding period becomes longer. However, if the market turns out to be bearish in the holding period, price decline may continue for the loser portfolio. Investors may reinforce their belief regarding the poor performance of the loser portfolio in the prior period. As a result, we would observe higher market sensitivity for the loser portfolio in the holding period. Such higher market sensitivity could be less likely for the winner portfolio. If the market is bullish in the holding period, certain investors may worry if the winner stocks are overpriced. In contrast, if the market is bearish in the holding period, some investors may be reluctant to adjust their positive assessment of the winner stocks. Thus, winner stocks could be less sensitive to market states in the holding period.

CONCLUSION

This paper examines the impact of market states on the profitability of momentum strategies using weekly data from the Taiwan Stock exchange over the 10-year period 1997-2006. The results indicate that market states in the formation period are positively associated with the profitability of the momentum strategies. The momentum profits are significantly positive following market gains in the formation period. In contrast, momentum profits appear to be negative following market losses in the formation period. The results are consistent with the overreaction theory developed in Daniel et al. (2004). Thus, market states in the formation period provide useful information regarding the profitability of momentum strategies in the subsequent holding period.

In addition, the empirical results indicate that market states in the holding period are negatively associated with the profitability of the momentum strategies. The momentum profits appear to be higher in a bearish holding period and lower for a bullish holding period. Moreover, the negative association between market states and momentum profits in the holding period appears to be driven by the higher market sensitivity of the loser portfolio than the winner portfolio. When compared to the winner portfolio, the loser portfolio appears to perform better in the bullish holding period. In contrast, the loser portfolio appears to perform worse in the bearish holding period. Thus, the market states in the holding period also provide information regarding the profitability of the momentum strategies.

It should be noted, however, that the empirical results documented in this paper reflect the behavior of traders in the Taiwan Stock market. The composition of traders in the Taiwan stock market indicates that individual traders account for a major part of around 70-90% of the trading volume in the sample period 1997-2006. Since these individual traders may have less access to information than institutional investors, the behavior of individual traders and institutional investors needs not be the same. Thus, any generalization of the empirical results obtained from the Taiwan Stock Exchange to other stock markets should be taken with care if the composition of traders differs drastically from that in the Taiwan stock market. Future research on how institutional investors and individual investors react in up and down markets is useful in enhancing our knowledge regarding this issue.

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IMPACT OF HEDGING PRESSURE ON IMPLIED VOLATILITY IN FINANCIAL TIMES AND LONDON STOCK EXCHANGE (FTSE) MARKET

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ABSTRACT

This paper examines the impact of net buying pressure and the event of 9/11 on the implied volatility of the U.K. FTSE 100 (Financial Times and the London Stock Exchange) index options. Our findings indicate that when effects such as financial leverage, information flow and mean reversion are held constant, the net buying pressure of the out-of-the-money put options plays a dominant role in determining the shape of the implied volatility function. Further, the event of 9/11 has a transitory influence on the implied volatility change. Our results also support the notion that hedging pressure can help explain the difference between implied volatility and realized volatility.

JEL: G11, G15

INTRODUCTION

The concept of implied volatility has received considerable attention from researchers in the past few decades. This research has addressed such questions as what drives the evolution of implied volatility and what contributes to the difference between implied volatility and realized volatility? Implied volatility is an inverse function of the option pricing models given the underlying spot asset price, exercise price, the risk-free rate of return, the remaining time to expiration, and the price of the option. Market participants such as arbitrageurs, speculators, and hedgers often treat implied volatility as market expectations in making their decisions. Arbitrageurs focus on implied volatility to profit from their actions, while hedgers pay attention to implied volatility to transfer risk to speculators. Thus, implied volatility is useful for traders to price and hedge their positions.

It is widely known that implied volatility observed in the market violates the constant volatility assumption of the Black-Scholes Option-Pricing Model (BSOPM). Many researchers have presented empirical evidence to support the notion that implied volatility exhibits persistent patterns of volatilities varying by strike, known as "volatility smile or sneer". For example, implied volatility exhibited a pattern known as "a smile" (Bollen and Whaley, 2004) in both the index option markets and the individual stock option markets prior to the market crash in October 1987. After the crash however, the pattern changed to exhibit a sneeze or skewed curve (see Dumas et al., 1998; Bollen and Whaley, 2004; Chan et al., 2004). The assumption of constant volatility in the BSOPM is relaxed in studies using deterministic models (Scott, 1987; Dupire, 1994), the GARCH generalized autoregressive conditional heteroskedasticity models (Heston and Nandi, 2000; Lehnert, 2003) and stochastic models (Hull and White, 1987; Bates, 1996, 2000). However, empirical research has shown that these models fail to explain the implied volatility smile in the global market (see Lamoureux and Lastrapes, 1993; Bates, 1996; 2000; Zhi and Lim, 2002).

Other factors such as time-varying volatility, a jump in stochastic volatility, market imperfections, and hedging pressure, have been used to explain the implied volatility pattern. However, Dumas et al., 1998; Bates, 2000; Szakmary et al., 2003 show that the time-varying, jump diffusion, and stochastic volatility models also fail to explain the implied volatility smile. From the market participants perspective, the hedging pressure theory proposed by Bollen and Whaley (2004) and documented by Chan et al. (2004), suggests that the net buying pressure of index put options caused by the limits to arbitrage mainly drives the index options premium to higher levels against a potential market crash. The result is non-constant volatility. This explanation of the volatility smile appears to be consistent with the empirical evidence observed in the US and Hong Kong markets.

This paper follows the Bollen and Whaley (2004) framework to examine whether the net buying pressure of put options influences implied volatility based on the U.K. FTSE 100 index options. In addition, we investigate whether the difference between implied volatility and realized volatility can be due to hedging pressure and the effects of the event of 9/11. Since the large demand in index put options is for the purpose of hedging portfolios from market crashes, the unexpected tragedy of the event of 9/11 could enhance hedging demand and therefore affect implied volatility. The remainder of this paper is organized as follows. The next section describes the data followed by the research methodology. The empirical results section discusses the impact of net buying pressure and the event of 9/11 on the implied volatility of the U.K. FTSE 100 index options. The last section concludes the paper.

REVIEW OF RELATED LITERATURE

Studies on the U.K. FTSE 100 (Financial Times and the London Stock Exchange) index options serve as an example. By comparing the stochastic volatility, the GARCH and the BSOPM based on the FTSE 100 index options, Lehar et al. (2002) find that the performance of the GARCH, unambiguously dominated both the BSOPM and the stochastic volatility models in terms of in-sample forecasts. However, the out-of-sample forecasts among these models showed significant differences. The predictions from most of the models showed sizable biases.

Other research on the phenomenon of the implied volatility smile focuses on the options market microstructure (Henstchel, 2003; Bollen and Whaley, 2004 and Chan et al., 2004). Market imperfection suggests that the small measurement error caused by finite quotations, bid-ask bounce effects, and non-synchronous prices between the index options and index value, result in large changes in the implied volatility (Henstchel, 2003). However, Guan and Ederington (2005) show that biased implied volatility in terms of forecasting is not attributed to measurement error. An asymmetric trading price around a bid-ask midpoint can only partially mitigate the smile of implied volatility when using the traded price to replace the bid-ask midpoint (Norden, 2003).

Shleifer and Vishny (1997) argued that the ability of professional arbitrageurs to explore mis-specified underlying financial assets is subject to the limitation of digesting the short or intermediate term loss. Liu and Lonstaff (2000) argued that margin requirements undermine the size of the potential profit taken by investors. These examples imply that market supply and demand are imbalanced without the market makers and that the prices of options will evolve with the dynamic supply, likewise, the implied volatility.

Bollen and Whaley (2004) argue that the buying pressure stems from the large demand of institutional investors in a particular index put option. This net buying pressure pushes market makers to increase the risk premium to hedge their increasing positions and risks. The findings of Chan et al. (2004) on the Hang Seng Index (HIS) options support the relationship between net buying pressure and the implied volatility proposed by Bollen and Whaley (2004). The authors discovered that net buying pressure is mainly from the demand in the OTM (out-of-the-money) put options. Guan and Ederington (2005) empirically discuss the information frown (the information content pattern of the implied volatility is a rough image of the implied volatility smile) in the S&P 500 index options. They find that the biased and inefficient implied volatility derived from the OTM put options can be due to hedging pressure instead of market imperfections. This evidence also supports the net buying pressure argument in the Bollen and Whaley (2004) study.

METHODOLOGY

Data

We obtained tick-by-tick data of the options on the FTSE 100 index for the year 2001 from EURONEXT. The options on the FTSE 100 index traded on the London International Financial Futures and Options Exchange (LIFFE) are of the European style and expire on the third Friday of the contract delivery month or the last trading day preceding the third Friday, when the third Friday is not a business day. The time-stamped options transaction data contain trading time (year, month, hour, minute and second), options premiums, types (puts or calls), trading volume, strike price and expiry date. The options premium and exercise price are quoted in index points. The contract multiplier is £10 per index point (see http://www.euronext.com for more details of tick-by-tick data)

The daily data on the London inter-bank offer rates (LIBOR) was obtained from the Yahoo Finance website with the term-to-maturity ranging from overnight to three months as a proxy for the risk-free interest rate. We obtained the annualized daily dividend rates from the FTSE 100 Group as a proxy of the dividend rates. The close values of the FTSE 100 index were obtained from the http://www.econstats.com website to form the daily continuously compounded return of the underlying asset series as follows:

$$R_t = \ln(\frac{S_t}{S_{t-1}}) \tag{1}$$

where S_t and S_{t-1} are the close values of the FTSE 100 index on day *t* and day *t-1* respectively. Further, we obtained data for the trading volume variable (V_t) from the UK Yahoo Financial website. We formed the trading volume of the FTSE 100 index as the sum of the trading volumes of the 87 constituents of the FTSE 100 index. Finally, to form a series of realized volatility, we obtained the price of the FTSE 100 index futures from the EURONEXT Company.

Given the above data set and the Black-Scholes (1973) formulae, we can back out the implied volatility as our dependent variable in examining the impact of net buying pressure on the implied volatility. To avoid the problems with thin trading and price distortion due to time decay, we excluded options with more than two months time remaining and options with less than four trading days remaining. The time decay effect is the phenomenon that options with a shorter time remaining until expiry deteriorate in value rapidly when holding other things constant. We also excluded options with a calculated implied volatility of more than 600% or less than 5%.

To obtain the net buying pressure on the implied volatility function, we classify the options into five different categories of moneyness for each option style—call and put respectively. Following Lehnert (2003), and Guan and Ederington (2005), we define moneyness as the ratio of the current index value to the option's strike price (i.e. moneyness = S/K, where *S* is the current index value and *K* is the option's strike price). The moneyness categories used in this study are shown in Table 1.

We obtain the realized volatility from the intraday prices of the FTSE 100 index futures based on the inverse function of the future fair price of $S = Fe^{-(R-D)T}$, where, *F* is the FTSE 100 index future's trading price at the trading time; *R* is the risk-free rate of interest; *D* is the dividend yield; and *T* is the time remaining to the expiration of the future contract which is less than three months. To be consistent with the options trading hours (8:00 to 17:30 GMT), we only use the data where trading hours range from 8:00 to 16:30 GMT. Further, the index value backed out from the index future price is preferred over the index price on the equity market for two reasons. One is that the future's price will converge to the value of the corresponding underlying expiration. The other is that relative to the cash market, the index futures market is more likely to be chosen by the traders to hedge their option positions. I n the cash market, the FTSE 100 index value is only recorded once every fifteen minutes due to lags in reporting the transactions of the constituents in the index. Therefore, the index value from the cash market is relatively stale owing to it lagging the futures price by a few minutes.

General Framework

We adopt the Bollen and Whaley (2004) framework to evaluate the impact of net buying pressure on implied volatility, based on the UK index options market. The general model is specified as follows:

$$\Delta\sigma_{\mathrm{t}} = \beta_0 + \beta_1 R_t + \beta_2 V_t + \beta_3 NBPP_t + \beta_4 NBPC_t + \beta_5 \Delta\sigma_{t-1} + \beta_6 D_{1t} + \beta_7 D_{2t} + \varepsilon_t \, , \tag{2}$$

where $\Delta\sigma_t$ is the change in the average implied volatility in a moneyness category from the close on day t-1 to the close on day t, R_t is the FTSE 100 index return from the close on day t-1 to the close on day t, V_t is the trading volume on the FTSE 100 index on day t in millions of pounds, $NBPP_t$ and $NBPC_t$ are the net buying pressure on the put and call option respectively, D_{1t} takes the value of one from 11 to 30 September 2001 and zero otherwise, and D_{2t} takes value of one after 10 September 2001 and zero otherwise.

The dependent variable in equation (2) measures the change in implied volatility, $\Delta\sigma_t$. In measuring the implied volatility, the bid-ask bounce effect (i.e. the trading option's price could occur either at the bid or ask) may result in calculated implied volatility shifting between the high and the low values and hence can potentially introduce measurement error. One way to avoid the bid-ask bounce effect in computing the implied volatility is to use the option's bid-ask mid point (see Corrado and Su, 1998; Bollen and Whaley, 2004). To avoid the bid-ask bounce effect and take into account the asymmetric trading price, we compute the implied volatility using the trading price for each traded option contract within a day by using the tick-by-tick data, which are time-stamped to the second. Following this, the average implied volatility based on all the calculated implied volatilities obtained from each traded option contract within the day will serve as our proxy of the day's implied volatility.

The independent variable, R_t, detects the relationship between the implied volatility and the underlying asset return. Following Bollen and Whaley (2004) and Chan et al. (2004), we employ the return on the underlying asset as our control variable. Theoretically, the movement in implied volatility negatively correlates with the returns of the underlying asset due to the leverage effect (see Black, 1976; Christie, 1982; Fleming et al., 1995).

We include the trading volume of the FTSE 100 index, Vt, as a control variable to reflect the information flow generated jointly from the implied volatility and trading volume. The expected sign of the trading volume variable is unclear. In general, the more new information introduced into the market, the greater the trading volume and the implied volatility (see Bollen and Whaley, 2004; Chan et al., 2004).

The variable, $\Delta\sigma_{t-1}$, which is lagged behind implied volatility by one period, is included to examine mean reversion as opposed to a random walk process. If the implied volatility is mean reverted, it will be forced back to its long-run mean by either known or unknown factors whenever it moves too far away from its mean. A one lagged change in the implied volatility ($\Delta\sigma_{t-1}$) is used in our study to control the effect of the mean reversion on the change in implied volatility. However, the expected sign on this variable is unclear. Bollen and Whaley (2004) argue under the case of the limits to arbitrage, the options risk premium will converge on the long run mean risk premium when market makers gradually rebalance their positions. In this case, the coefficient of $\Delta\sigma_{t-1}$ ($\beta 5$) should be negative. However, if the impact of the net buying pressure on the options risk premium is caused by the learning process (referred to as the trading activities, which follows the market expectation of future volatility), the option price and implied volatility should behave stochastically. Hence, the change in the implied volatility caused by the change in market expectations would be independent of the previous level of implied volatility and the coefficient of $\Delta\sigma_{t-1}$ (β_5) should be zero (see Bollen and Whaley, 2004; Chan et al., 2004).

Net buying pressure is defined as the difference between the number of buyer-motivated traded contracts and the number of seller-motivated traded contracts. Bollen and Whaley (2004) define the buyer-motivated (seller-motivated) contracts as the option contracts with a trading price higher (or lower) than the middle point of the prevailing bid-ask spread. Therefore, the midpoint of the prevailing bid-ask is preferred for obtaining the observations for the net buying pressure variables in our study.

Framework for the Impact of Net Buying Pressure on IVF

In this section, we transform the general framework in equation (2) into four different models according to the option's moneyness:

$$\Delta\sigma_{all,t} = \beta_0 + \beta_1 R_t + \beta_2 V_t + \beta_3 NBPP_{all,t} + \beta_4 NBPC_{all,t} + \beta_5 \Delta\sigma_{all,t-1} + \varepsilon_t \tag{3}$$

$$\Delta\sigma_{otmc,t} = \beta_0 + \beta_1 R_t + \beta_2 V_t + \beta_3 NBPP_{otm,t} + \beta_4 NBPC_{otm,t} + \beta_5 \Delta\sigma_{otmc,t-1} + \varepsilon_t \tag{4}$$

$$\Delta\sigma_{otmp,t} = \beta_0 + \beta_1 R_t + \beta_2 V_t + \beta_3 NBPP_{otm,t} + \beta_4 NBPC_{atm,t} + \beta_5 \Delta\sigma_{otmp,t-1} + \varepsilon_{1t} \tag{5}$$

$$\Delta\sigma_{otmp,t} = \beta_0 + \beta_1 R_t + \beta_2 V_t + \beta_3 NBPP_{otm,t} + \beta_4 NBPP_{atm,t} + \beta_5 \Delta\sigma_{otmp,t-1} + \varepsilon_{1t} \tag{6}$$

Equation (3) evaluates the overall effect of the explanatory variables on the change in the implied volatility. If net buying pressure significantly affects the change in implied volatility, the coefficients, β_3 and β_4, should be significantly greater than zero. The large demand in the OTM index puts mainly drives the index options risk premium to higher levels against the potential stock market decline (Bollen and Whaley, 2004). We use equations (4) to (6) to examine the impact of the OTM puts' net buying pressure on the change in the implied volatility.

Equations (4) and (5) assess the effect of net buying pressure on both implied volatilities of the call and put options across different moneynesses, respectively. The net buying pressure of the OTM put should affect the OTM put itself and subsequently on the overall change in implied volatility regardless of option type (put or call). If these are true and can be applied to the FTSE 100 index options, β_3, the coefficient of the net buying pressure of the OTM puts in equation (4), must be significantly different from zero or greater than β_4 for the OTM calls. Conversely, β_3, the coefficient for the OTM index puts in equation (5), should be greater than β_4 since the impact of the net buying pressure on the implied volatility comes from the OTM puts instead of calls. Equation (6) determines whether the net buying pressure from the OTM puts drives the evolution of the implied volatility when controlling the effects of the net buying pressure for the ATM puts which replaced the ATM calls in equation (5).

Impact of 9/11 Event On Implied Volatility

To examine the impact of the 9/11 event on the changes in the implied volatility, we adopt the following specifications, analogous to those discussed above.

$$\Delta\sigma_{\mathrm{all,t}} = \beta_0 + \beta_1 R_t + \beta_2 V_t + \beta_3 NBPP_{all,t} + \beta_4 NBPC_{all,t} + \beta_5 \Delta\sigma_{all,t-1} + \beta_6 D_{1t} + \beta_7 D_{2t} + \varepsilon_t \tag{7}$$

$$\Delta\sigma_{\mathrm{otmc,t}} = \beta_0 + \beta_1 R_t + \beta_2 V_t + \beta_3 NBPP_{otm,t} + \beta_4 NBPC_{otm,t} + \beta_5 \Delta\sigma_{otmc,t-1} + \beta_6 D_{1t} + \beta_7 D_{2t} + \varepsilon_t \tag{8}$$

$$\Delta\sigma_{\mathrm{otmp,t}} = \beta_0 + \beta_1 R_t + \beta_2 V_t + \beta_3 NBPP_{otm,t} + \beta_4 NBPC_{atm,t} + \beta_5 \Delta\sigma_{otmp,t-1} + \beta_6 D_{1t} + \beta_7 D_{2t} + \varepsilon_t \tag{9}$$

$$\Delta\sigma_{\mathrm{otmp,t}} = \beta_0 + \beta_1 R_t + \beta_2 V_t + \beta_3 NBPP_{otm,t} + \beta_4 NBPP_{atm,t} + \beta_5 \Delta\sigma_{otmp,t-1} + \beta_6 D_{1t} + \beta_7 D_{2t} + \varepsilon_t \quad (10)$$

The dummy variable D_{1t} is designed to examine whether the implied volatility has signal to the salient event as well as the succeeding market shock, whereas the dummy variable, D_{2t} is designed to examine whether the effect of the 9/11 event is transient.

The impact of the 9/11 event transferred to the implied volatility via the net buying pressure could vary in the pre-9/11, post-9/11, and during the period of 9/11. For example, if the market is more volatile during the period of 9/11 which is [9/9, 9/30], then the coefficient, β_6, should be significantly different from zero. If the impact of the net buying pressure caused by the event of 9/11 is transient, then β_7 should not differ from zero.

Bias between Implied Volatility and Realized Volatility

To determine whether the hedging pressure causes the bias between implied volatility and realized volatility, we follow Szakmary et al. (2003), Bollerslev and Zhou (2004), and Guan and Ederington (2005) information content method with the following equations:

$$RV_t = \gamma_0 + \gamma_1 IV_t + \varepsilon_t \quad (11)$$

$$RV_t = \gamma_{01} + \gamma_{11} IV_{1t} + \varepsilon_t \quad (12)$$

$$RV_t = \gamma_{011} + \gamma_{111} IV_{11t} + \varepsilon_t \quad (13)$$

where RV_t is the realized volatility, IV_t is the implied volatility before isolating the effects of hedging pressure, IV_{1t} is the implied volatility after isolating the effects of hedging pressure, and IV_{11t} is the implied volatility after isolating the effects of market imperfections.

In equations (11) to (13), the actual daily realized volatility is defined as the annualized standard deviation of the natural logarithm of the FTSE 100 index return change as:

$$RV_t = \sqrt{365 \times [\frac{1}{N-1} \sum (R_t - \overline{R})^2]} \quad (14)$$

where $R_t = \ln(\frac{S_t}{S_{t-1}})$.

Guan and Ederington (2005) argue that under the case of market efficiency, the realized volatility (RV_t) fluctuates around the market expectation (EV_t) and specifies as:

$$RV_t = EV_t + \varepsilon_t \quad (15)$$

Further, if assumptions made in Black and Scholes (1973) are true, implied volatility would be completely consistent with market expectations and hence the following equation holds:

$$IV_t = EV_t \quad (16)$$

As noted by Guan and Ederington (2005), combining equation (11) with (15) and (16), the coefficient in equation (11), γ_1, which exhibits the ability of the implied volatility to capture the information before isolating the effect of hedging pressure, should be equal to one. The expected value of R^2 equals $\frac{Var(EVt)}{Var(EVt)+Var(\varepsilon t)}$ (where R^2 is the fraction of the sample variance of the dependent variable explained by the independent variable). However, Guan and Ederington (2005) further argue that implied volatility is influenced by factors other than market expectation, and hence equation (16) should no longer hold, but instead:

$$IV_t = EV_t + \eta_t \tag{17}$$

Suppose one of the factors influencing implied volatility is hedging pressure. According to Guan and Ederington (2005), the estimated coefficient of IV_t in equation (11), γ_1, which is equivalent to $\frac{Var(EVt)}{Var(EVt)+Var(\eta_t)}$, would be downward biased from one if the variances of η_t were different cross-sectionally. The R^2, which is equivalent to $(\frac{Var(EVt)}{Var(EVt)+Var(\eta_t)})(\frac{Var(EVt)}{Var(EV_t)+Var(\varepsilon_t)})$, is also downward biased from $\frac{Var(EVt)}{Var(EVt)+Var(\varepsilon t)}$.

Accordingly, if the impact of hedging pressure on implied volatility is true, the calculated implied volatility with the effects of the net buying pressure from the index puts will not converge to one unless the market makers completely rebalance their position. In other words, the ability of the implied volatility contaminated by the hedging pressure is weakened in terms of capturing the market information. We adopt the following two-stage least squares technique to estimate the information content in implied volatility. Firstly, we regress the contaminated implied volatility equation on the instrumental variable (Z_t) as follows:

$$IV_t = \gamma_0^{'} + \gamma_1^{'} Z_t + \varepsilon_{t3} \tag{18}$$

Secondly, we use the problem-free estimators ($\gamma_0^{'} + \gamma_1^{'} Z_t$) to proxy for the variable of the contaminated implied volatility, IV_t, in equation (11). Since the contaminated part has been isolated, the expected value of the coefficient of the variable (IV_{t1}), γ_{11} in equation (12), is greater than γ_1 in equation (11). In general, it is unrealistic to expect γ_{11} to be one because only the impact of the hedging pressure is isolated. It is possible that the implied volatility is influenced by other factors than the hedging pressure, such as market imperfections caused by non-synchronous prices, bid-ask bounce effect, and other unknown factors.

The instrumental variables method can also be used to examine the impact of market imperfections on implied volatility via equation (13). We use the averaged implied volatilities from the OTM and DOTM calls as our instrumental variable (i.e. $Z_t = 1/2(IV_{OTMC,t} + IV_{DOTMC,t})$) when examining the impact of net buying pressure on the implied volatility (see Bollen and Whaley, 2004; Guan and Ederington, 2005). For the impact of market imperfections on the implied volatility, we follow Guan and Ederington (2005) using the one-lagged implied volatility ($Z_t = IV_{t-1}$) as the instrumental variable.

EMPIRICAL RESULTS

Summary of Data

Table 1 reports the average implied volatilities and the daily net buying pressure of the FTSE 100 index options over the year 2001 in terms of the moneyness category. The results show the IVF index of the puts is decreasing monotonically across the S/K-ratio categories. On the contrary, the IVF index of the calls is increasing monotonically across the same categories. Table 1 also shows that the index IVF, regardless of the option type, decreases from Category 1 to the lowest point of Category 3. From this point onward, the IVF increases to the level of Category 5, similar to Category 1. The implied volatility of the Category 1 options (DOTM puts and DITM calls) is 112.42%, which is 160.38% higher than the average implied volatility of Category 3 (DITM puts and DOTM calls) 42.82%. In general, the three different patterns in the shapes of the FTSE 100 index options can be due to the more expensive calls corresponding to the lower strikes and the more expensive puts corresponding to the higher strikes. As a result, the IVF of puts skewed more rightward, but the calls skewed more leftward in terms of categories based on the S/K-ratio.

In addition, the data in Table 1 shows that the daily net buying pressure of the put options is about 14% higher than that of call options. Among the put options, the more out-of-the money, the greater the net buying pressure. The OTM put option (Category 4) has the highest daily average net buying contracts at 91, and the highest net purchase ratio of 8.4%. That is, 52% of total net buying pressure of put options is due to the OTM put options.

Further, Table 1 also demonstrates that the net buying pressure of both put and call options correlate with the change in the implied volatility of all options. There is also an intriguing outcome, where the net buying pressure positively correlates with the change in the implied volatility for all the moneneyness put options except Category 1. On the contrary, the net buying pressure negatively correlates with the change in the implied volatility for most of the call options. Furthermore, the OTM put options (Categories 4 and 5) show greater correlation between net buying pressure and the change in implied volatility than do the ITM put options (Categories 1 and 2), while the opposite is true for the call options. For example, the correlation for the OTM put options is 0.193 but only 0.0733 for the ITM put options; inversely, the figure is 0.0091 for the OTM call options but -0.2652 for the ITM call options. Furthermore, when combining the last three columns of Table 1, it is interesting to note that for the call options, the greater the net buying pressure, the less the correlation between net buying pressure and the change in the implied volatility.

Empirical Evidence of Impact of Net Buying Pressure on Implied Volatility

Table 2 shows the signs of coefficients for the control variable (R_t) are mixed and most of them are not significant. Only the sign of the coefficient for the variable R_t in equation (3) is negative at the less than 1% level of significance. These findings are consistent with the Figlewski and Wang's (2000) results but inconsistent with the results of Bollen and Whaley (2004) and Chan et al (2004). Thus, our evidence suggests that financial leverage had no relationship with the change in implied volatility for the FTSE 100 index during the sampling period in the year 2001. One possible explanation for these empirical results is that the leverage effect on the return volatility is asymmetrical (Dumas, Fleming and Whaley, 1998). The higher ratio of equity to debt has a powerful effect on return volatility, while the lower ratio of equity to debt affecting return volatility is weak. Since Bollen and Whaley (2004) used a long period sampling data (June 1988 through December 2000) and the S&P 500 index value tended to increase during this period, the financial leverage with the higher ratio of equity to debt severely affected the return volatility. Chan et al (2004) also adopted a long sampling period of 1993 to 2000 and the HSI experienced an upward trend. Their results are similar to Bollen and Whaley's (2004) findings. On the contrary, our sampling period is only one year and as the FTSE 100 index value tended to decrease dramatically, the financial leverage with the lower ratio of equity to debt had almost no affect on the change in the implied volatility.

Table 1: Average Implied Volatility, Daily Net Buying Pressure and Correlation between Net Buying Pressure and Change in Implied Volatility across Different Moneyness

S/K-Ratio Interval	Code	Options Type	Moneyness	Abbreviation	Average Implied Volatility	NBP Contracts	NBP Ratio (%)	ρ
<0.925	1	c	Deep-out-of-the-money	DOTM	23.561	47	4.34	-0.0106
		p	Deep-in-the-money	DITM	201.28	7	0.64	-0.0091
		All Option Types			112.42	54	4.99	
[0.925, 0.975)	2	C	Out-of-the-money	OTM	30.668	65	6.00	0.0091
		P	In-the-money	ITM	84.698	-1	0.09	0.0733
		All Option Types			57.683	64	5.91	
[0.975, 1.025]	3	c	At-the-money	ATM	41.409	19	1.75	-0.0846
		p	At-the-money	ATM	44.239	57	5.26	0.2494
		All Option Types			42.824	76	7.02	
(1.025, 1.075]	4	c	In-the-money	ITM	85.212	-6	0.56	-0.2652
		p	Out-of-the-money	OTM	36.082	91	8.40	0.1930
		All Option Types			60.647	85	7.85	
>1.075	5	c	Deep-in-the-money	DITM	185.45	29	2.68	-0.1177
		p	Deep-out-of-the-money	DOTM	37.562	21	1.93	0.0412
		All Option Types			111.51	50	4.62	

This table reports the five categories of the option moneynesses. The classification is based on the ratio of S to K. S is for the FTSE 100 index price on day t in 2001 and K for the strike price of the options on FTSE 100 index at the expiry day. In addition, c is for the European call options on the FTSE 100 index and p for European put options. For the calls, the greater the S/K is, the more likely to be exercised on the expiry day. The reverse is true for the puts. The average implied volatility based on the FTSE 100 index options across the moneynesses, the daily net buying pressure, and the correlation between the net buying pressure and change in the implied volatility (ρ) are reported in the last four columns of the table. The daily average net purchase of the options in terms of the number of net buying contracts is defined as the number of buyer-motivated contracts less the number of seller-motivated contracts. The net buying ratio in terms of absolute percentage value is defined as net purchase contracts divided by daily total trading contracts.

The data in Table 2 shows the signs of the coefficients for the variable V_t are also mixed and all are insignificant. To check whether these results are sensitive to the tremendous trading volume, we use the natural logarithm of the trading volume to run the regression and the results are identical. This evidence supports the results of Bollen and Whaley's (2004) study using a different index option based on a different sample period. In their study, in spite of the corresponding estimates for most of the individual stocks being positively significant, five of six coefficients of the trading volume for the S&P 500 index in their six regressions were negatively insignificant. This evidence suggests that the information effect for the individual equity market does not always happen to the aggregate equity market. Our findings that the trading volume has no influence on the change in the implied volatility reconfirm that the aggregate market does not follow the learning process.

The results in Table 2 offer a number of interesting insights into the impact of the net buying pressure on the change in the implied volatility. The results of equation (3) show that the coefficient estimate of the net buying pressure of put options, β_3, is positive and significant at the less than 5% level of significance. On the contrary, β_4, the coefficient estimate of the net buying pressure of call options, is not significant. In addition, the results show that β_3 is significantly greater than β_4. This evidence suggests that the net buying pressure of put options has greater influence on the implied volatility movement than does the net buying pressure of call options.

In equation (4), the results show the coefficient estimates of the net buying pressure for both OTM put and call options, β_3 and β_4, are positive but not significant. To make our testing more robust, we also used the net buying pressure of the ATM call and put options to run the regression with the same results. These findings suggest that the net buying pressure of options does not affect the change in implied volatility for the OTM call options. This evidence is inconsistent with the results of past studies (Bollen and Whaley, 2004; Chan et al., 2004). There are two possible explanations for our findings differing from the two previous empirical results. One is that, as discussed previously, the impact of the net buying pressure on the implied volatility is not even across all moneyness options. In particular, the net buying pressure does not affect the change in the implied volatility for the OTM call options. The other is that our research is based on the FTSE 100 index options with a downward trend (bearish market) over the period of 2001, whereas the two previous studies were based on the S&P 500 index and HSI options with upward trends (bullish markets) over several years respectively.

With respect to equation (5), the fourth column of Table 2 shows that the coefficient of the net buying pressure for the OTM put options, β_3, is 0.006002 with a t-statistic of 3.81, which is significant at the less than 1% level of significance. However, the estimate of β_4 for the ATM call options is negative but not significant. This evidence is consistent with the results of the previous studies (Bollen and Whaley, 2004; Chan et al, 2004). It proposes that the net buying pressure of the OTM puts influences the change in the implied volatility of the OTM put option itself relative to that of the ATM call option. In addition, the equality of these two coefficients (β_3=β_4) in equation (5) is strongly rejected at the 1% level of significance by executing the Wald coefficient restriction test. This means the net buying pressure is due to the limits to arbitrage instead of the learning process since β_3 is significantly greater than β_4. As discussed above, if the investors' trading activity follows the expectation of future volatility, there is no reason for them to prefer one moneyness option to another. Furthermore, the ATM option should carry more weight in determining the shape of the implied volatility since it is more sensitive to volatility. Since our ATM option ranges from 0.975 to 1.025 according to the S/K-ratio, the ATM options are regarded as moneynesses that are more informative. On the contrary, the limits to arbitrage result in the net buying pressure on the OTM puts being more important than others due to hedging the market crash.

Comparing equation (5) to equation (6), the only change is replacing the ATM call options in equation (5) for the ATM put options in equation (6) (see Table 2). Therefore, equation (6) is an attempt to further test whether the change in the implied volatility for the OTM put option stems from the net buying pressure of the put option when holding the net buying pressure of the OTM put options constant. The results of the coefficient estimates (β_3 and β_4) of the net buying pressures of both OTM and ATM put options are 0.004825 and 0.006472 respectively with significance of less than 1%. When these results are compared to the corresponding results of equation (5), the net buying pressure of the ATM put options influences the change in the implied volatility of the OTM put options more than does the net buying pressure of the ATM call options. The net buying pressure of the put options has more weight in determining the shape of the implied volatility for the FTSE 100 index options (see Table 2). All the coefficient estimates of one lagged change in the implied volatility corresponding to the option sources are negative and significant at the less than 1% level. They hover around a value of -0.428 regardless of option type and moneyness. This means the price will return to about 40% of the previous level on the next trading day. Hence, the net buying pressure is caused by the limits of arbitrageurs instead of the learning process and its impact on the change in the implied volatility is transitory.

However, the above results could be due to measurement error. Bollen and Whaley (2004) argue that the changes in the implied volatility for $\Delta\sigma_t$ and $\Delta\sigma_{t-1}$ are based on three consecutive days. Therefore, the implied volatility for day t-1 has to be used twice with the opposite sign in the calculation of both $\Delta\sigma_t$ and $\Delta\sigma_{t-1}$. To some degree there will be a negative serial correlation in the observed change in implied volatility due to the measurement error including the bid-ask spreads and the non-syncronous record between the index and option price.

Table 2: Impact of Net Buying Pressure on Implied Volatility

Equation	**3**	**4**	**5**	**6**
Dependent Variable	$\Delta\sigma_{all,t}$	$\Delta\sigma_{otmc,t}$	$\Delta\sigma_{otmp,t}$	$\Delta\sigma_{otmp,t}$
NBP_1 Source	All put	OTM put	OTM put	OTM put
NBP_2 Source	All call	OTM call	ATM call	ATM put
β_0	1.892331	-26.47512	1.320781	0.941685
	(0.621)	(-1.283)	(0.486)	(0.355)
β_1	1.522972***	-2.90083	0.265344	0.154311
	(-3.154)	(-0.885)	(0.602)	(0.357)
β_2	-3.98E-11	4.45E-10	-2.94E-11	-2.85E-11
	(-0.778)	-1.285	(-0.645)	(-0.641)
β_3	0.001296**	0.00765	0.006002***	0.004825***
	(2.193)	(0.642)	(3.814)	(3.058)
β_4	0.001163	-0.004436	-0.001934	0.006472***
	(1.193)	(-0.433)	(-0.696)	(3.328)
β_5	-0.384998***	-0.492087***	-0.427709***	-0.420164***
	(-6.372)	(-8.502)	(-7.013)	(-7.048)
R^2	0.198617	0.253182	0.211952	0.24731
Adj- R^2	0.180648	0.236586	0.19444	0.230583
Numbers of Observations	229	231	231	231
F-Statistic(β_3=β_4)	0.012251	0.54441	6.150947***	0.355084

*This table reports the empirical results of the impact of the net buying pressure on the change in the implied volatility in different moneyness categories according to equations (3) through (6). The dependent variable is change in implied volatility. The independent variables are return on the underlying asset, r_t, trading volume of the FTSE 100 index, v_t, net buying pressure variables, NBP_1 and NBP_2, and one-period lag of implied volatility, $\Delta\sigma_{t-1}$. The coefficients of independent variables are as follows. β_0 is the intercept term. β_1 to β_5 are coefficient of r_t, V_t, NBP_1 and NBP_2, and $\Delta\sigma_{t-1}$, respectively. The *, **, and *** denote significant level of 10%, 5% and 1% respectively. T-statistics are reported in parentheses.*

Table 3 summarizes the results of equations (7) through (10) including the impact of 9/11 on the change in the implied volatility. Table 3 offers a number of intriguing findings. Note first that the estimates of the coefficients of the net buying pressure in equations (7), (9) and (10) are significantly positive, which is consistent with the results in equations (3), (5) and (6). The magnitudes of the net buying pressure of the OTM put options increase by about 1%, 2.25% and 3.03% respectively. This suggests that the tragedy of 9/11 exerted net buying pressure on the change in implied volatility through the net buying pressure of the OTM put options. The estimates of coefficients β_6 and β_7 represent the dummy variables (D_{1t}, 1 for the period of [9/9, 9/30] and zero otherwise; dummy D_{2t} is one for after 9/10, 0 otherwise), and further confirms that the 9/11 event exerted net buying pressure on changes in the implied volatility. The results in Table 2 show the signs of β_6 are mixed but negative for β_7 in equations (7) through (10). However, the estimate of β_6 in equation (9) is about 4.98 with a less than 5% level of significance. The estimate coefficient β_6 in equation (10) is about 4.25 and significant at the less than 10% level of significance. The findings indicate that the effect of the tragedy of 9/11 occurring in the US did spillover to the UK index option market during the period of [9/9, 9/30] and the ensuing jump fear made the market more volatile through the net buying pressure of the OTM put options. However, the shock was transitory. This is confirmed by estimates of the coefficients β_7 in equations (7) through (10) being non-significantly different from zero.

Table 3: Impact of 9/11 Event on Implied Volatility via Net Buying Pressure

Equation	**7**	**8**	**9**	**10**
Dependent Variable	$\Delta\sigma_{all,t}$	$\Delta\sigma_{otmc,t}$	$\Delta\sigma_{otmp,t}$	$\Delta\sigma_{otmp,t}$
NBP_1 Source	All put	OTM put	OTM put	OTM put
NBP_2 Source	All call	OTM call	ATM call	ATM put
β_0	2.951928	-29.1396	2.750469	2.085402
	(0.864)	(-1.345)	(0.972)	(0.753)
β_1	-1.477515***	-2.92329	0.351762	0.237305
	(-3.035)	(-0.886)	(0.798)	(0.548)
β_2	-3.98E-11	5.11E-10	-5.42E-11	-4.67E-11
	(-0.723)	(1.353)	(-1.103)	(-0.97)
β_3	0.001309**	0.00731	0.006137***	0.004971***
	(2.184)	(0.609)	(3.909)	(3.146)
β_4	0.001217	-0.00442	-0.002038	0.006203***
	(1.243)	(-0.43)	(-0.736)	(3.179)
β_5	-0.387558***	-0.49142***	-0.439837***	-0.43119***
	(-6.39)	(-8.454)	(-7.213)	(-7.216)
β_6	1.736652	-5.64257	4.983275**	4.253643*
	(0.446)	(0.093)	(4.194)	(3.16)
β_7	-1.124654	-2.23426	-0.909519	-1.03182
	(0.624)	(0.055)	(0.536)	(0.719)
R^2	0.177752	0.253847	0.226633	0.258356
Adj- R^2	0.151708	0.230425	0.202357	0.235076
Numbers of Observations	229	231	231	231
F-Statistic(β_3=β_4)	0.005717	0.506807	6.581213***	0.196475

This table reports the empirical results of the impact of 9/11 event on the change in the implied volatility in different moneyness categories via net buying pressure according to equations (7) through (10). The dependent variable is change in implied volatility. The independent variables are return on the underlying asset, r_t *, trading volume of the FTSE 100 index,* v_t*, net buying pressure variables,* NBP_1*and* NBP_2*, one-period lag of implied volatility,* $\Delta\sigma_{t-1}$*, and dummy variables,* $\Delta\sigma_{t-1}$*, and* D_{2t}*. If the data period is between 9/9 and 9/30,* D_{1t} *takes the value of one on day t, and zero otherwise.* D_{2t}*is equal to one for the period of post-9/11, zero otherwise. The coefficients of independent variables are as follows.* β_0*is the intercept term.* β_1*to* β_7 *to are coefficient of* r_t *,* v_t*,* NBP_1*and* NBP_2*,* $\Delta\sigma_{t-1}$*,* D_{1t}*, and* D_{2t}*, respectively. The * , ** , and *** denote significant level of 10%, 5% and 1% respectively. T-statistics are reported in parentheses.*

In summary, the above tests demonstrate strong statistical support for the contention that the net buying pressure of put options drives the implied volatility higher against a market crash when controlling for the effects of financial leverage, information flow and mean reversion. Furthermore, the nature of the change in the implied volatility is options specific for the FTSE 100 index options. Net buying pressure of the OTM put options plays a dominant role in determining the shape of the implied volatility.

Hedging Pressure as an Explanation of Bias of the Implied Volatility from the Realized Volatility

The correlation between the net buying pressure of the put options and the implied volatility minus the realized volatility is approximately 0.15. The coefficient is approximately 0.11 for the correlation between the net buying pressure of the call options and the implied volatility less realized volatility. This suggests that that implied volatility biased away from realized volatility is most likely caused by the net buying pressure of the put options when considering the net buying pressure of the call options, which has no influence on the change in the implied volatility as shown in the previous section.

Table 4 reports the empirical results of equations (11), to (13). The F-statistics of 223 and 165.82 indicate that the impacts of both of our instrument variables are not weak. According to Stock and Watson (2003), if the first-stage F-statistic exceeds 10, the instrument should not be weak. Our findings from Table 4 show that γ_I, the estimate of the coefficient of IV_t in equation (11), is about 0.17 with a t-statistic of 1.8, which is significant at the less than 10% level of significance. However, $\gamma_{II,}$ the estimate of the coefficient of $IV_{tI,}$ is approximately 0.33 and significant at the less than 1% level of significance. Therefore, after using the instrument variable to isolate the effect of hedging pressure on the implied volatility of all options, $\gamma_{II,}$ the coefficient of the implied volatility, increased by about 17% relative to γ_I. In addition, the R^2 in instrumental variable regression (12) also increased by nearly 20% relative to OLS regression (11). This evidence suggests that hedging pressure does contribute to implied volatility biasing away from the realized volatility.

In contrast, the instrumental estimates based on the instrumental variable $Z_t = IV_{t-1}$ in equation (13) is about 0.2, but not significant. The R^2 is about 0.02, which is almost the same as the OLS in equation (3). This evidence suggests that there is no obvious measurement error effect. The result also confirms that the mean reversion cannot be due to the measurement error discussed previously.

Our findings from these two instrumental variables tests indicate that the difference between implied volatility and realized volatility is caused by net buying pressure rather than measurement error. However, hedging pressure is not able to be completely responsible for their difference because the coefficient restriction tests strongly reject both $\gamma_I = 1$ and $\gamma_{II} = 1$. Therefore, there are other unknown factors contributing to the difference between implied volatility and realized volatility.

Table 4: Bias between Implied Volatility and Realized Volatility

Equation	**11**	**12**	**13**
Independent Variable	IV_t	IV_{tI}	IV_{tII}
Coefficient Value	0.169678*	0.327599***	0.202429
T-Statistic	(1.80)	(6.16)	(1.51)
R-Squared	0.021897	0.219752	0.019202
Adjusted R-Squared	0.015151	0.213559	0.011828
First Stage F-Statistic		223.3963	165.8223
Number of Observations	147	128	135
F-Statistic(γ=1)	77.73368	159.8399	35.31285

*This table reports the empirical results of the information contained in the implied volatility according to the equations (11) through (13) and estimated using the two-stage least squares regression via instrumental variables method. The * , ** , and *** denote significant level of 10%, 5% and 1% respectively. T-statistics are reported in parentheses.*

CONCLUSIONS

In this paper, we adopt the Bollen and Whaley (2004) framework to examine whether the net buying pressure of put options influences implied volatility based on the U.K. FTSE 100 index options. Further, we investigate whether the hedging pressure and the effects of the 9/11 event on implied volatility can explain the difference between implied volatility and realized volatility. Finally, we examine whether the biases in using implied volatility to forecast volatility is due to hedging pressure based on the instrumental variable regression.

Our findings are generally consistent with the results from Bollen and Whaley (2004) and Chan et al (2004) and hence support the hedging pressure theory in the UK market. The empirical results show that the FTSE 100 index options-based implied volatility derived from the BSOPM exhibits the classical smile surface during the period of year 2001. The implied volatility function displays a steep slope. The tests statistics show that put options are under heavier net buying pressure relative to call options. The OTM put options have the highest net buying pressure.

Our regression results indicate that the evolution of implied volatility is options specific. Put options, particularly the OTM put options, play a dominant role in determining the shape of implied volatility. Based on the results of a negative correlation between implied volatility and its lagged one, and that the change in implied volatility is sensitive to both options type and moneyness specific, the net buying pressure is caused by the limits to arbitrage instead of the learning process.

In addition, we find that the salient event of 9/11 influenced the change in implied volatility for the OTM put options. The impact is transitory because the dummy variable estimate is significantly positive only during the period of 9/9 through 9/30. This evidence is also consistent with the mean reversion. Finally, our results show that the difference between implied volatility and realized volatility stems from the net buying pressure instead of measurement error. Therefore, hedging pressure can explain the steep slope of the implied volatility.

In this paper, we only consider the implied volatility of options where the underlying asset is the stock index. It is therefore interesting to extend this research to consider the behavior of implied volatility for option on individual stocks and examine whether any significant differences exist in the underlying asset, which is not the stock index. Finally, another possible venue of future research is to extend this study to international markets and to more recent.

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THE LONG-TERM PERFORMANCE OF PARENT FIRMS AND THEIR SPIN-OFFS

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ABSTRACT

This study examines the four-year stock performance of firms that undertake a spin-off. The theoretical motivations for spin-offs have been widely documented in the literature; however, an empirical examination of the aftermarket performance of spin-offs across a protracted bear market remains an unexamined topic. I find that spin-offs and their parents consistently outperform market indices from the closing price on the first day of public trading to their four-year anniversaries. These findings are important because the existence of price patterns during a market correction can serve as an investment hedge within a mean-variance efficient portfolio.

JEL: G30, G38

INTRODUCTION

Complete and perfect market theory predicts that a spin-off announcement should not alter firm value. However, if management is unable to replicate the role of financial markets, then capital may be misallocated and a spin-off, by improving investment decisions, may enhance the value of the divested assets. Schipper and Smith (1983) contend that the creation of publicly traded firms results in new information that empowers stakeholders to more closely monitor the activity of managers, thereby reducing agency costs and enhancing shareholder wealth. The prevailing view of spin-offs in corporate finance theory is that the price of a firm's stock should reflect the market's best estimate of the firm's long-run value. If, however, the stock value of a diversified corporation sends a weak signal of the productivity of any given division, then a single-product stock could dominate a diversified-product stock.

A spin-off is defined as a segment of a corporation that is established as an independent company, and stockholders receive shares in this new entity on a pro-rata basis (Allan, 2001). Several studies have found empirical evidence of the positive short-term announcement returns surrounding the public offering of spin-offs. (Hite and Owers, 1983; Schipper and Smith, 1983). However, the long-term performance of a portfolio invested in spin-offs across a protracted bear market remains a largely unexamined topic in the literature. There are several important reasons why the study of the long-run performance of spin-offs is of interest. First, from an investor's perspective, the existence of price patterns during a market correction may present investment opportunities for active traders. Spin-offs represent an important hedge if they produce excess returns during a market contraction. Finally, from a corporate finance perspective, the cost of external capital for newly issued spin-offs depends not only on the issuance cost of going public, but also on the returns that investors expect in the secondary market (Ritter, 1991).

To summarize the empirical findings of this paper, the average four-year holding period returns for a portfolio of 101 spin-offs that went public from 1999 to 2003 is 29.78%. The holding period return is measured from the closing market price on the first day of issuance to the spin-off's 48-month anniversary. Moreover, two of the four years of seasoning occurred across the bear market experienced in 2001 and 2002. I document that every dollar invested in a portfolio of spin-offs purchased at the closing market price on the first day of trading and held until the four-year anniversary results in a terminal wealth of $1.2978 (29.78% gain); while every dollar invested in the equally weighted stock market index over the same period results in a terminal wealth of just $.9150 (8.50 loss). In the long-run, spin-offs outperformed the market. The remainder of the paper is organized as follows: in the next section, I discuss the relevant literature. Next, I discuss methodological considerations. The paper continues with a discussion of the data and the presentation of the empirical results. The paper closes with concluding comments.

LITERATURE REVIEW

The literature on spin-offs identifies a number of motivations for undertaking these corporate restructurings. First, spin-offs often emerge in connection with technological innovations. Second, spin-offs are sometimes pursued to increase a subsidiary's direct access to the capital markets, instead of competing for limited parental funds as a non-core subsidiary within a multi-line company (Cooper, 1985). Finally, spin-offs are often economically efficient since the market will frequently value the parent and spin-off at a higher combined market capitalization than just the parent alone. Upward equity analysts' valuations following a spin-off are predictable when the parent and spin-off firms are in vastly different business lines or if there are different earnings growth prospects.

It is important to note that firms can also increase their focus by selling unrelated core assets to other firms rather than initiate a spin-off (Desai and Jain, 1999). However, there are several inherent advantages to concentrating on a sample of spin-offs. First, asset sales are often motivated by liquidity constraints or a desire to pay down debt. However, unlike the sale of assets, a spin-off does not involve an exchange of cash. Thus, a spin-off is generally not motivated because of the parent firm's desire to generate immediate cash. Third, Vijh (2002) finds that asset sales are frequently undertaken to provide a cheap source of funds, not necessarily to improve efficiency. However, a study of spin-offs is free from this confounding effect. Finally, since both the parent and the subsidiary trade as separate entities after the spin-off, their subsequent performance can be analyzed separately. This type of analysis is not feasible in asset sell-offs.

Previous studies find significant abnormal announcement returns ranging from 2.9% to 3.3% on the announcement of spin-off depending upon the benchmark index and event window. These studies include Hite and Owers (1983), Miles and Rosenfeld (1983), Schipper and Smith (1983), Linn and Rozeff (1985), Copeland, Lemgruber and Mayers (1987), Vijh (2002), and Krishnaswami and Subramaniam (1999). These gains are attributed to improved managerial efficiency, reduced information asymmetry, and relaxed regulatory and tax constraints. The elimination of negative synergies is the most commonly cited reason for positive stock price reaction to spin-offs. Schipper and Smith (1983) argue that, just as there are economies of scale, there may be diseconomies of scale by combining disparate assets under one corporate umbrella. As the firm becomes more complex, the costs of decision management and decision control increase to the point where they may offset economies gained from increasing firm size. Hite and Owers (1983), Miles and Rosenfeld (1984), and Desai and Jain (1999), among others, note that spin-offs improve managerial efficiency by reducing the potential for misallocation of capital, eliminating cross subsidies, and enabling improved investment decisions. Krishnaswami and Subramaniam (1999) suggest that many spin-offs are motivated by the desire to mitigate the information asymmetry problem, as the market is forced to value the stock of a company that is really a portfolio of several dissimilar lines of businesses. The findings Desai and Jain (1999) that the announcement returns for focus-increasing spin-offs are significantly greater than the non-focus-increasing spin-offs are also consistent with an information asymmetry hypothesis.

At least two published academic studies have examined the long-term aftermarket performance of spin-offs prior to this study. Cusatis, Mikes, and Woolridge (1993) find significant long-term excess returns for a sample of spin-offs from 1965 to 1988. However, the majority of their excess returns are limited to spin-offs that were acquired within three years of their going public. Desai and Jain (1999) document significant short-term excess returns for spin-offs in the twelve months following public issuance. Their analysis contains 155 spin-offs that originated during the 1975 to 1991 period. My research differs from these studies because of a more recent sample period from 1999 to 2003. This period is distinct from other decades due to the two-year bear market that lasted from 2001 to 2002. In addition, my sample is free of the slight survivorship bias observed in Cusatis, Mikes, and Woolridge (1993). In addition, to ensure that my return calculations are robust and not methodology specific, I employ several sensitivities on the matched market return portfolios.

Many rationales for the positive announcement effects of spin-offs have been proposed in the literature. Allen, Lummer, McConnell and Reed (1995) argues that spin-offs benefit the firm since, after the spin-off, the equity values of the securities traded provide a much "cleaner" signal of managerial productivity than when the two divisions were part of a combined firm. The argument is that this enables the firm to provide better incentives for firm management based on the stock price of the individual firms. Habib, Johnsen and Naik (1997) argue that spin-offs improve the quality of the information managers and uninformed investors can infer from the prices of the firm's traded securities, therefore leading to an increase in the expected price of the firm's equity. The firm may be undervalued if the market cannot observe the cash flows of each individual division of that firm. Therefore, the firm that needs external financing may resort to divestures such as spin-offs so that it can raise capital at a fair market price after the divesture.

DATA AND METHODOLOGY

I identify the sample of parent firms that undertake spin-offs from the following sources: (i) stock distributions by firms trading on the NYSE, Amex, and NASDAQ that the Center for Research in Security Prices (CRSP) identifies as spin-offs, and (ii) news wires and articles on Lexis-Nexis, Barron's and the Wall Street Journal that report spin-off transactions. I next match the spin-off sample firms with their monthly stock prices from CRSP. To avoid survival bias, I calculate returns for event firms until they are delisted on CRSP. This results in a sample of 101 spin-offs that went public between the years of 1999 to 2002. For the spin-offs that went public between 1999 and 2002, the long-term returns are calculated from issuance until their four-year anniversary. This calculation applies to 82 of 101 spin-off firms in my sample.

I define a market correction as a drop of market levels of at least 10%, but not more than 20%. The most recent example of a correction was the stock market downturn during third quarter 2001. Dismal labor and retail numbers pushed the stock market into a correction. By September 17, 2001, the first day of trading after the September 11 attacks, the Dow Jones Industrial Average had plunged 684.81 points to 8,921. That loss officially pushed the Dow into a bear market, which lasted until December 2002. Shortly thereafter, the stock market downturn of 2002 pushed the Dow and NASDAQ from 10,000 and 2,000 levels in March, to five- and six- year lows of 7,200 and 1,100 by October 2002.

To evaluate the long-run performance of the parents of spin-offs, two measures are used: four-year cumulative average adjusted returns (CAR) are calculated in excess of the market benchmark and four-year buy and hold returns are calculated for sample firms and market index. The methodology used in this paper is similar to that used in other long-term return studies (Rauterkus and Song (2005); Barber and Lyon (1997)). Monthly market-adjusted returns are calculated as the monthly raw return on a stock minus the monthly benchmark market return for the corresponding 21-day trading period.

Fama (1998) documents that long-term return estimation is often sensitive to the methodology used. To avoid methodology specific bias, sensitivities are run on market benchmarks to include (1) an index of the CRSP equally weighted stock market index, and (2) an index of the CRSP value-weighted index. Canina, Michaely, Thaler and Womack (1998) observed that the equally weighted CRSP index sometimes results in an upward overestimation of long-term returns, therefore the value-weighted index is used as a robustness check to avoid methodology-specific bias. Therefore, the benchmark-adjusted return (ar_{it}) for stock i in event month t is calculated for the spin-off from initial pricing to the firm's 48-month anniversary, defined as

$$ar_{it} = r_{it} - r_{ew} \tag{1}$$

$$ar_{it} = r_{it} - r_{vw} \tag{2}$$

where r_{it} is the stock return for event firm i in month t, r_{ew} is an index comprised of equally weighted stocks and r_{vw} is the value-weighted index from CRSP.

The average benchmark-adjusted return on a portfolio of n stocks for event month t is the equal-weighted average of the benchmark-adjusted returns for

$$\breve{\underset{.}{A}}R_t = 1/N \sum_{i=1}^{N} ar_{it} \tag{3}$$

where N is the number of spin-off firms.

The cumulative benchmark-adjusted aftermarket performance (CAR) from event month q to event month s, where q is January of the event year and s is December, is represented by the equation:

$$CAR_{q,s} = \sum_{t=q}^{s} \breve{\underset{.}{A}}R_t \tag{4}$$

As an alternative to the CAR, which implicitly assumes monthly portfolio rebalancing, a 48-month holding period return, R_i, for firm i is calculated as:

$$R_i = \prod_{t=1}^{48} (1 + r_{it}) - 1 \tag{5}$$

These equations measures the total return from a buy and hold strategy where a stock is purchased at the first closing market price after a spin-off goes public and held for a four-year period. To interpret this four-year total return, I compute the buy and hold excess returns (BHER) for firm i as:

$$BHER_i = R_i - R_{mkt} \tag{6}$$

where R_i is the holding period return for \$1 invested in a portfolio of spin-offs and R_{mkt} is the holding period return for \$1 invested in the market portfolio. A BHER that is greater than zero is interpreted as a portfolio of the parent firms of spin-offs outperforming the market and a BHER of less than zero indicates a portfolio of the parent firms of spin-offs underperforming.

RESULTS

Table 1- Panel A reports that the top performing spin-off over the sample period, Cavco, experienced the highest annualized return of 371%. Cavco Industries went public at an opening list price of \$12.70. Cavco Industries primarily focuses on the production of manufactured homes in the southwestern United States. Although not reported in Table 1, I analyze the industry affiliations of the 101 parent firms that engaged in a spin-off and the 101 subsidiaries that were spun-off, and find that the distribution of subsidiaries is spread evenly across sixteen distinct industries with no one industry having a concentration of more than four spin-offs. Based on Lexis-Nexis news runs, the motives most often cited for spinning-off a subsidiary are improvement of business focus, enhanced market valuation of the separate entities, and increased access to capital markets. These motives are similar to the academic explanations for spin-offs presented in the literature review section of this paper.

Table 1: High and Low Performing Corporate Spin-offs

Spin-off/Parent	Spin Date	Initial Price ($)	Annual Return (%)	Result
Panel A: High Performing Firms				
Cavco/Centex	06/08	12.70	371.4	Active
Mind Speed/Conexant Sys	06/08	2.05	304.0	Active
Gen Probe/Chergal	09/08	6.50	219.4	Active
Brookfield Homes/Brookfield	01/08	8.50	195.1	Active
Hdsn Highland/Mnstr VW	03/08	9.00	174.7	Active
Medco Health	08/08	22.60	110.9	Active
Marine Products/RPG	02/08	3.07	86.1	Active
Fordin/Can Pacific	02/08	22.00	83.0	Active
Genesis Healthcare/Genesis	11/08	15.50	74.0	Active
MWE Development/Magna Intl	08/08	17.65	70.0	Active
Plains E&P/Plains Res	12/08	9.20	71.5	Active
Tender Lvg Cr/Staff Blds	10/99	0.10	68.8	Active
Ceva/DSP Group	11/08	5.10	64.6	Active
Levitt/BankAtlantic	12/08	16.00	63.0	Active
ENPRO/Goodrich	05/08	8.45	57.5	Active
Duck Head/Delta Wdide	06/00	4.50	49.7	Active
Panel B: Low Performing Firms				
VelocityHI/BREProps	08/00	2.00	-99.2	Active
Key3 Media/Softbank	08/00	5.50	-95.7	Delisted
APW/Actuant	07/00	37.00	-97.6	Delisted
ANC Rental/AutoNation	06/00	6.50	-92.6	Delisted
Wllms Com/Wllms Cons	09/99	23.00	-92.4	Delisted
Tikcro Tech/Orckit Com	06/00	15.00	-91.1	Delisted
Gentek/GeneralChem.	04/99	13.00	-82.9	Delisted
Seera Nova/Intelligroup	07/00	10.63	-81.0	Acquired
Crstlne Cptl/Hst Mrrit	12/98	14.50	-79.0	Active
Circle.com/Snyder Comm	10/99	13.00	-77.9	Acquired
eLoyalty/Tech sol	02/00	350.00	-63.1	Active
Lanier Worldwide/Harris	11/99	6.50	-47.5	Acquired
McData/EMC	02/08	53.50	-45.2	Active
Overhill Farms/Overhill	11/08	3.00	-39.0	Active
VIAlta/ESS Tech	08/08	1.40	-38.2	Active
Synavant/IMS Health	08/00	12.50	-33.5	Acquired
Roxio/Adaptec	05/08	13.00	-30.6	Active
Evercel/Energy Res	03/99	3.13	-30.4	Active

The top performing parent firms and their spin-offs are presented from a sample of over one hundred firms. These firms went public during the years 1999-2003. Firms are ranked by highest annualized returns from high to low during the January 1, 1999 to December 31, 2003 sample period.

Table 1 – Panel B reports that approximately half of the low performing spin-offs declared bankruptcy within three years of their issuance. The empirical literature finds the performance of spin-offs with respect to growth and survival is generally above the average survival rate on an initial public offering. This result is, however, conditional on the parent being financially and strategically healthy. The positive effect of a strong parent-progeny relationship is consistent with evolutionary organizational theory's predictions that routines and procedures are inherited by the spin-off from the parent (Helfat and Lieberman, 2002). Other research documents the importance of a deliberate planning process as a key predictor factor in the survival of the spin-off. Research has found that this direct support is more

important in predicting the stand-alone success than either the product line or type of business. Furthermore, Klepper (2001) finds there are much higher observed survival rates from parental initiated spin-offs when compared to entrepreneurial initiated spin-offs.

As a cautionary tale to potential spin-off investing, Table 1 – Panel B reports that Velocity experienced the lowest annualized return of -99.2% of any firm in the sample. Velocity is a high-speed internet company based in San Francisco that went public in August 2000. Shortly after issuance, Velocity entered bankruptcy protection in the early fall of 2001. In conjunction with this bankruptcy announcement, Velocity declared that it entered into an asset purchase agreement with Reallinx of Dallas. Velocity sold its major assets for $350,000. However, given the company's outstanding debt, Velocity did not expect any distributions to stockholders following the transaction because of the size of the company's outstanding debt.

Table 2 – Panel A reports the average matching firm-adjusted returns (AR) and cumulative average value-weighted stock market index adjusted returns (CAR) for the first 48 months after the offering date for Spin-Off firms. Thirty-six of the 48 monthly average adjusted returns are positive. The positive average adjusted returns, after a slight decrease in the first 10 months of seasoning, increases to 137.83% by the end of month 48. Focusing first on the raw returns, a positive initial return of 11.71% is followed by monthly average raw returns varying between negative 3.48% and positive 8.51%. The cumulative average raw return peaks at 101.64% in month 48. The poor performance of the stock index is strongly influenced by the bear market that lasted through December 2002 and by the September 11, 2001 tragedies.

Table 2 – Panel B reports the average matching firm-adjusted returns (AR) and cumulative average market index adjusted returns (CAR) for the first 48 months after the offering date for the Parents of Spin-offs. Twenty-nine of the 48 monthly average adjusted returns are positive, while 19 are negative. The cumulative average market adjusted returns, after a slight decrease in the first four months of seasoning, increases to 105.64% by the end of month 48. The cumulative average raw return peaks at 122.32% in month 40. It is important to note that the poor performance of the entire market was influenced by the bear market experienced from September 2001 to December 2002 and by the September 11, 2001 tragedy. Yet, the spin-offs and their parent firms beat their matching value-weighted market indices.

Table 3-Panel A reports the average matching firm-adjusted returns (AR) and cumulative average market adjusted returns (CAR) for Spin-offs after the first 48 months of their offering date. Thirty-five of the 48 monthly average adjusted returns are positive. The positive average adjusted returns, after a slight decrease in the first 10 months of seasoning, increases to 137.83% by the end of month 48. Focusing first on the raw returns, a positive initial return of 15.53% is followed by monthly average raw returns varying between negative 16.44% and positive 14.37%. The cumulative average raw return peaks at 150.93% in month 40. The measurement of the long-run performance of stocks is especially sensitive to the benchmark employed. This is commonly documented in event studies using long windows, as indicated by Dimson and Marsh (1986). Since the vast majority of the Spin-offs that trade on NASDAQ, a natural candidate for the most appropriate benchmark portfolio is the small stock market index. This index has the advantage that the industry mix more closely matches the sample spin-offs than does either the AMEX or NYSE.

Table 2: Value-Weighted Cumulative Adjusted Returns for Parent Firms and their Spin-offs

Month of	***Panel A***		***Panel B***	
Seasoning	**Spin-offs AR %**	**Spin-offs CAR %**	**Parents AR %**	**Parents CAR %**
1	11.71	11.71	-4.73	-4.73
2	8.51	20.22	0.23	-4.50
3	4.22	24.44	-4.15	-8.65
4	4.17	28.61	8.15	-0.50
5	0.79	29.40	11.92	11.42
6	2.96	32.36	-0.01	11.41
7	5.04	37.40	6.78	18.19
8	3.39	40.79	-3.94	14.25
9	-1.84	38.96	-2.95	11.30
10	-3.18	35.78	-7.13	4.17
11	4.56	40.34	3.55	7.73
12	3.87	44.21	3.47	11.20
13	2.96	47.16	5.24	16.44
14	1.59	48.76	-7.99	8.44
15	6.01	54.76	9.48	17.92
16	3.19	57.96	23.96	41.89
17	5.52	63.48	5.90	47.78
18	-1.98	61.49	-1.91	45.87
19	-0.24	61.25	4.08	49.95
20	4.44	65.70	-2.00	47.95
21	-2.67	63.03	12.47	60.42
22	-3.32	59.72	-0.07	59.71
23	-0.63	59.09	3.45	63.15
24	4.81	63.91	2.50	65.66
25	2.38	66.28	12.19	77.85
26	-1.00	65.28	18.43	96.28
27	-1.42	63.86	2.18	98.46
28	4.34	68.20	-6.86	91.60
29	1.80	70.00	7.50	99.10
30	5.79	75.79	0.06	99.16
31	1.13	76.92	-0.49	98.67
32	1.13	78.05	4.88	103.54
33	1.50	79.55	-6.06	96.94
34	-1.25	78.29	-12.58	84.36
35	4.23	82.52	6.16	90.52
36	2.03	84.55	13.37	103.89
37	3.00	87.55	6.03	109.91
38	1.91	89.47	2.44	112.35
39	2.55	92.02	0.00	112.35
40	2.93	94.95	8.89	121.24
41	0.73	95.68	1.08	122.32
42	2.47	98.15	-0.62	121.70
43	-0.98	97.18	-7.13	114.57
44	-1.40	95.78	-10.01	104.56
45	1.02	96.80	-4.52	100.04
46	1.23	98.03	-16.92	83.11
47	1.44	99.46	9.89	93.00
48	2.18	101.64	12.63	105.64

Average market-adjusted returns (AR) and cumulative average returns (CAR), for a sample of parent firms and their spin-offs that went public from 1999 to 2003 until their 48-month anniversary. The market-adjusted return (ar_{it}) for stock i in event month t is defined as $ar_{it} = r_{it} - r_{iL,t}$ where r_{it} is the stock return for event firm i in month t and $r_{iL,t}$ is the value stock market index.
The average benchmark-adjusted return across stocks for event month t is the equal-weighted average of the adjusted returns for event month t for spin-offs $\bar{A}R_t = 1/N \sum_{i=1}^{NL}$ where N is the number of spin-off firms.

Table 3. Equally-Weighted Cumulative Adjusted Returns for Parent Firms and their Spin-offs

Months of	**Panel A**		**Panel B**	
Seasoning	**Spin-off AR %**	**Spin-off CAR %**	**Parents AR %**	**Parent CAR %**
1	15.53	15.53	-8.83	-8.83
2	14.37	29.90	-1.27	-10.11
3	0.44	30.34	-7.79	-17.90
4	9.97	40.31	0.00	-17.70
5	2.04	42.35	22.60	4.86
6	-5.44	36.92	6.46	11.32
7	7.93	44.85	9.97	21.30
8	3.21	48.05	3.47	24.78
9	-12.12	35.93	-4.02	20.75
10	-3.38	32.55	-16.33	4.41
11	8.41	40.96	-5.50	-1.07
12	10.46	51.42	1.24	0.17
13	5.06	56.48	13.80	13.97
14	4.17	60.65	0.07	13.22
15	14.50	75.15	-8.19	5.04
16	10.86	86.01	36.52	41.56
17	1.01	87.03	32.05	73.62
18	0.36	87.39	1.60	75.20
19	1.47	88.86	3.80	79.01
20	3.56	92.41	-3.99	75.02
21	7.01	99.42	15.82	90.84
22	5.88	105.30	12.25	103.95
23	3.84	109.14	10.74	113.83
24	1.96	111.10	5.54	119.38
25	-2.13	108.97	11.23	130.06
26	-7.72	101.24	39.85	170.47
27	6.92	108.16	27.03	197.49
28	11.70	119.86	-12.37	185.13
29	-2.53	117.34	0.00	185.25
30	7.68	125.02	10.10	195.32
31	-13.33	111.69	0.00	195.97
32	-3.67	108.02	10.80	206.76
33	-3.43	104.58	6.40	213.20
34	1.63	106.22	-20.90	192.20
35	13.15	119.37	-13.90	178.27
36	0.61	119.98	18.70	197.04
37	5.23	125.21	20.90	217.99
38	8.18	133.39	10.50	228.65
39	9.41	142.79	-1.20	227.30
40	8.14	150.93	15.00	242.33
41	-16.44	134.48	10.80	253.21
42	0.80	135.28	7.70	260.92
43	-7.29	127.99	0.00	261.07
44	0.90	128.89	-17.60	243.44
45	3.64	132.53	-3.50	239.90
46	9.72	142.25	-30.00	209.81
47	-1.68	140.58	-12.70	197.07
48	-2.74	137.83	28.60	225.64

Average market-adjusted returns (AR) and cumulative average returns (CAR), for a sample of parent firms and their spin-offs that went public from 1999 to 2003 until their 48-month anniversary. The market-adjusted return (ar_{it}) for stock i in event month t is defined as $ar_{it} = r_{it} - r_{iL,t}$ *where r_{it} is the stock return for event firm i in month t and $r_{iL,t}$ is the equally weighted market index. The average benchmark-adjusted return across stocks for event month t is the equal-weighted average of the adjusted returns for event month t for spin-offs*

$$\breve{A}R_t = 1/N \sum_{i=1}^{NL}$$ *where N is the number of spin-off firms.*

Table 3 – Panel B reports the average matching firm-adjusted returns (AR) and cumulative average market index returns (CAR) for the Parents of Spin-offs after the first 48 months after their offering date. Focusing first on the raw returns, a negative initial return of -8.83% is followed by monthly average raw returns varying between negative 30.00% and positive 39.85. The cumulative average raw return peaks at 243.44% in month 48. While the excess CAR returns remain regardless of market benchmark used, my results confirm a slight upward bias in the non-value weighted indexes first reported by Canina, Michaely, Thaler and Womack (1998). Thirty-one of the 48 monthly average adjusted returns are positive. The positive average adjusted returns increases to 39.85% by the end of month 48.

Table 4 - Panel A reports the four-year holding period returns for the parent firms of our spin-off sample. The median spin-off four-year return is 2.978% contrasted with .996% for the value weighted market index and .915% for the equally weighted market index. In other words, every dollar invested in a portfolio of spin-offs purchased at the closing market price on the first month of trading results in a terminal wealth of $1.03, while every dollar in the value-weighted market index would result in a terminal wealth of $.996 and a $.915 terminal wealth for the equally-weighted market index.

Table 4 - Panel B reports the four-year holding period returns for the parent firms of our spin-off sample. The median spin-off four-year return is 1.015% contrasted with .996% for the value weighted market index and .914% for the equally weighted market index. In other words, every dollar invested in a portfolio of spin-offs purchased at the closing market price on the first month of trading results in a terminal wealth of $1.01, while every dollar in the value-weighted market index would result in a terminal wealth of $.996 and a $.914 terminal wealth for the equally-weighted market index.

Table 4: Four-Year Holding Period Returns for Parent Firms and their Spin-offs and their Matched Market Returns

Market Index Type	*Panel A* Four-year Holding Period Returns, in %		Spin-off BHER	*Panel B* Four-year Holding Period Returns, in %		Parent BHER
	Spin-off Firms	**Market Index**		**Parent Firms**	**Market Index**	
	(1)	(2)	**(1) – (2)**	(1)	(2)	**(1) – (2)**
Value-weighted	2.978	0.996	**1.982**	1.015	0.996	**0.019**
Equally-weighted	2.978	0.915	**2.063**	1.015	0.915	**0.100**

Four-year holding period returns are calculated as [Π (1+ r_{it}) - 1] where *r* is the return on stock *i. The CRSP return tapes are the source of the monthly returns. To avoid survivorship bias, the total return is calculated until the delisting date for event firms that are delisted before the four-year anniversary. The corresponding matching index is calculated over the same period as that of the event firm. Buy and hold excess returns are calculated as event firm return minus the market return. Buy and hold excess returns are abbreviated in the table as BHER.*

The extant empirical literature provides several reasonable explanations for the excess returns documented in this study. For example, Allen, Lummer, McConnell and Reed (1995) theorize that the primary benefit derived from a spin-off is the removal of significant negative synergies between the parent and subsidiary. This is particularly apparent when two diverse business units are separated. Another explanation for the superior performance of spin-offs could be related to the documented positive announcement returns found for spin-offs that originate due to an acquisition and the subsequent streamlining of the parent firm. This explanation suggests that positive abnormal returns represent the re-creation of value that was destroyed at the time of the parent firm acquisition. My empirical results could also be explained by the informational symmetry explanation of spin-offs, first reported by Krishnaswami and Subramaniam (1999). Their study finds that the abnormal returns following the issuance of a spin-off is typically larger than other new initial public offerings because the spun-off firm has higher levels of information asymmetry. Similarly, Hakansson (1982) notes that if all information regarding the future prospects of the spun-off entity is not fully disclosed at the time of a spin-off, then those holding private information may have the opportunity to earn positive excess returns in stock transactions following the spin-off.

CONCLUSION

This paper documents a strategy of investing in spin-offs at the end of the first day of public trading and holding them until their four-year anniversary. A portfolio invested fully in spin-offs would have left an investor with $1.298 cents relative to just $.996 for each dollar invested in the value-weighted market portfolio and $.915 for each dollar invested in the equally-weighted market portfolio. While a portfolio invested fully in the parents of spin-offs would have left an investor with $1.015 cents relative to just $.915 for each dollar invested in the equally weighted stock market portfolio and $.996 for each dollar invested in the value-weighted market portfolio. These findings are surprising because the parents and their spin-offs excess returns are observed during the bear market experienced in 2001 and 2002.

The finding that spin-off offerings over perform, on average, implies that the costs of raising external equity capital is high for these firms. The high transaction costs of raising external equity capital are similar to that documented for initial public offerings (Muscarella and Vetsuypens, 1990). Consequently, the small growth companies that predominate among spun-off firms face a higher cost of equity capital than is true for more established firms (Ritter, 1991).

The paper checks the robustness of the cumulative average market-index adjusted returns (CARs) by performing sensitivities on both the value- and equally-weighted indices. However, one limitation of this research is the specificity of its conclusions is limited to the corporate environment that spurned the restructurings in the early nineties. An extension of the current study would examine other down market periods wherein pronounced spin-off restructurings were undertaken by the parent firms of spin-offs. A final extension of this work would involve using a time-series analysis to examine the after performance of spin-offs over longer time intervals (greater than 10 years).

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