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THE EFFECT OF EXCHANGE RATE RISK ON THE CONDITIONAL RELATIONSHIP BETWEEN BETA RISK AND RETURN IN INTERNATIONAL EQUITY MARKETS

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ABSTRACT

The paper examines the effect of exchange rate risk on the conditional relationship between beta risk and return in international equity markets from January 1978 through September 2004. We use an extension of the model introduced by Pettengill, Sundaran, and Mathur (PSM Model, 1995) and adapted by several authors afterwards. The empirical results show evidence in international markets that are compatible with the PSM model and some international studies addressing returns that are unhedged against exchange rate risk. However, when this risk is controlled and hedged with forward contracts, the conditional relationship between beta risk and return appears asymmetric and presents a lower beta risk premium than the one takes place under unhedged returns in up-market months. A main business implication of the findings follows: international equity market administrators and portfolio managers can defend themselves against exchange risk by using forward contracts, particularly in world stock market conditions similar to those discussed throughout the paper.

JEL: G12, G15

INTRODUCTION

The effect of exchange rate risk on stock market returns has become an important issue for several economic agents, especially those interested in diversifying risks, when needed and feasible, by investing across different world stock markets. Simple observation of international stock markets shows that many investors spend substantial resources to control the foreign exchange risk, which is associated with their international stock market investments, under the belief that it is a source of risk to be hedged away. Thus, hedging foreign exchange rate risk will be a valuable financial strategy if investors price such risk in international stock markets. However and surprisingly, typical methods for testing whether the exchange rate risk is priced in the world stock market do not address the possibility that such risk could be differently priced under up and down world stock market periods. Currency exposure could lead to different stock market risk premiums depending on previous world stock market conditions. For instance, investors' optimism or pessimism may result in asymmetric responses under such conditions. This situation could generate differences on their portfolio returns' expectations due to the potential impact that changes in foreign exchange rate may have on foreign stock markets. Therefore, if investors price differently currency exposure in such world stock market conditions then this result may lead them to use financial hedging strategies in order to 1) protect their returns against exchange rate exposure and, thus, 2) eventually prevent financial portfolio problems or going bust.

Our main hypothesis is that if exchange rate risk were priced in an international context then there would be a positive and higher beta risk premium in up-market months and a negative and higher beta risk premium in down-market months. The above applies when investors do not hedge their international portfolio against exchange rate risk as compared to the situation when investors hedge their portfolios by using forward contracts. We assume that hedging may reduce both market (systematic) and non-market risk. We follow a methodology based on an extension of Pettengill, Sundaran, and Mathur (PSM Model, 1995). They are one of the first authors in examining the stock market risk premiums under up and down market periods. However, they also ignore both the effects that exchange rate risk and hedging strategies may have on stock market returns under such market conditions. Thus, the above issue constitutes the main objective in this study.

We start by examining the period from 1973 to 2003. This period includes the Bretton Woods (1973) and the Jamaica Agreements (1976). These agreements established a set of rules for the international monetary system, where flexible exchange rates were acceptable to the IMF members, and central banks could intervene in the exchange markets in order to control unwarranted volatilities. Despite the regulations, however, this period has not been exempt of exchange rate volatility. To achieve the stated objective, we use two approaches in order to isolate the effect of exchange rate risk. We follow the methodology of Pettengill, Sundaran, and Mathur (1995) in the first approach in order to estimate an international conditional relationship between market beta risk and return without controlling exchange rate risk. This approach assumes an unhedged estimation of security returns. Then, we control the exchange rate risk in the second approach by using forward currency contracts.

We organize the remainder of the paper as follows. Section 2 briefly discusses a review of the literature on this issue. Section 3 presents the methodology. Section 4 shows the empirical results and finally Section 5 discusses the paper conclusions and implications.

LITERATURE REVIEW

Several studies have considered the effects of exchange rate risk on asset returns when examining international asset pricing models assuming both unhedged returns and exchange rate risk as a pervasive independent explanatory factor (Solnik, 1974; Sercu, 1980; Stulz; 1981; Adler and Dumas, 1983; Solnik, 1997). Yet, the empirical evidence is twofold. On the one side, the results from testing unconditional asset pricing models are not conclusive. Early studies (e.g., Hamao, 1988; Jorion, 1991) found no evidence in favor of the effects of exchange risk pricing on the Japanese or the US stock markets. More recent studies (e.g., Carrieri and Majerbi, 2006), however, show significant unconditional exchange risk premium in emerging stock markets. On the other side, the results from testing time-varying conditional asset pricing models generally support the hypothesis that foreign exchange risk is priced in the stock markets of major developed countries (Dumas and Solnik, 1995; De Santis and Gerard, 1998; Choi et al., 1998; Doukas et al., 1999; Carrieri, 2001). Previous studies, however, do not asses the issue that exchange risk premia may eventually differ under up and down stock market periods when this risk is controlled on asset returns (dependent variable) by using forward contracts. The traditional approaches based on conditional time-varying and unconditional models do not consider previous possibility.

In the capital asset pricing literature, is also possible to find a variety of studies that address the issue of up and down beta risk premia. Wiggins (1992), Bhardwaj and Brooks (1993) and Pettengill, Sundaram, and Mathur (1995) suggested a potential explanation for the flat unconditional relationship between beta risk and security returns in the U.S. equity market. More recently, there has been an increasing flow of empirical research validating a conditional relationship rather than an unconditional relationship for these two variables, beta risk and return (Cheng, 2005; Conover, Friday, and Howton, 2000; Faff, 2001; Howton and Peterson, 1998; Hung and Shackelton, 2004; Jensen and Mercer, 2002; Pettengill, Sundaran, and Mathur, 2002; Tang and Shum, 2003). The work of Pettengill et al., (1995) is one of the first in recognizing that a conditional relationship between market beta risk and return may take place in up and down stock market periods and that a systematic relationship must exist between market beta risk and return for the further to be a useful measure of risk. They assumed that both variables depend on whether the excess market return is positive or negative. In fact, using 55 years of U.S monthly stock return data, Pettengill, Sundaran, and Mathur (1995) show that beta market risk is priced when sample period is divided into up and down market months. Granted, the CAPM indicates a systematic and positive tradeoff between market beta and expected return. Yet, in line with rational expectations and Pettengill et al. (1995), there should be a positive relationship between realized returns and market beta during positive market-excess return periods and a negative relationship during negative market-excess return periods.

More specifically, let's assume that the expected return for the portfolio is the mean of the distribution for all possible returns for that portfolio in a given period. Thus, the expected value for its return distribution must contain a non-zero probability of realizing a return below the risk-free rate for market portfolio or for any other portfolio with a positive market beta. Otherwise, no investor would hold risk free bonds. In addition, portfolios with higher market betas have higher expected returns because of higher risks, that is, there must be some level of realized return for which the probability of exceeding that particular return is greater for the low market beta portfolio than for the high market beta portfolio. This logic can help understand why some investors carry on low-market beta portfolios. Indeed, returns for high market beta portfolios are less than returns for low market beta portfolios when the realized market return is less than the risk free rate. Therefore, the main Pettengill et al's prediction works out: there should be a positive relationship between realized returns and market beta during positive market-excess return periods and a negative relationship during negative market-excess return periods.

In an international setting, the implication of this conditional relationship has also been a matter of study. Empirical support was found for a significant positive relationship between beta and return in up-market months and a significant but negative relationship in down market months in U.K.'s main international developed stock markets (Fletcher, 1997, 2000), Japan's equity markets (Hodoshima, Garza-Gomez, and Kunimura, 2000), and Latin American equity markets (Sandoval and Saens, 2004).

The main drawback of previous studies, however, is that they do not take into account the effect that exchange rate risk may have on such conditional relationship after controlling the exchange rate risk effect on asset returns, as dependent variable, by using forward currency contracts. In an international setting, currency exposure may lead to different stock market risk premiums during up and down world stock market periods. As previously mentioned, investors' optimism or pessimism may result in asymmetric responses under such conditions. This behavior could imply differences on their portfolio returns' expectations given the potential effect that the foreign exchange rate volatility may have on foreign stock markets. The study of this issue is the focus of this article.

METHODOLOGY

Exchange Rate Volatility and Unhedged Security Returns

This section presents the analysis of the effect of exchange rate variability on the total risk of foreign stock markets. The analysis takes the viewpoint of a U.S. investor investing in the U.S. and sixteen major foreign stock markets, i.e., Australia, Austria, Belgium, Canada, Denmark, France, Germany, Italy, Japan, Netherlands, Norway, Singapore, Spain, Sweden, Switzerland and the UK. The international portfolio strategy initially assumes uncovered returns against exchange rate risk, and uses Eun and Resnick's (1988) methodology, which includes six years of weekly data and the experience of seven countries. This study uses 32 years of monthly data from 17 major countries.

The Effect of Exchange Rate Variability

The dollar rate of return, $\widetilde{R}_{jUS\$}$, that an U.S. investor can make from an unhedged investment in the jth foreign stock market is given by

$$\tilde{R}_{jUS\$} = (1+\tilde{R}_{j\$})(1+\tilde{e}_j) - 1 \tag{1}$$

$$\tilde{R}_{jUS\$} = \tilde{R}_{j\$} + \tilde{e}_j + \tilde{R}_{j\$}\tilde{e}_j \tag{2}$$

where $\tilde{R}_{j\$}$ is the local stock market return, which is measured in domestic currency, $\tilde{e}_j$ is the rate of appreciation of the domestic currency against the dollar, and the symbol "~" indicates a stochastic variable. Note that the cross-product, $\tilde{R}_{j\$}\tilde{e}_j$, in equation (2) is close to zero in value, so $\tilde{R}_{jUS\$}$ can be estimated by

$$\tilde{R}_{jUS\$} \approx \tilde{R}_{j\$} + \tilde{e}_j \tag{3}$$

From equation (3), the variance of the dollar rate of return can be estimated by

$$Var(\tilde{R}_{jus\$}) \approx Var(\tilde{R}_{j\$}) + 2Cov(\tilde{R}_{j\$}, \tilde{e}_j) + Var(\tilde{e}_j) \tag{4}$$

The previous analysis can be extended into a portfolio context. The variability of dollar portfolio returns, $Var(\tilde{R}_{pUS\$})$, can be estimated as

$$Var(\tilde{R}_{pUS\$}) = \sum_{j=1}^{n}\sum_{k=1}^{n} w_j w_k Cov(\tilde{R}_{jUS\$}, \tilde{R}_{kUS\$}) \;\forall\; \mathrm{j, k} = 1,\ldots\ldots\ldots\ldots, \mathrm{n} \tag{5}$$

where w_j shows the percentage of wealth invested in the *j*th stock market. It is noted from equation (3) that

$$\begin{aligned} Cov(\tilde{R}_{jUS\$}, \tilde{R}_{kUS\$}) &\approx Cov\left[(\tilde{R}_{j\$} + \tilde{e}_j), (\tilde{R}_{K\$} + \tilde{e}_K\right] \\ &\approx Cov(\tilde{R}_{j\$}, \tilde{R}_{k\$}) + Cov(\tilde{R}_{j\$}, \tilde{e}_k) + Cov(\tilde{R}_{k\$}, \tilde{e}_j) + Cov(\tilde{e}_j, \tilde{e}_k) \end{aligned} \tag{6}$$

By merging equation (6) into equation (5), the former can be estimated as follows

$$Var(\tilde{R}_{pUS\$}) \approx \sum_{j=1}^{n}\sum_{k=1}^{n} w_j w_k Cov(\tilde{R}_{j\$}, \tilde{R}_{k\$}) + 2\sum_{j=1}^{n}\sum_{k=1}^{n} w_j w_k Cov(\tilde{R}_{j\$}, \tilde{e}_k) + \sum_{j=1}^{n}\sum_{k=1}^{n} w_j w_k Cov(\tilde{e}_j, \tilde{e}_k) \tag{7}$$

Equation (7) shows that the total portfolio risk depends on (a) the covariances among the local stock market returns (local currency returns), (b) the cross-covariances among the local stock market returns and the exchange rate variations, and (c) the covariances among the exchange rate variations. The exchange rate variability contributes to the total portfolio risk through the second and third terms of equation (7). We can observe that if the second and the third terms are largely positive (negative), then the exchange rate variability will increase (decrease) the overall portfolio risk.

The effect of the exchange rate risk can be analyzed by forming an equally weighted portfolio from the seventeen stock markets under study. In order to study what has been the trend decomposition of the total portfolio risk, we examine 28 moving subperiods using monthly data from January 1973 to September 2004. Each subperiod includes 60 months (5 years). The first subperiod starts in January 1973 and ends in

December 1977. The second period covers from January 1974 through December 1978 and so on. The $Var(\widetilde{R}_{pUS\$})$ decomposition using equation (7) is shown in Table 1.

Table 1: Decomposition of Portfolio Risk (Absolute Contribution in Squared Percent)

Component/Period	73-77	74-78	75-79	76-80	77-81	78-82	79-83
A. $\sum_{j=1}^{n}\sum_{k=1}^{n} w_j w_k Cov(\tilde{R}_{j\$},\tilde{R}_{k\$})$	15.755	14.384	10.039	6.998	6.770	7.494	8.364
B. $2\sum_{j=1}^{n}\sum_{k=1}^{n} w_j w_k Cov(\tilde{R}_{j\$},\tilde{e}_k)$	2.536	0.869	1.317	1.801	1.502	2.974	3.163
C. $\sum_{j=1}^{n}\sum_{k=1}^{n} w_j w_k Cov(\tilde{e}_j,\tilde{e}_k)$	4.297	4.486	4.282	4.511	5.765	6.463	4.237
D. $Var(\tilde{R}_{pUS\$})$	22.588	19.739	15.638	13.310	14.037	16.931	15.764

Component/Period	80-84	81-85	82-86	83-87	84-88	85-89	86-90
A. $\sum_{j=1}^{n}\sum_{k=1}^{n} w_j w_k Cov(\tilde{R}_{j\$},\tilde{R}_{k\$})$	9.698	8.569	9.284	19.752	19.590	19.319	25.359
B. $2\sum_{j=1}^{n}\sum_{k=1}^{n} w_j w_k Cov(\tilde{R}_{j\$},\tilde{e}_k)$	2.289	0.306	-0.121	-5.876	-5.341	-5.803	-5.222
C. $\sum_{j=1}^{n}\sum_{k=1}^{n} w_j w_k Cov(\tilde{e}_j,\tilde{e}_k)$	4.480	5.847	6.538	6.564	7.307	7.418	5.792
D. $Var(\tilde{R}_{pUS\$})$	16.467	14.722	15.701	20.440	21.556	20.934	25.929

Component/Period	87-91	88-92	89-93	90-94	91-95	92-96	93-97
A. $\sum_{j=1}^{n}\sum_{k=1}^{n} w_j w_k Cov(\tilde{R}_{j\$},\tilde{R}_{k\$})$	25.087	16.180	17.356	17.454	11.635	10.356	12.044
B. $2\sum_{j=1}^{n}\sum_{k=1}^{n} w_j w_k Cov(\tilde{R}_{j\$},\tilde{e}_k)$	-6.187	-5.519	-5.780	-6.219	-5.867	-5.998	-5.075
C. $\sum_{j=1}^{n}\sum_{k=1}^{n} w_j w_k Cov(\tilde{e}_j,\tilde{e}_k)$	6.435	7.269	6.833	5.827	6.054	4.241	3.074
D. $Var(\tilde{R}_{pUS\$})$	25.335	17.930	18.409	17.062	11.822	8.599	10.043

Component/Period	94-98	95-99	96-00	97-01	98-02	99-03	00-04
A. $\sum_{j=1}^{n}\sum_{k=1}^{n} w_j w_k Cov(\tilde{R}_{j\$},\tilde{R}_{k\$})$	17.125	16.447	19.234	23.476	24.807	22.010	22.958
B. $2\sum_{j=1}^{n}\sum_{k=1}^{n} w_j w_k Cov(\tilde{R}_{j\$},\tilde{e}_k)$	-5.296	-4.183	-3.866	-2.038	0.932	1.456	0.737
C. $\sum_{j=1}^{n}\sum_{k=1}^{n} w_j w_k Cov(\tilde{e}_j,\tilde{e}_k)$	3.015	2.775	2.252	2.069	1.389	0.889	0.909
D. $Var(\tilde{R}_{pUS\$})$	14.844	15.039	17.620	23.507	27.128	24.355	24.604

Each entry in Table 1 shows the decomposition of total risk associated to an equally weighted portfolio according to its absolute contribution in squared percent. Row A shows the covariances among the local stock market returns (local currency returns). Row B shows cross-covariances among the local stock market returns and the exchange rate variations and row C shows the total portfolio risk measured by the variance of portfolio returns in US\$. The decomposition is presented for 28 sub-periods starting with the sub-period 1973-77 and finishing with the sub-period 2000-04.

Table 1 shows that the exchange rate risk (measured by the sum of component B and C) reinforces the portfolio risk during the first ten subperiods (from subperiod 1973-1977 to subperiod 1982-1986). During the first subperiod, exchange rate changes account for 19.0% of the risk of an equally weighted portfolio through its own covariance and for an extra 11.2% through its cross-covariances with the stock market returns. This trend is relatively stable up to subperiod 1982-1986 where exchange rate changes account for 41.7% of the risk of the portfolio through its own covariances and for a negative percentage (-0.8%) through its cross-covariances with the stock market returns. With no exchange rate risk, the portfolio variance would have been 15.755 and 9.284 squared percent for period 1973-1977 and 1982-1986, respectively, as compared to 22.588 and 15.701 squared percent, respectively, under the presence of exchange rate risk. This implies that, while the local stock market risk may be significantly diminished by constructing an international equally weighted portfolio, much of the exchange risk is nondiversifiable.

The results reported in Table 1 and Table 2 reveal important changes starting in period 1983-1987. It can be observed that components B and C almost offset each other in many cases during 18 subsequent subperiods. That is, the cross-covariances among the local stock market returns and the exchange rate variations behave with the opposite sign to the covariances among the exchange rate variations. Such trend implies that exchange rate risk did not have a significant effect on the total portfolio risk in many subperiods. The percentage of total portfolio risk explained by the cross-covariances among the local stock market returns ranges between 90.9% in the subperiod 1984-1988 and 120.4% in the subperiod 1992-1996. Surprisingly, however, exchange rate risk contributes to decrease the overall portfolio risk instead of increasing it in some subperiods as shown in Figure 1.

In summary, the subperiods between 1973 and 1986 are characterized by relatively higher exchange rate volatility where components B and C reinforce each other. Conversely, the subperiods between 1983 and 2004 show relatively lower exchange rate risk where components B and C offset each other. It should be noted that, important international economic agreements took place in order to stabilize the international monetary system during the subperiod 1983-2004. On the one hand, under the Plaza Agreement (New York, 1985) France, Japan, Germany the U.K. and the U.S. agreed that the dollar should depreciate against major currencies to solve a growing U.S. trade deficit, reason why they expressed their willingness to intervene in the exchange market to achieve this goal. On the other hand, under the Louvre Accord (Paris, 1987) Canada, France, Japan, Italy, Germany the U.K. and the U.S. agreed to cooperate in achieving greater exchange rate stability and to closely consult and coordinate their macroeconomic policies because of their concern that the dollar may fall too far. In fact, exchange rates have become more stable since the Louvre Accord.

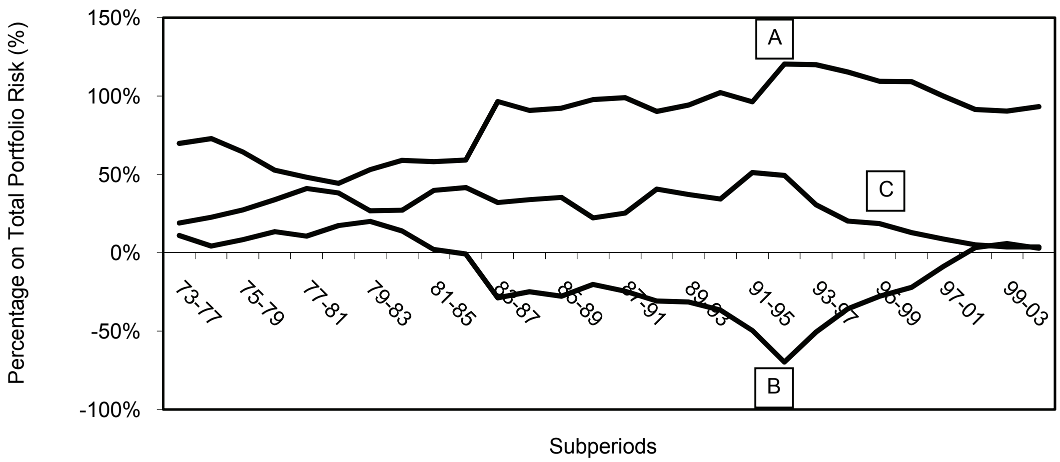
The previous analysis can serve as a good basis to formulate two hypotheses. 1) The exchange rate risk is a market risk from subperiods 1973-1977 to 1982-1986. Consequently, international equally weighted unhedged portfolios will have higher systematic risk premiums than those comparative hedging portfolio under the conditional relationship between beta risk and return applied by Pettengill, Sundaran, and Mathur (1995). 2) The exchange rate risk is neither a market nor a non-market risk from the subperiods 1983-1987 to 2000-2004. Consequently, there are not significant differences between hedging and not hedging portfolio returns against exchange rate risk. In practice, U.S. investors using forward contracts to control exchange risk in international equally weighted hedging portfolios, as implied in the previous hypothesis, will obtain no significant differences in terms of systematic risk and return compared to those comparative international equally weighted unhedged portfolio under the conditional relationship between beta risk and return applied by Pettengill, Sundaran, and Mathur (1995).

Table 2: Decomposition of Portfolio Risk (Relative Contribution in Percentage)

Component/Period	73-77	74-78	75-79	76-80	77-81	78-82	79-83
A. $\sum_{j=1}^{n}\sum_{k=1}^{n} w_j w_k Cov(\tilde{R}_{j\$},\tilde{R}_{k\$})$	69.8%	72.9%	64.2%	52.6%	48.2%	44.3%	53.1%
B. $2\sum_{j=1}^{n}\sum_{k=1}^{n} w_j w_k Cov(\tilde{R}_{j\$},\tilde{e}_k)$	11.2%	4.4%	8.4%	13.5%	10.7%	17.5%	20.1%
C. $\sum_{j=1}^{n}\sum_{k=1}^{n} w_j w_k Cov(\tilde{e}_j,\tilde{e}_k)$	19.0%	22.7%	27.4%	33.9%	41.1%	38.2%	26.8%
D. $Var(\tilde{R}_{pUS\$})$	100%	100%	100%	100%	100%	100%	100%
Component/Period	**80-84**	**81-85**	**82-86**	**83-87**	**84-88**	**85-89**	**86-90**
A. $\sum_{j=1}^{n}\sum_{k=1}^{n} w_j w_k Cov(\tilde{R}_{j\$},\tilde{R}_{k\$})$	58.9%	58.2%	59.1%	96.6%	90.9%	92.3%	97.8%
B. $2\sum_{j=1}^{n}\sum_{k=1}^{n} w_j w_k Cov(\tilde{R}_{j\$},\tilde{e}_k)$	13.9%	2.1%	-0.8%	-28.7%	-24.8%	-27.7%	-20.1%
C. $\sum_{j=1}^{n}\sum_{k=1}^{n} w_j w_k Cov(\tilde{e}_j,\tilde{e}_k)$	27.2%	39.7%	41.7%	32.1%	33.9%	35.4%	22.3%
D. $Var(\tilde{R}_{pUS\$})$	100%	100%	100%	100%	100%	100%	100%
Component/Period	**87-91**	**88-92**	**89-93**	**90-94**	**91-95**	**92-96**	**93-97**
A. $\sum_{j=1}^{n}\sum_{k=1}^{n} w_j w_k Cov(\tilde{R}_{j\$},\tilde{R}_{k\$})$	99.0%	90.2%	94.3%	102.3%	96.4%	120.4%	119.9%
B. $2\sum_{j=1}^{n}\sum_{k=1}^{n} w_j w_k Cov(\tilde{R}_{j\$},\tilde{e}_k)$	-24.4%	-30.8%	-31.4%	-36.4%	-49.6%	-69.8%	-50.5%
C. $\sum_{j=1}^{n}\sum_{k=1}^{n} w_j w_k Cov(\tilde{e}_j,\tilde{e}_k)$	25.4%	40.5%	37.1%	34.2%	51.2%	49.3%	30.6%
D. $Var(\tilde{R}_{pUS\$})$	100%	100%	100%	100%	100%	100%	100%
Component/Period	**94-98**	**95-99**	**96-00**	**97-01**	**98-02**	**99-03**	**00-04**
A. $\sum_{j=1}^{n}\sum_{k=1}^{n} w_j w_k Cov(\tilde{R}_{j\$},\tilde{R}_{k\$})$	115.4%	109.4%	109.2%	99.9%	91.4%	90.4%	93.3%
B. $2\sum_{j=1}^{n}\sum_{k=1}^{n} w_j w_k Cov(\tilde{R}_{j\$},\tilde{e}_k)$	-35.7%	-27.8%	-21.9%	-8.7%	3.4%	6.0%	3.0%
C. $\sum_{j=1}^{n}\sum_{k=1}^{n} w_j w_k Cov(\tilde{e}_j,\tilde{e}_k)$	20.3%	18.5%	12.8%	8.8%	5.1%	3.7%	3.7%
D. $Var(\tilde{R}_{pUS\$})$	100%	100%	100%	100%	100%	100%	100%

Each entry in Table 2 shows the decomposition of total risk associated to an equally weighted portfolio according to its relative contribution in percentage. Row A shows the covariances among the local stock market returns (local currency returns). Row B shows cross-covariances among the local stock market returns and the exchange rate variations and row C shows the total portfolio risk measured by the variance of portfolio returns in US$. The decomposition is presented for 28 sub-periods starting with the sub-period 1973-77 and finishing with the sub-period 2000-04.

Figure 1: Decomposition of Portfolio Risk (%)



Line A shows covariances among the local stock market returns (local currency returns) on total portfolio risk. Line B shows cross-covariances among the local stock market returns and the exchange rate variations on total portfolio risk. Line C shows covariances among the exchange rate variations on total portfolio risk

Exchange Rate Volatility and Hedging Security Returns

The previous section showed that exchange rate risk could be nondiversifiable at least in some subperiods within the period under study. Such observation opens the possibility to consider the use of forward contracts as a choice to control exchange rate risk. Suppose that a U.S. investor sells the expected foreign currency proceeds forward. At the maturity of the forward contract, the U.S. investor will exchange the uncertain dollar return $(1+E(\widetilde{R}_{j\$}))(1+\widetilde{e}_j)-1$ for the dollar return $(1+E(\widetilde{R}_{j\$}))(1+f_j)-1$, where $E(\widetilde{R}_{j\$})$ is the expected rate of return on the *j*th foreign stock market (in foreign currency) and f_j is the forward exchange premium, which is defined as $(F_j / S_j)-1$, where F_j and S_j are the forward and spot exchange rates in U.S. dollar equivalents, respectively. F_j is estimated following the International Interest Rate Parity (IIRP) as $F_j = S_j(1+r_{US})/(1+r_j)$, where r_{US} and r_j are the risk-free interest rates in the U.S. and country *j*, respectively. However, the unexpected foreign currency proceeds $(\widetilde{R}_{j\$}-E(\widetilde{R}_j))$ will be exchanged for U.S. dollar at an uncertain future spot exchange rate. Therefore, the dollar rate of return for the hedging strategy, $\widetilde{R}^H_{j\ US\$}$, is given by

$$\widetilde{R}^H_{j\ US\$} = \{[1+E(\widetilde{R}_{j\$})](1+f_j)+[\widetilde{R}_{j\$}-E(\widetilde{R}_{j\$})](1+\widetilde{e}_j)\}-1 \tag{8}$$

$$\widetilde{R}^H_{j\ US\$} = \widetilde{R}_{j\$}+f_j+\widetilde{R}_{j\$}\widetilde{e}_j+E(\widetilde{R}_{j\$})(f_j-\widetilde{e}_j) \tag{9}$$

Because the third and fourth terms of equation (9) are small in value, this equation can be estimated by

$$\widetilde{R}^H_{j\ US\$} \approx \widetilde{R}_{j\$}+f_j \tag{10}$$

Therefore, the variance of the dollar hedged portfolio returns, $Var(\widetilde{R}^H_{p\ US\$})$, can be estimated approximately as:

$$Var(\widetilde{R}^{H}_{p\ US\$}) \approx \sum_{j=1}^{n}\sum_{k=1}^{n} w_j w_k Cov(\widetilde{R}^{H}_{j\ US\$}, \widetilde{R}^{H}_{k\ US\$}) \approx \sum_{j=1}^{n}\sum_{k=1}^{n} w_j w_k Cov[(\widetilde{R}_{j\$} + f_j), (\widetilde{R}_{k\$} + f_k)]$$
$$\approx \sum_{j=1}^{n}\sum_{k=1}^{n} w_j w_k Cov(\widetilde{R}_{j\$}, \widetilde{R}_{k\$}) \qquad (11)$$

Equations (7) and (11) and the results presented in Table 1 and 2 show that the hedging strategy may result in a lower portfolio variance for subperiods from 1973-1977 to 1982-1986. That is because of presence of exchange risk (where components B and C reinforce each other) results in a higher risk under an international equally weighted unhedged portfolio, reason why this risk can be reduced by using a hedging strategy. The hedging strategy using forward contracts seems to be one of the cases where portfolio risk can be reduced without adversely affecting portfolio return because the forward exchange premium is an unbiased estimator of the future change of the exchange rate, i.e., $f_j \approx E(\widetilde{e}_j)$. However, it is not clear that the hedging strategy may result in a lower portfolio variance for subperiods from 1983-1987 to 2000-2004, mainly because the presence of exchange risk (where components B and C almost offset each other) does not affect significantly the international equally weighted undhedged portfolio risk. Therefore, similar results in terms of portfolio performance can be expected for these periods either for unhedged or hedging strategies.

Model Specification and Econometric Methodology

This section presents the model specification and the econometric methodology used to test the pricing models proposed. The model specification starts with the zero-beta CAPM used by Black (1972), which predicts that:

$$E(R_i) = \gamma_0 + \gamma_1 \beta_i \qquad (12)$$

where $E(R_i)$ is the expected return on portfolio i, $\beta_i = Cov(R_i, R_m)/Var(R_m)$ is the beta of portfolio i, γ_0 is the expected return on the portfolio which has a zero covariance with the market portfolio, and γ_1 is the expected risk premium of the market portfolio. Yet, in an international setting, a CAPM extension requires some extra assumptions, such as that the capital markets are integrated and the international interest rate and purchasing power parities hold.

The next step relates to the estimation of the unconditional relationship between beta risk and return (URBRR). To accomplish this, the next analysis is performed in order to know if the CAPM model tested by Fama and MacBeth (1973) exhibits a positive relationship between realized portfolio returns and betas. The econometric tests are conducted in two stages. The first stage requires the stock market index beta estimation. Due to the documented presence of infrequent trading in many stock markets, individual betas are estimated using the aggregated coefficients method proposed by Dimson (1979) under two scenarios affecting the seventeen stock markets under analysis. The first scenario relates to unhedged stock market returns, which are estimated using equation (1). The second scenario considers hedged stock market returns, which are estimated using equation (10). Thus, the unhedged and hedged betas are estimated for each stock market index over the 5 years (60 months) moving subperiods, starting with the subperiod from January 1973 to December 1977. The consistent estimates of individual betas for the first subperiod are then used as predictors for the subsequent subperiods up to subperiod 1999-2003. The last estimates of betas are used as predictors for year 2004.

In the second stage, a pooled cross-sectional OLS regression for equation (13) is estimated for subperiods from 1978 through 1986 and from 1987 through 2004, using the CAPM specified by Black (1972):

$$R_{jt} = \gamma_{0t} + \gamma_{1t}\beta_j + \mu_{jt} \tag{13}$$

where R_{jt} is the return on portfolio *j* in month *t*, β_j is the beta of portfolio *j*, and μ_{jt} is a random error term. Both, individual stock markets returns and betas from the previous stage (under the two scenarios, hedged and unhedged returns) are used in this stage to estimate portfolio returns and portfolio betas through equation (13). Thus, equation (13) gives estimates of the average values of monthly coefficients γ_{0t} and γ_{1t} for the tested periods, coefficients that are then tested to see if they are significantly different from zero. The main prediction obtained from equation (13) is whether β_j is the only cross-sectional variable that explains the relationship between portfolio returns and risk.

There is some evidence that additional factors may also contribute to explain cross-sectional return variations. Jegadeesh and Titman (1993), Banz (1981) and Fama and French (1992, 1996) conclude that momentum, size and book-to-market (B/M), respectively, can contribute to explain cross-sectional return variations in U.S. samples. Rouwenhorst (1999) finds general evidence for the three previous factors in emerging markets using univariate analysis. Marshall and Walker (2000) add the size-effect in the Chilean stock market. Yet, the methodologies used in previous studies omit controlling for the sign of the market premium. In contrast, Hodoshima, Garza-Gomez, and Kunimura (2000) and Sandoval and Saens (2004) use multivariate approaches in an international context and find evidence that factors (such as momentum, size and B/M) have little effect in explaining international cross-sectional return variations after controlling their effects on the conditional relationship between beta risk and return, as applied by Pettengill, Sundaran and Mathur (1995).

Furthermore, to test the Pettengill, Sundaran, and Mathur's (1995) conditional relationship between beta risk and return, the tested periods are split into up and down market months. If the realized market portfolio return is above the risk-free return (up market), portfolio betas and returns are positively related. However, if the realized market return is below the risk-free return (down market), portfolio betas and returns are inversely related. Consequently, regression coefficients for equation (14) are estimated in order to know if a systematic conditional relationship between beta and returns exists.

$$R_{jt} = \gamma_{0t} + \gamma_{2t} D\beta_j + \gamma_{3t}(1 - D)\beta_j + e_{it} \tag{14}$$

where $D = 1$ *if* $(R_{Mt} - R_{ft}) >= 0$, $D = 0$ *if* $(R_{Mt} - R_{ft}) < 0$. R_{Mt} is the market portfolio return, and R_{ft} is the risk-free rate, in week *t*. The derived hypotheses from this equation are: $H_0 : \gamma_2 = 0$ versus $H_A : \gamma_2 > 0$ and $H_0 : \gamma_3 = 0$ versus $H_A : \gamma_3 < 0$. Herein γ_2 and γ_3 are the average values of the coefficients γ_{2t} and γ_{3t}, respectively. The statistical significance of these coefficients can be tested using standard t-tests. It should be noted that Pettengill, Sundaran, and Mathur (1995) assume that the conditional relationship [equation (14)] does not imply a positive relationship between risk and return and that two conditions are necessary in order to test a positive relationship between risk and return. Those conditions are: (1) the excess market return are positive on average and (2) the beta risk premium in up markets and down markets are symmetrical. Because $\alpha_2 \leq 0$, the symmetry hypothesis can be specified as follows: $H_0 : \gamma_2 + \gamma_3 = 0$ versus $H_A : \gamma_2 + \gamma_3 \neq 0$. All the derived hypotheses can be tested using a Wald test, which accounts for an absolute significant difference between the γ_2 and the γ_3 coefficients.

<u>Data and Descriptive Statistics for Stock Market Indices</u>

The overall tested period begins in January 1978 and ends in September 2004. The following data sets are used: 1) Monthly returns in U.S. dollars for the seventeen equity indices under study from Morgan Stanley Capital International (MSCI). The seventeen equity indices are from Australia, Austria, Belgium, Canada, Denmark, France, Germany, Italy, Japan, Netherlands, Norway, Singapore, Spain, Sweden, Switzerland, the UK and USA. 2) Treasury bills and government bonds rates for the countries included in the sample from the Federal Reserve Bank of Chicago, and 3) World equity index from the International Monetary Fund (IMF) databases available in electronic sources. All stock markets returns are calculated in local currency and U.S. dollars. The monthly return on the MSCI world index is used as a proxy for the market portfolio. The monthly return on a 3-month U.S. Treasury Bill is used as the risk-free asset chosen due to data availability.

Table 3 sets forth MSCI summary statistics for the 17 stock markets and the world index under the scenario of unhedged returns for the period from January 1978 to September 2004. These stock markets constitute a significant part and, thus, are representative of the world stock market as a whole. The statistics shown include the average monthly return, the standard deviation and Dimson's average beta risk from the 28 moving subperiods examined. Dimson's betas are estimates with respect to the MSCI World index.

The average monthly returns range between 0.80% for Spain and 1.38% for Sweden. The standard deviations range between 4.17% for the World Index and 7.77% for Singapore. Notably, the World Index exhibits the smallest returns standard deviation across the stock markets, which is consistent with the assertion that potential benefits are derived from international diversification. The Dimson's average betas range between 0.498 for Austria and 1.306 for Singapore, country that also shows the highest total risk and the highest systematic risk.

Table 3: Summary Statistics – Unhedged Returns

Stock Market	Mean Return %	Standard Deviation %	Dimson's Average Beta
Australia	0.87%	6.68%	1.148
Austria	0.81%	6.48%	0.498
Belgium	0.86%	5.74%	0.696
Canada	0.81%	5.62%	1.067
Denmark	0.98%	5.48%	0.538
France	1.08%	6.55%	1.009
Germany	0.81%	6.41%	0.658
Italy	1.13%	7.41%	1.052
Japan	0.85%	6.73%	1.285
Netherlands	0.91%	5.23%	0.829
Norway	1.07%	7.57%	0.760
Singapore	0.94%	7.77%	1.306
Spain	0.80%	6.64%	0.974
Sweden	1.38%	7.15%	1.174
Switzerland	0.93%	5.29%	0.836
UK	0.86%	5.42%	1.026
USA	0.84%	4.42%	0.851
World	0.80%	4.17%	

Table 3 sets forth the mean return %, standard deviation % and Dimson's average beta for the 17 stock markets and the world index under the scenario of unhedged returns for the period from January 1978 to September 2004. Source: MSCI Website (http://www.msci.com)

Table 4 shows similar MSCI summary statistics for the 17 stock markets and the world index under the scenario of hedged returns for the same period (from January 1978 to September 2004). The results are similar although slightly lower for hedged returns compared to those observed in Table 3 for unhedged returns. The average monthly returns range between 0.68% for Spain and 1.34% for Sweden. The standard deviations range between 4.17% for the World Index and 7.41% for Italy. The Dimson's average betas range between 0.440 for Austria and 1.217 for Singapore. The lower indices indicate the probability of achieving a better systematic risk-return trade off when using a hedging strategy for an internationally diversified portfolio. This assertion is tested in the next section.

Table 4: Summary Statistics – Hedged Returns

Stock Market	Mean Return %	Standard Deviation %	Dimson's Average Beta
Australia	0.70%	5.46%	1.015
Austria	1.00%	6.35%	0.440
Belgium	0.80%	5.69%	0.632
Canada	0.69%	4.98%	0.965
Denmark	0.69%	5.28%	0.500
France	1.00%	6.25%	0.922
Germany	0.83%	6.18%	0.613
Italy	0.95%	7.41%	0.933
Japan	0.88%	5.38%	0.833
Netherlands	0.80%	5.38%	0.775
Norway	0.91%	7.27%	0.715
Singapore	1.09%	7.22%	1.217
Spain	0.68%	6.57%	0.917
Sweden	1.34%	7.15%	1.077
Switzerland	0.93%	4.80%	0.754
UK	0.66%	4.81%	0.848
USA	0.84%	4.42%	0.851
World	0.80%	4.17%	

Table 4 sets forth the mean return %, standard deviation % and Dimson's average beta for the 17 stock markets and the world index under the scenario of hedged returns for the period from January 1978 to September 2004. Source: MSCI Website (http://www.msci.com)

EMPIRICAL RESULTS

This section presents the research results from testing each proposed model: URBRR-UR, URBRR-HR, CRBRR-UR, and CRBRR-HR.

Unconditional Relationship between Beta Risk and Return Assuming Unhedged Returns (URBRR-UR)

The analysis focuses on the unconditional relationship between beta risk and return for the entire tested period and two subperiods under the scenario of unhedged returns. The subperiods are from January 1978 to December 1986 and from January 1987 to September 2004. Table 5 shows the average of the monthly coefficients of the intercept γ_{0t} and the slope γ_{1t} and their respective t statistics after running a pooled cross-sectional OLS regression. The latter is estimated for unhedged equity index returns for the 17 countries on a constant and their respective predicted betas according to equation (13). The country betas are estimated according to the aggregated coefficients method proposed by Dimson (1979). The t statistics (in parentheses) indicate whether the average value of the coefficient equals zero. The t statistics are corrected for heteroskedasticity and autocorrelation effects using Newey and West's (1987) method.

Table 5: Unconditional Relationship Assuming Unhedged Returns

Coefficient	Jan 1978 – Sept 2004	Jan 1978- Dec 1986	Jan 1987- Sept 2004
γ_0	0.00873 (4.12)***	0.00908 (2.81)***	0.01024 (3.62)***
γ_1	0.00071 (0.33)	0.00419 (1.28)	-0.00289 (-0.98)

*Table 5 shows the estimated parameters after running a pooled cross-sectional OLS regression for equation (13), which is estimated assuming unhedged returns for the entire period from 1978 through September 2004 and also for those subperiods indicated in the table. T-statistics are reported in parenthesis. ***, **, and * indicate significance at the 1, 5 and 10 percent levels respectively.*

The results reported in Table 5 evidence that there is no a significant relationship between beta risk and return in international unhedged equity returns for the entire tested period and the two subperiods. The annualized estimated beta risk premium is only 0.85%, which is statistically no significant. These results are in line with those published by Fama and French (1992) and subsequent studies that find a flat association between beta risk and portfolio return in the U.S. and other stock markets. One possible explanation for the flat relationship, as suggested by Pettengill, Sundaran, and Mathur (1995), is the inclusion of realized returns (past returns) in the tests rather than expected returns.

Unconditional Relationship between Beta Risk and Return Assuming Hedged Returns (URBRR-HR)

A similar analysis is performed for the unconditional relationship between beta risk and return for the same tested period and two subperiods under the scenario of hedged returns. The same method applied to produce Table 5 is used to generate data for Table 6. This table shows that there is a significant but negative relationship between beta risk and return in international hedged equity returns for the entire tested period. The annualized estimated beta risk premium is -6.72%, which is statistically significant. Yet, the results are different in each subperiod. There is no significant relationship between beta risk and hedged stock returns in the first subperiod, whereas there is a significant but negative relationship in the second subperiod. These results suggest that systematic risk is negatively priced when U.S. investors use forward contracts to hedge exchange rate risk. However, it can also be said that the tests are biased because they use realized returns (past returns) instead of expected returns.

Table 6: Unconditional Relationship Assuming Hedged Returns

Coefficient	Jan 1978 – Sept 2004	Jan 1978- Dec 1986	Jan 1987- Sept 2004
γ_0	0.01333 (6.99)***	0.01519 (6.40)***	0.01233 (4.01)***
γ_1	-0.00560 (-2.62)***	-0.00219 (-0.77)	-0.00700 (-2.11)**

*Table 6 shows the estimated parameters after running a pooled cross-sectional OLS regression for equation (13), which is estimated assuming hedged returns for the entire period from 1978 through September 2004 and also for those subperiods indicated in the table. T-statistics are reported in parenthesis. ***, **, and * indicate significance at the 1, 5 and 10 percent levels respectively.*

Conditional Relationship between Beta Risk and Return Assuming Unhedged Returns (CRBBR-UR)

A pooled cross sectional OLS regression is estimated for unhedged stock market index returns from 17 countries on a constant and their respective predicted betas under up and down markets according to equation (14) for the entire period and two subperiods. The betas for each country are estimated according to the aggregated coefficients method proposed by Dimson (1979). Table 7 includes the average of the monthly risk premiums in up market months γ_2 and in down market months γ_3. The Wald test is an indication of whether there is an absolute significant difference between γ_2 and γ_3 coefficients. The table also reports the included observations, the number of cross sections used, the total panel observations used in the tests, and the adjusted R^2. The t statistics (in parentheses) indicate whether the average value of γ_2 and γ_3 is significantly positive or negative. The t statistics have been corrected for heteroskedasticity and autocorrelation effects using Newey and West's (1987) method.

Results shown in Table 7 for the entire period are consistent with the results supporting the conditional relationship between beta risk and return. There exist a positive and significant relationship between return and beta risk in up market months and a significant but negative relationship in down market months. The annualized estimated beta risk premium is 30% in up market months and -38.1% in down market months, and both risk premiums result statistically significant at any significant level. In addition, the Wald test indicates that the hypothesis of a symmetric relationship between up market and down market months cannot be rejected at the 5% significance level.

The results for the two subperiods are consistent with those for the entire period in terms of a positive and significant relationship between return and beta risk in up market months and a significant but negative relationship in down market months. As Pettengill, Sundaran, and Mathur (1995) pointed out, the conditional relationship between beta risk and return implies that high beta risk equity countries like Singapore will show higher returns than low beta risk equity countries in up market months and lower returns in down market months. Yet, the hypothesis of a symmetric relationship between up and down market months is rejected in the second subperiod, because the evidence is less favorable in terms of symmetric risk premiums between up and down market months, even though there is support for a conditional relationship between beta risk and return for the entire period. These findings may be explained by the behavior of those investors who feel relatively pessimistic when facing business prospects under down world stock market months and, viceversa, optimistic when facing up world stock market months in the last subperiod. Thus, beta risk can be a useful indicator in international stock allocations, particularly when investors try to identify aggressive and defensive stock markets.

Table 7: Conditional Relationship Assuming Unhedged Returns

Coefficient	Jan 1978 – Sept 2004	Jan 1978- Dec 1986	Jan 1987- Sept 2004
γ_2	0.02500 (4.61)***	0.02565 (8.28)***	0.02288 ((7.94)***
γ_3	-0.03175 (-11.78)***	-0.02414 (-6.99)***	-0.03802 (-13.40)***
Wald test Ho: $\gamma_{2+}\gamma_3 = 0$	2.79187*	0.06439	7.86504***
p-value	0.09481	0.79971	0.00507
Included Observations	321 months	108 months	213 months
# of cross section used	17 countries	17 countries	17 countries
Total panel observations	5457 observations	1836 observations	3621 observations
Adjusted R^2	0.2021	0.1582	0.2315

*Table 7 shows the estimated parameters after running a pooled cross-sectional OLS regression for equation (14), which is estimated assuming unhedged returns for the entire period from 1978 through September 2004 and also for those subperiods indicated in the table. T-statistics are reported in parenthesis. ***, **, and * indicate significance at the 1, 5 and 10 percent levels respectively.*

Conditional Relationship between Beta Risk and Return Assuming Hedged Returns (CRBRR-HR)

A similar analysis is performed for the conditional relationship between beta risk and return assuming hedged returns. Results reported in Table 8 for the entire period are also consistent with the conditional model applied by Pettengill, Sundaran, and Mathur (1995). There is a positive and significant relationship between return and beta risk in up market months and a significant but negative relationship in down market months. The annualized estimated beta risk premium is 20.8% in up market months and -39.4% in down market months, and both risk premiums result statistically significant at any significant level. However, the Wald test indicates that the hypothesis of a symmetric relationship between up market and down market months is rejected at the 5% significance level. It should be noted that the beta risk premium is lower in up market months and slightly similar in down market months under hedged returns against exchange rate risk compared to similar indices under unhedged returns (compare Table 8 with Table 7). Previous results just apply to the case of up market months.

In addition, it should be noted that the above findings are mainly explained by the influences brought into play in the first subperiod. In this subperiod, the beta risk premium (16.7% annual) in up market months is significantly lower to the same index under unhedged returns (30.8% annual), which offers partial support to our first hypothesis (see Section 2). In other words, there is support for up but not for down market months at least in the first subperiod. The remaining beta risk premiums are similar in both subperiods, which offer support to our second hypothesis (see Section 2). These results suggest that the potential benefits from a hedging strategy against exchange rate risk can be captured mainly in the first subperiod under up market months. These findings are in line with the arguments discussed in the first part of the paper.

Table 8: Conditional Relationship Assuming Hedged Returns

Coefficient	Jan 1978 – Sept 2004	Jan 1978- Dec 1986	Jan 1987- Sept 2004
γ_2	0.01736 (8.21)*	0.0139 (4.87)*	0.01965 (6.06)*
γ_3	-0.03287 (-14.65)*	-0.02190 (-6.58)*	-0.03741 (-11.76)*
Wald test Ho: $\gamma_2 + \gamma_3 = 0$	14.88451***	2.19180	8.44535***
p-value	0.00012	0.13892	0.00368
Included Observations	321 months	108 months	213 months
# of cross section used	17 countries	17 countries	17 countries
Total panel observations	5457 observations	1836 observations	3621 observations
Adjusted R^2	0.1481	0.0787	0.1865

*Table 8 shows the estimated parameters after running a pooled cross-sectional OLS regression for equation (14), which is estimated assuming hedged returns for the entire period from 1978 through September 2004 and also for those subperiods indicated in the table. T-statistics are reported in parenthesis. ***, **, and * indicate significance at the 1, 5 and 10 percent levels respectively.*

CONCLUSIONS AND IMPLICATIONS

This study examines the conditional relationship between beta risk and return in international equity markets from January 1978 to September 2004 and focuses on the effect of exchange rate risk on this conditional relationship. Consistent with previous studies, the paper finds that there exist a flat unconditional relationship between beta risk and return when returns are unhedged against exchange rate risk. Conversely, there is a negative relationship when the returns are hedged using forward contracts. However, a note of caution is suggested in Pettengill, Sundaran, and Mathur (1995). These results can be biased when realized returns are used in the tests rather than expected returns.

When the conditional relationship between beta risk and returns in international equity markets is examined, there is a significant positive relationship in up market months and a significant but negative relationship in down market months. Under unhedged returns, this evidence is consistent with those reported in Pettengill, Sundaran, and Mathur (1995), Fletcher (1997, 2000), Hodoshima, Garza-Gomez, and Kunimura (2000) and Sandoval and Saens (2004) that provide support for the conditional relationship between beta risk and return in the U.S., U.K., main international developed stock markets, Japan, and Latin American equity markets, respectively. This research also finds that the conditional relationship between beta risk and return appears symmetric for the entire period under unhedged returns but that it turns asymmetric when the returns are hedged against exchange rate risk. These results suggest that, on average, U.S. investors feel more “optimistic” by hedging their internationally diversified portfolio instead of maintaining their international investments without exchange rate risk control.

However, it is important to note that these findings are mainly influenced by the market conditions that prevailed on the first subperiod (1978-1986), in which the beta risk premium in up market months is significantly lower than similar indices under unhedged returns. The remaining beta risk premiums are similar in both subperiods. Such observations have important implications for controlling exchange rate volatility in an international portfolio context. Potential benefits from a hedging strategy against exchange rate risk were captured mainly in the first subperiod and under up market months.

Furthermore, in the first period, the stock and foreign currency markets exhibited cross-covariances among the local stock market returns and the exchange rate variations that reinforce each other. This trend is reversed in the second subperiod (1987-2004). Thus, markets conditions as those prevailing in the first subperiod offer a good opportunity to control and, therefore, diminish exchange rate risk without adversely affecting portfolio return by using forward contracts.

Overall, the paper suggests that forward contracts are useful tools for controlling exchange rate risk when it takes the form of market risk. Forward contracts allow firms to reduce this risk when using exchange rates in their operations and, thus, can transfer this benefit to the stock markets. These contracts represent useful mechanisms to hedge this risk, and thus, it would explain why they have already been used attempting to obtain better payoffs for internationally managed portfolios. Thus, an implication for business policy practices follows: international equity market administrators and portfolio managers can defend themselves against exchange risk (as a market risk) by using forward contracts, particularly in world stock market conditions that are similar to those discussed throughout the paper.

REFERENCES

Adler, M., & Dumas, B (1983) "International Portfolio Choice and Corporation Finance: A Synthesis," *Journal of Finance*, vol. 38, p. 925-984

Banz, R., (1981) "The Relationship between Return and Market Value of Common Stocks," *Journal of Financial Economics*, vol. 9, p. 3-18

Bhardwaj, R. & Brooks, L (1993) "Dual Betas from Bull and Bear Markets: Reversal of the Size Effect," *Journal of Financial Research*, vol. 16(4), p. 269-283

Black, F., (1972) "Capital Market Equilibrium with Restricted Borrowing," *Journal of Business*, vol. 45, p. 444-455

Carrieri, F., (2001) "The Effects of Liberalization on Market and Currency Risk in the EU," *European Financial Management*, vol. 7, p. 259-290

Carrieri, F., & Majerbi, B (2006) "The Pricing of Exchange Risk in Emerging Stock Markets," *Journal of International Business Studies*, vol. 37, p. 372-391

Cheng, P., (2005) "Asymmetric Risk Measures and Real Estate Returns," *Journal of Real Estate Finance & Economics*, vol. 30(1), p. 89-102

Choi, J., Hiraki, T., and Takezama, N (1998) "Is Foreign Exchange Risk Priced in the Japanese Stock Market? *Journal of Financial and Quantitative Analysis*, vol. 33, p. 361-382

Conover, C., Friday, H., and Howton, S (2000) "An Analysis of the Cross-Section of Returns for REITs using a Varying-Risk Beta Model," *Real Estate Economics*, vol. 28, p. 141-163

De Santis, G., & Gérard, B (1998) "How Big is the Premium for Currency Risk," *Journal of Financial Economics*, vol. 49, p. 375-412

Dimson, E., (1979) "Risk Measurement when Shares are Subject to Infrequent Trading," *Journal of Financial Economics*, vol. 7, p. 197-226

Doukas, J., Hall, P., and Lang, L (1999) "The Pricing of Currency Risk in Japan," *Journal of Banking and Finance*, vol. 23, p. 1-20

Dumas, B., & Solnik, B (1995) "The World Price of Foreign Exchange Risk," *The Journal of Finance*, vol. 50, p. 445-479

Eun, C. & Resnick, B (1988) "Exchange Rate Uncertainty, Forward Contracts and International Portfolio Selection," *Journal of Finance*, vol. 43, p. 197-215

Faff, R., (2001) "A Multivariate Test of a Dual-Beta CAPM: Australian Evidence," *The Financial Review* vol. 36(4), p. 157-174

Fama, E. & MacBeth J (1973) "Risk, Return and Equilibrium," *Journal of Political Economy*, vol. 8, p. 607-636

Fama, E. & French, K (1992) "The Cross-Section of Expected Stock Returns," *The Journal of Finance*, vol. 47, p. 427-465

Fama, E. & French, K (1996) "Multifactor Explanations of Asset Pricing Anomalies," *The Journal of Finance*, vol. 51, p. 55-84

Fletcher, J., (1997) "An Examination of the Cross-Sectional Relationship of Beta and Returns: UK Evidence," *Journal of Economics and Business*, vol. 49, p. 211-221

Fletcher, J., (2000) "On the Conditional Relationship between Beta and Return in International Stock Returns," *International Review of Financial Analysis*, vol. 9, p. 235-245

Hamao, Y., (1988) "An Empirical Examination of Arbitrage Pricing Theory: Using Japanese Data," *Japan and the World Economy* vol. 1, p. 45-61

Hodoshima, J., Garza-Gómez, X., and Kunimura M (2000) "Cross-Sectional Regression Analysis of Return and Beta in Japan," *Journal of Economics and Business*, vol. 52, p. 515-533

Howton, S. & Peterson, D (1998) "An Examination of Cross-Sectional Realized Stock Returns Using Varying-Risk Beta Model," *The Financial Review*, vol. 33(3), p. 199-212

Hung, D., Shackelton, M., and Xu, X (2004) "CAPM, Higher Co-Moment and Factor Models of UK Stock Returns," *Journal of Business Finance & Accounting*, vol. 31(1), p. 87-112

Jegadeesh, N. & Titman, S (1993) "Returns to Buying Winners and Selling Losers: Implications for Stock Market Efficiency," *The Journal of Finance*, vol. 48, p. 65-91

Jensen, G. & Mercer, J (2002) "Monetary Policy and the Cross-Section of Expected Stock Returns," *Journal of Financial Research*, vol. 25(1), p. 125-139

Jorion, P., (1991) "The Pricing of Exchange Rate Risk in the Stock Market," *Journal of Financial and Quantitative Analysis*, vol. 26, p. 363-376

Marshall, P. & Walker, E (2000) "Day-of-the-Week and Size Effects in Emerging Markets: Evidence from Chile," *Revista de Análisis Económico*, vol. 15, p. 89-108

Pettengill, G., Sundaram, S., and Mathur, I (1995) "The Conditional Relation between Beta and Returns," *Journal of Financial and Quantitative Analysis*, vol. 30(1), p. 101-116

Pettengill, G., Sundaram, S., and Mathur, I (2002) "Payment for Risk: Constant Beta versus Dual-Beta Models" *The Financial Review*, vol. 37 (2), 123-135

Ross, S., (1976) "The Arbitrage Theory of Capital Asset Pricing," *Journal of Economic Theory*, vol. 13, p. 341-360

Rouwenhorst, K., (1999) "Local Return Factors and Turnover in Emerging Stock Markets," *The Journal of Finance*, vol. 54, p. 1439-1464

Sandoval, E. & Saens, R (2004) "The Conditional Relationship between Portfolio Beta and Return: Evidence from Latin America," *Cuadernos de Economía*, vol. 41, p. 65-89

Sercu, P., (1980) "A Generalization of the International Asset Pricing Model," *Journal de l'Association Francaise de Finance*, vol. 1, p. 91-135

Solnik, B., (1974) "The International Pricing of Risk: An Empirical Investigation of the World Capital Market Structure," *Journal of Finance,* vol. 29, p. 365-378

Solnik, B., (1997) "The World Price of Foreign Exchange Rate Risk: Some Synthetic Comments," *European Financial Management*, vol. 3, p. 9-22

Stulz, R.M., (1981) "A Model of International Asset Pricing," *Journal of Financial Economics*, vol. 9, p. 383-406

Tang, G. and Shum, W (2003) "The Relationship between Unsystematic Risk, Skewness and Stock Returns during up and down Markets," *International Business Review*, vol. 12(5), p. 523-541

Wiggins, J., (1992) "Betas in up and down Markets," *The Financial Review* vol. 27(1), p. 107-123

DO HERDING BEHAVIOR AND POSITIVE FEEDBACK EFFECTS INFLUENCE CAPITAL INFLOWS? EVIDENCE FROM ASIA AND LATIN AMERICA

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ABSTRACT

A considerable amount of research has focused on herding behavior vis-à-vis international capital, either by focusing on theory or by applying simple statistical analyses, but most studies have ignored factors that trigger international capital inflows. In essence, any connection between theory and empirical evidence has not been validated. In this paper, we test two primary drivers of capital inflows to emerging markets, namely herding behavior and positive feedback effects. Data from Asia and Latin America are used for our empirical study. There is significant evidence of positive feedback and herding behavior in both stable and highly volatile countries.

JEL: F21, G11, G15

INTRODUCTION

Global capital flows play a very important role in international integration and global financial management. A series of financial crises in Mexico, Asia, Russia, and Argentina, which revealed the volatility of international capital flows, especially to emerging markets, was blamed as a trigger. This motivated an academic literature on contagion effects (propagation of crises), "contagion," and the international transmission of crises, the "financial contagion," a phenomenon associated with dependencies among countries that allow market shocks in one country to affect other countries, often on a regional basis.

Most studies on capital flows emphasize the effects of changes at the macroeconomic level. Some studies, for example, have explored the effects of capital inflows on exchange rates and stock prices, while others have investigated their causes and other consequences.[1]

In the wake of both the 1994 Mexican and the 1997 Asian financial crises, there was a surge in the number of researchers attempted to find the root causes of these crises in a comprehensive and systematic manner. This was especially the case following the Asian crisis when interest had reached its peak.[2] A general consensus was reached: both financial crises were triggered by capital inflows followed by an abrupt change in confidence in large measure due to capital flights, a view later confirmed by the International Monetary Fund (IMF) in 1998. Interestingly enough, the cornerstone of most studies has been the impact of the shock brought on by the sudden sharp capital reversal in regional economies. This fact had actually been ignored in the early 1990s when substantial inflows of capital arrived through domestic financial channels (Radelet and Sachs, 1998, Rajan et al., 2003, Shen and Hsieh, 2000).

Although it is hard to determine whether co-movements are irrational or excessive, empirical work has been able to document patterns in the vulnerability of countries to volatility and to identify possible channels through which contagion is transmitted. Neither the exact cause of this volatility nor the best international financial architecture for guiding the movement of international capital is yet known. Yet reducing volatility and contagion has been an important stated objective of recent reforms (Dornbusch et al., 2000).

Therefore, analyzing the behavior of investors would become a very crucial issue. In the aftermath of the recent crises in emerging markets, considerable attention has focused on the question of whether herding behavior by international investors leads to excessive volatility in the flow of capital to developing countries.

Regarding the behavior of investors, the ‘positive feedback’ and herding behavior are usually analyzed. The ‘positive feedback’ (‘positive momentum’ or just ‘momentum’) trading strategy, refers to an investor buys past winners and sells past losers. Herding is the tendency for investors of a particular group to mimic each other’s trading.

Even though a considerable amount of research has focused on herding behavior vis-à-vis international capital, either by focusing on theory or by applying simple statistical analyses, most studies have ignored factors that trigger international capital inflows. In essence, any connection between theory and empirical evidence has not been validated.

The purpose of this paper intends to fill this gap. Moreover, the existing literature either focuses on individual economies or examines the determinants of FDI flows. This study adopts aggregate mutual fund data that belong to private capital flows rather than the more stable FDI flows and enables us to extend the sample period from 1996 to 2004 for selected countries in Asia and Latin America. How herding behavior and feedback effects affect capital inflows and how to analyze the causes of these behaviors will be investigated in this study.

The reason for choosing Asia and Latin America is that these two groups of countries account for over 75% of all international capital flows. Rajan et al. (2003) argued that South Asian countries paid higher interest rates before the crisis in order to attract international capital inflows, but their hypothesis has not been investigated for Asia as a whole. Additionally, we wish to examine whether there is a similar pattern in Latin American countries.

It is an interesting observation that a financial crisis in one country may very well cross borders and spread into other countries in the same region. Moreover, should the financial crisis not be harnessed, a regional crisis may even turn into a global one. Second, not only does this research study the behavior of international capital flows but also analyzes the causes of these behaviors. Third, previous studies have either adopted long period and highly accessible data or used high frequent data with short time span. This study takes the mutual fund data sets with the longest time frame possible, and in so doing, is able to capture a more robust analysis.

The remainder of this paper is organized as follows. Section 2 outlines the literature review. Section 3 provides the empirical model. Data description will be in section 4. Section 5 reports and analyzes the empirical results. The final section presents the conclusions.

LITERATURE REVIEW

International capital flows stem from both external (push) and internal (pull) factors. Calvo, Leiderman, and Reihart (1993) found that declines in U.S. interest rates were correlated with foreign reserves accumulation and real exchange rate appreciation that they used as proxies of capital inflows to Latin America in the early 1990s. They reported that external factors were the primary determinants of capital inflows to developing countries in that period. Based on their empirical study on country risk, Fernandez-Arias (1996) studied the determinants and sustainability of widespread capital inflows to middle-income countries after 1989 and confirmed that they were vulnerable to *external factors*, like international interest rate.

Taking a different approach from Fernandez-Arias' methodology, Berg and Pattillo (1998) subsequently found that *domestic factors* may have played a larger role in the 1997 East Asian currency crises. McKinnon (1999) concluded that higher interest premiums were the main reason for huge capital inflows to South Asia. McKinnon (1999) pointed out that these high interest differentials may have been associated with some kind of "Peso problem", i.e. a dramatic devaluation in Asia. Along similar lines, Rajan et al. (2003) recently found that countries in their South Asia group were paying higher interest rates than the London Inter-Bank Offered Rate (LIBOR). For example, in 1992, the interest rate differentials in South Asian countries were 19.83% in Indonesia, 5.11% in Malaysia, 15.28% in the Philippines and 7.97% in Thailand. These substantial capital inflows represented a very high percentage of the Gross Domestic Product (GDP) of Indonesia (8.3%), Malaysia (45.8%), the Philippines (23.1%) and Thailand (51.5%). Once these inflows abruptly reverse into capital flights, those countries suffered extremely grave economic consequences. McKinnon (1999), and Rajan et al. (2003) have reiterated that the main factor inducing capital inflows was high interest differentials in those countries before the crisis.

However, after the crisis -- from 1999 onwards, large sums of global mutual funds have been again injected into Asia, a phenomenon that has only received little attention until now. In another study, De Long et al. (1990) concluded that in the presence of noise traders, even rational investors may want to engage in positive feedback trading, and in the process, destabilize the market.

Much of the research has focused on Korea to shed light on questions relating to the trading strategy of investors, such as Choe et al. (1999), Kim and Wei (2002a) and Kim and Wei (2002b). Choe et al. (1999) find strong evidence of positive feedback trading and herding by foreign investors before the Korea's economic crisis. During the crisis period, herding falls, and positive feedback trading by foreign investors mostly disappears. Kim and Wei (2002a) find that although offshore funds trade more frequently, they do not, as a group, engage in positive feedback trading. Kim and Wei (2002b) find increased herding after the outbreak of the crisis. Froot et al. (2001) also get a similar result, that is, the factors that affect fund flow are based on the previous return, and the price sensitivity of regional stocks has a positive and massive impact on overseas fund inflow.

Instead of focusing on a particular country,[3] Borensztein and Gelos (2003) use a data set collected by Emerging Markets Funds Research, Inc., on the monthly geographic asset allocations of 467 funds active in developing countries over the period January 1996 to March 1999. They find that the degree of herding among funds is statistically significant, but moderate. Herding is not more prevalent during crises than during tranquil times. Funds tend to follow momentum strategies, selling past losers and buying past winners. The authors also present some evidence that suggests that increased herding measures are associated with higher stock return volatility but caution against pushing this conclusion too far.

But Bikhchandani and Sharma (2001) maintained that the herding behavior results from an obvious tendency by investors to copy the behavior of others. Fundamentals-driven spurious herding could arise, for example, if interest rates suddenly rise and stocks become less attractive investments. It can also be argued that herding takes place when investors are reacting to known public information -- that is, a rise in interest rates. Thus, the definitions of positive feedback and herding behavior should not be lumped together. As Bikhchandani and Sharma (2001) noted, however, much of the previous work does not test the validity of specific models of the causes of herding behavior. Also, in Bikhchandani and Sharma (2001), one cannot distinguish among the different causes of herding behavior directly from the analysis of a data set on asset holdings and price changes, as it is difficult, if not impossible, to discern the motive behind trade that is not driven by "fundamentals".

THE EMPIRICAL MODEL

In order to examine how positive effect and herding behavior affect capital flows and the causes behind them, this study constructs three models. Equation (1) intends to find whether the positive feedback effect does exist, in which an investor buys past winners and sells past losers.

$$Kl_{i,t} = \alpha_0 + \beta_{i,t-k} Stock_{i,t-k} + \gamma_{i,t-k} Exch_{i,t-k} + \varepsilon_{i,t} \qquad (1)$$

KI denotes change in monthly capital flow. *Stock* stands for monthly stock market returns of a given country. A positive coefficient of *Stock* implies that the level of capital inflows tends to move in the same direction with stock return, and the positive feedback effect is confirmed. *Exch* denotes variations in the exchange rate. The exchange rate is quoted in direct quotation, that is, a positive value for *Exch* represents a depreciation of the currency.

Furthermore, we determine whether "herding behavior" takes place. As mentioned in Bikhchandani and Sharma (2001), herding behavior is an obvious intent by investors to copy the behavior of others. Therefore, we construct a feasible model, an autocorrelation model modified from model (1) and represented as model (2).

$$Kl_{i,t} = \alpha_0 + \alpha_{i,t-k} KI_{i,t-k} + \beta_{i,t-k} StockR_{i,t-k} + \gamma_{i,t-k} Exch_{i,t-k} + \varepsilon_{i,t} \qquad (2)$$

A positive coefficient of *KI* means that previous high (low) capital flows subsequently induced even higher (lower) flows, indicative of typical herding behavior.

Here, it is important to rule out fundamentals-driven spurious herding, which means capital flows are driven by fundamental issues rather than investors copying the previous strategies of others. This is achieved by adding the control variables, interest rate, real GDP growth rate and foreign reserves and is written as:

$$\begin{aligned} KI_{i,t} &= \alpha_0 + \alpha_{i,t-k} KI_{i,t-k} + \beta_{i,t-k} Stock_{i,t-k} + \gamma_{i,t-k} Exch_{i,t-k} + \omega_{i,t-k} INT_{i,t-k} \\ &+ \rho_{i,t-k} RGDP_{i,t-k} + \lambda_{i,t-k} FR/IMP_{i,t-k} + \theta_{i,t-k} M2/FR_{t-k} + \varepsilon_{i,t} \qquad (3) \end{aligned}$$

Different from model (2), KI is weighted by the GDP of the specific country. In order to be consistent with the data frequency of the control variables, we adopt quarterly data for model (3). Hence, monthly data is adopted both in model (1) and model (2); but in model (3) turns to be quarterly data. *INT* denotes the interest rate differentials of LIBOR U.S. dollar lending. $RGDP$ refers the real GDP growth rate. FR/IMP, foreign reserves divided by value of imports, is a measure of the strength of reserve. $M2/FR$, broad money supply divided by foreign reserves, is a measure of financial liberalization.[4]

In order to account for a possible two-way causality between capital inflow and stock market returns, we perform the augmented Granger causality tests for Equations (2) and (3). Since the only difference between Equations (1) and (2) is the past values of capital inflow, a test for Equation (1) is not needed. Theoretically, when $Z_{t-k}, K = -1,0,1,...,T,$ do not affect y_t, then

$$E(y_t / z_{t-k,} y_{t-l}) = E(y_t / y_{t-l}),\ k = -1,0,1,...T;\ l = 1,2,...T.$$

We regress the *STOCK* on the future, present and past values of *KI*, all other variables, and *STOCK* own lags. Hence, the estimated equation without control variables is

$$STOCK_{i,t} = \alpha_0 + \alpha_{i,t} KI_{i,t-k} + \gamma_{i,t-j} Exch_{i,t-j} + \beta_{i,t-l} Stock_{i,t-l} + \varepsilon_{i,t},$$
$$k = -1, 0, 1, 2, 3;\ j = 0, 1, 2, 3;\ l = 1, 2, 3, \tag{4}$$

and with control variables:

$$STOCK_{i,t} = \alpha_0 + \alpha_{i,t} KI_{i,t-k} + \gamma_{i,t-j} Exch_{i,t-j} + \beta_{i,t-l} Stock_{i,t-l} + \omega_{i,t-j} INT_{i,t-j}$$
$$+ \rho_{i,t-j} RGDP + \lambda_{i,t-j} RFR / IMP_{i,t-j} + \theta_{i,t-j} M2 / FR_{i,t-j} + \varepsilon_{i,t},$$
$$k = -1, 0, 1, 2, 3;\ j = 0, 1, 2, 3;\ l = 1, 2, 3\,(5) \tag{5}$$

If coefficient of $KI_{i,t-k}$ is not statistically significant, then a two-way causality between capital inflows and stock returns is not a concern for our subsequent estimations.

DATA DESCRIPTION

Sample

We use the *Emerging Portfolio Fund Research* (*EPFR*) Global mutual fund database, a leading provider of mutual fund data, research and consulting. The EPFR tracks 10,000 international mutual funds in emerging and the U.S. markets with funds worth $5 trillion in assets, including offshore and U.S.-registered funds. The sample covers the January 1996-October 2004 period. In comparison with previous studies, this time frame provides the most comprehensive data set to our best knowledge. The data set covers selected emerging markets in Asia and Latin America. The Asian group consists of 12 countries, namely Bangladesh, China, Hong Kong, India, Indonesia, Korea, Malaysia, the Philippines, Singapore, Sri Lanka, Taiwan and Thailand. The Latin American group is made up of 7 countries, i.e., Argentina, Brazil, Chile, Colombia, Mexico, Peru and Venezuela. As mentioned earlier, these two groups occupy more than 75% of all international capital flows.

Table 1 presents the descriptive statistics of the variables for Asia and Latin America. It shows that capital flows into Asia is larger than those in Latin America, though the latter yields a higher stock return and interest rate differential. This implies that mutual fund managers have preferences for the relative steady economic environment (RGDP) and a sounder financial environment (M2/FR).

Worthy of mention is that the Korean economy was significantly hit by the Asian financial crisis and led to a nearly 50% decline of real GDP growth rate in the first quarter of year 1998. A similar situation happened in Indonesia, where a significant devaluation of Indonesian peso occurs to the maximum value of 135.50%. In the same period across the Pacific Ocean, Latin America tried to attract international capital to stay through raising interest rate. For example, the lending rate in Peru reached a high level of 424% in the second quarter of 1997.

The Capital Flows Volatility

This study adopts the volatility model, championed by Parkinson, German and Klass (1980) and Kuo and Chi (2000). We find that the volatility threshold of Asia is 7.01 and that of Latin America is 12.11. If the volatility level of one country is higher than its threshold value, it is deemed a volatile country; otherwise, it is considered stable.

In Asia, the volatile countries are China, Hong Kong, Korea, Singapore and Taiwan, and in Latin America, they are Argentina, Brazil and Mexico. According to Table 2**,** these volatile countries share some 77% of all mutual funds in Asia, and an even higher, 83%, in Latin America.

Table 1: Descriptive Statistics

Region	Variable	Mean	Stdev	Max	Min
Asia	KI (million)	2485	2869	13883	0
	Stock Return rate (%)	1.49	16.54	112.63	-48.86
	Interest Rate Differential (%)	6.67	4.90	29.98	-0.82
	Change Rate of Exchange (%)	1.47	8.59	135.50	-35.45
	Real GDP (%)	-3.10	11.16	25.04	-75.86
	FR/Import (quarters)	1.93	1.06	5.15	0.29
	M2/FR	5.36	13.35	58.33	0.00
Latin American	KI (million)	1517	2129	9046	-7299
	Stock Return rate (%)	4.53	20.93	201.55	-64.25
	Interest Rate Differential (%)	42.21	71.69	424.96	0.00
	Change Rate of Exchange (%)	2.88	9.77	91.00	-14.60
	Real GDP (%)	-32.77	328.94	3669.98	-2448.58
	FR/Import (quarters)	3.00	1.15	6.67	0.71
	M2/FR	1.34	1.35	4.77	0.00

KI denotes capital flows and is collected from EPFR. Stock return rate (%) = [(Stock Index)$_t$ -(Stock Index)$_{t-1}$] / (Stock Index)$_{t-1}$ X 100. Interest Rate Differential is defined as the domestic lending rate minus LIBOR US dollar lending rate. Change rate of exchange (%) = [(Exchange)$_t$ -(Exchange)$_{t-1}$] / (Exchange)$_{t-1}$ X 100. Real GDP stands for the real GDP growth rate and is defined as the nominal GDP growth rate minus CPI rate. FR/Import denotes the foreign reserve divided by the import values on the quarter basis M2/FR denotes the broad money divided by the foreign reserve on the quarter basis.

Interest Differentials, Exchange Rates and Stock Indices

Table 3 presents the interest differentials of the Asian and Latin American countries from 1996 to 2004. In Asia, the highest average interest differential was for Indonesia at 16.8%, more than double that of the Philippines (8.33%) and triples that of Thailand (5.24%). By way of comparison, the Latin American countries had considerably higher average interest differentials, most notably Brazil at 64.28%, and all other countries above 10% with the exception of Chile. On the basis of this finding, it can be argued that interest rate differential is one of the major pulling forces that attract investors.

Thus, it may be considered a key financial indicator linked to capital flows that investors watch over time so that they can make sound investment decisions. This finding also amplifies the point that, before their respective crisis, it was more profitable to invest in most of the selected Latin American countries than in the Asian ones.

It is worth pointing out that crucial monetary reform measures were put into place in 1994 to tighten economic policy and restrain consumption while attempting to fight against inflation. The interest rate differentials reached 7.74% from 6.30% in 1993 and 1994. A series of macroeconomic reforms reduced the interest rate before the crisis. This may have been one of the reasons that China was not a part of the Asian crisis.

Although the interest rate differentials of China did not rise after the 1997 crisis and financial institutions were not independent, bank credit expanded rapidly, culminating in an unsound credit system. With this state of affairs, the potential risk for the Chinese banking system was progressively increasing.

Table 2: Descriptive Statistics of Global Mutual Funds Monthly Inflows to Asian and Latin American Countries (1996/01~2004/11) (%)

Countries	Mean	Maximum	Minimum	Volatility
Asia				
Bangladesh	0.01	0.05	0.00	0.05
China	**6.08**	10.64	1.16	**9.48**
Hong Kong	**25.05**	35.45	16.72	**18.73**
India	3.74	5.19	2.77	2.42
Indonesia	2.32	3.71	1.19	2.52
Korea	**20.06**	28.65	3.76	**24.89**
Malaysia	5.20	8.34	2.43	5.91
Philippines	1.55	4.45	0.57	3.88
Singapore	**11.15**	17.26	8.01	**9.25**
Sri Lanka	0.09	0.25	0.04	0.21
Taiwan	**14.92**	19.65	10.29	**9.36**
Thailand	4.55	6.94	2.82	4.12
Average	7.29	10.83	3.83	**7.01**
Latin America				
Argentina	**6.18**	14.07	0.46	**13.61**
Brazil	**40.33**	50.97	27.43	**23.54**
Chile	7.89	11.34	5.61	5.73
Colombia	0.86	2.75	0.01	2.74
Mexico	**36.54**	52.29	22.22	**30.07**
Peru	2.32	4.80	0.80	4.00
Venezuela	1.55	5.27	0.20	5.07
Average	13.67	20.21	8.10	**12.11**

Volatility was first proposed by Parkinson, German and Klass (1980) and Kuao and Chi (2000) and is defined as (maximum – minimum) / N. The threshold of volatility is 7.01 for Asia and 12.11 for Latin America. In Asia, the volatile countries are China, Hong Kong, Korea, Singapore and Taiwan, and in Latin America, are Argentina, Brazil and Mexico.

Table 3: Interest Rate Differentials in Asian and Latin American Countries (%)

	1996	1997	1998	1999	2000	2001	2002	2003	2004	Average
Asia										
China	4.64	2.98	0.89	0.49	-0.63	2.12	3.55	4.14	4	2.46
Hong Kong	3.06	3.84	3.50	3.14	3.02	1.4	3.24	3.83	3.42	3.16
Indonesia	13.78	**16.16**	**26.65**	22.31	11.97	14.82	17.19	15.76	12.54	**16.80**
Korea	3.4	**6.22**	**9.78**	4.04	2.06	3.98	5.01	5.07	4.32	**4.88**
Malaysia	4.5	**4.97**	**6.63**	3.21	1.19	3.4	4.77	5.13	4.47	**4.25**
Philippines	9.4	**10.62**	**11.28**	6.42	4.42	8.67	7.38	8.3	8.5	**8.33**
Singapore	0.82	0.66	1.94	0.44	-0.65	1.93	3.61	4.14	3.72	1.85
Taiwan	3.11	7.99	8.92	2.67	1.12	3.26	3.77	2.93	1.89	**3.96**
Thailand	7.96	**7.99**	**8.92**	3.62	1.35	3.52	5.12	4.77	3.92	**5.24**
Latin America										
Argentina	5.07	3.58	5.14	4.56	7.36	25.95	50.51	17.57	2.38	13.57
Brazil	na	72.53	80.86	73.96	53.1	55.86	61.71	65.5	50.68	**64.28**
Chile	11.93	10.01	14.67	6.14	11.11	10.13	6.59	4.6	0.73	8.43
Colombia	36.55	28.56	36.74	19.29	15.06	18.96	15.16	13.61	10.68	**21.62**
Mexico	30.95	16.48	20.86	17.26	13.2	11.04	7.03	5.33	2.82	13.89
Peru	20.63	24.3	25.3	24.31	24.18	18.67	13.56	12.63	10.09	**19.30**
Venezuela	33.97	18.03	40.85	25.65	21.47	20.69	35.41	23.61	14.1	**25.98**

The Interest Rate Differential is the spread between the lending rate of a country (IFS line 60p) and the London Eurodollar lending rate (IFS line 60D). The data for Taiwan are retrieved from the Taiwan Economic Journal. The numbers in bold indicate that the value is higher than the level during the respective crisis.

Table 4 presents the results of the exchange rate variations, where a significant variation stands for a large devaluation during the sample period. Based on the exchange rate, the countries that experienced a financial crisis were those that had a large devaluation of currency between 1996 and 2004. These were

Indonesia, Korea, Malaysia, the Philippines and Thailand and most Latin America countries except Peru.

Table 4: The Exchange Rate Variations in Asian and Latin American Countries

	Exchange rate of local currency		Exchange rate of local currency against		Exchange volatility	
	End	Average	End	Average	End	Average
Asia						
China	8.30	8.31	8.28	8.28	-0.24%	-0.36%
Hong Kong	7.74	7.73	7.78	7.79	0.51%	0.77%
Indonesia	2377.75	2342.30	9253.80	8938.85	289.3%	281.7%
Korea	839.02	804.45	1050.80	1145.32	25.1%	42.3%
Malaysia	839.02	804.45	1050.80	1145.32	25.1%	42.3%
Philippines	26.29	26.22	56.18	56.04	113.3%	113.0%
Singapore	1.40	1.41	1.64	1.69	17.3%	19.8%
Taiwan	27.50	27.46	32.23	33.43	17.0%	21.1%
Thailand	25.55	25.34	39.18	40.22	53.6%	58.1%
Latin America						
Argentina	1	1	2.96	2.92	196.0%	192.0%
Brazil	1.04	1.01	2.65	2.93	154.8%	190.9%
Chile	424.97	412.27	559.83	609.37	31.4%	47.8%
Colombia	1005.33	1036.69	2412.1	2628.61	139.1%	153.8%
Mexico	7.85	7.6	11.26	11.29	43.9%	48.3%
Peru	2.6	2.45	3.28	3.41	26.4%	39.4%
Venezuela	476.5	417.33	1918	1891.33	302.8%	353.8%

The data for exchange rate are retrieved from the IMF's IFS, the November 2005 version, Line af. The positive numbers shown in the volatility column represent currency devaluation.

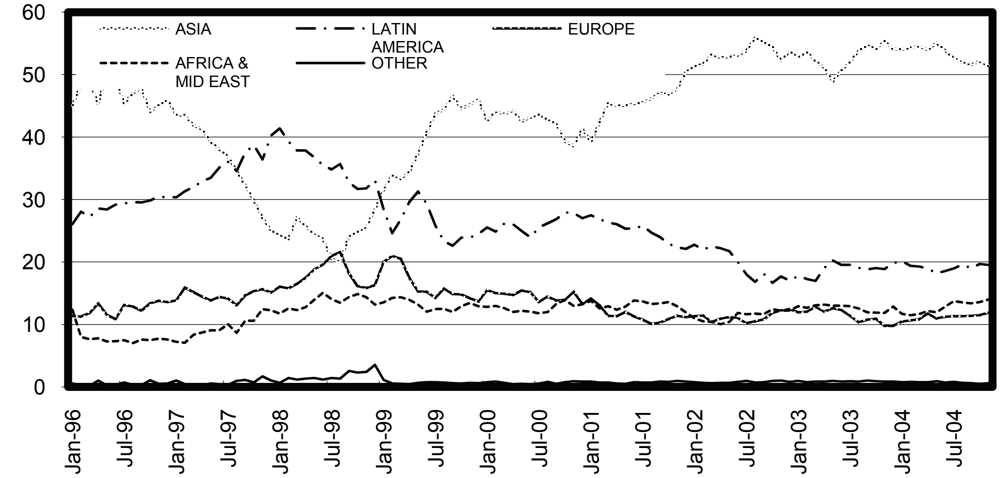
There is no question that the Philippines were the most volatile with a devalued position of 113%. This was followed by depreciation in Thailand (58%) and Singapore (42%). While all the other Asian countries were in a stage of depreciation, China was experiencing the opposite effect. In the Latin American countries, notably Argentina and Brazil, devaluation was much higher than in the Philippines, reaching 190%. In order to avoid the devaluation of invested currency, international capitals were transferred from the Latin American countries to the Asian region during 1999-2004. Figure 1 shows this trend while Fig 2 depicts the trend between capital inflows and the stock indices in Asia and Latin America. Clearly, capital inflows to Asia and Latin America follow a co-integrated pattern with that of the stock markets, but the pattern is more pronounced in Asian than in Latin American countries.

In Asia, this is more apparent in the case of Hong Kong, where capital inflows were also pegged to the Heng Seng Index. By contrast, in Brazil, the pattern is not as significant as that in Hong Kong. On the whole, it is evident that stock markets do have a certain influence on capital flows. This is in line with the theoretical argument that increased stock returns inevitably induce capital inflows. Simply put, the greater stock returns are, the greater are capital inflows.

EMPIRICAL FINDINGS

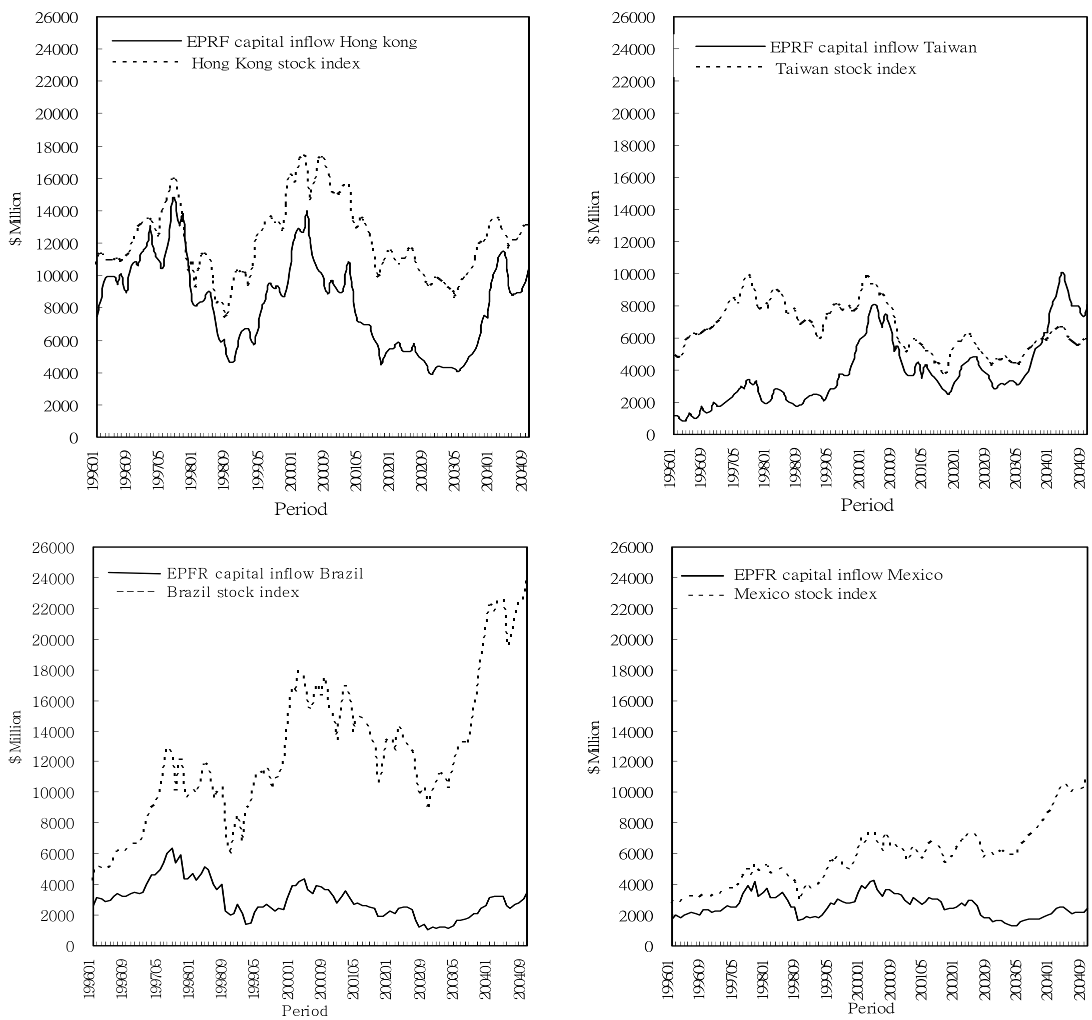
Tables 5 and 6 provide the results of the random effects regressions on models (1) and (2) for the positive feedback effect and the herding behavior effect, respectively. Table 5 reports the impacts of geographic location. The coefficients of the stock market returns (*Stock*) in different regions are all consistently positive for Asia and Latin America. Especially in Asia, the coefficients remain positive even in the two lagged periods of the stock market returns, implying the existence of a positive feedback effect.

Figure 1: Percentage of Global Mutual Funds invested in Different Regions (1996-2004)



Source: The Emerging Portfolio Fund Research (EPFR) Global mutual fund database

Figure 2: Comparison of the *EPFR* Capital Flows with the Stock Index of Selected Countries in Asia (Hong Kong and Taiwan) and Latin America (Mexico and Brazil)



A positive coefficient for *Stock* means that when stock returns increase (decrease), the level of capital inflows also increases (decreases), which matches the definition of the herding behavior effect is reported in Table 5: the coefficient of capital inflow (*KI*) with one lagged period for Asia is significantly positive, but for Latin America, they are significantly negative. Thus, managers with mutual funds invested in Asiamost probably intend to copy the behavior of other investors, an indication of herding behavior. It

Table 5: The Estimated Results for Models (1) and (2)-- By Geographic Location

	Asia		Latin America			Asia		Latin America	
	A	b	a	b		a	b	a	b
Stock	5.0083***	5.0790***	4.6444***	4.6973***	KI_{t-1}	0.1692***	0.1682***	0.0489	-0.1632*
	(6.008)	(6.081)	(8.948)	(9.003)		(5.827)	(5.793)	(1.315)	(-1.783)
$Stock_{t-1}$	6.9622***	7.006***	0.4417	0.4934	KI_{t-2}	0.0264	0.0276	-0.0231	0.0791
	(8.297)	(8.320)	(0.861)	(0.956)		(0.909)	(0.949)	(-0.620)	(0.854)
$Stock_{t-2}$	5.9856***	5.9056***	-1.3115**	-1.2536**	Stock	4.6808***	4.7468***	4.6437***	19.521***
	(7.232)	(7.110)	(-2.571)	(-2.435)		(5.680)	(5.767)	(8.941)	(9.088)
Exch		0.0159		-0.1259	$Stock_{t-1}$	5.9944***	6.0147***	0.3269	3.5853
		(0.317)		(-0.522)		(7.130)	(7.156)	(0.605)	(1.276)
$Exch_{t-1}$		-0.0577		-0.1312	$Stock_{t-2}$	4.6544***	4.5501***	-1.1806**	-0.5246
		(-1.148)		(-0.512)		(5.515)	(5.393)	(2.210)	(-0.186)
$Exch_{t-2}$		-0.0530		0.0005	Exch		0.0124		2.5889
		(-1.054)		(0.002)			(0.251)		(1.589)
					$Exch_{t-1}$		-0.0605		0.2703
							(-1.220)		(0.132)
					$Exch_{t-2}$		-0.0431		-1.3327
							(-0.869)		(-0.650)
Hausman Test (P-value)	0.950	0.951	0.392	0.328	Hausman Test (*P*-value)	0.930	0.930	0.772	0.829
R^2	0.096	0.095	0.015	0.029	R^2	0.116	0.115	0.026	0.397
Observation	1212	1212	714	714	Observati	1212	1212	714	714

*The regressions are estimated using the Generalized Method of Moments (GMM) with random effects for the whole sample of specific countries and for the 1996–2004 period. The dependent variable is capital flow (KI) collected by the EPFR. Stock stands for stock market returns of the specific country. Exch denotes exchange rate variation. The exchange rate is quoted in direct quotations; that is, positive Exch refers to a depreciation of currency. The Hausman test is a test of systematic differences between coefficients of the fixed effects and the random effects regression. We report the p-value of the Hausman test statistic. A constant and year dummies are included but are not reported. T values are between brackets. * Significance at the 10% level. ** Significance at the 5% level. *** Significance at the 1% level.*

Implies that capital persistently flows into Asia and has accumulated to a significant level. This is evidence in the coefficients of the stock market returns. All values, even in the two lagged periods, are significantly positive in Asia but are only remained significantly positive in Latin America in the current period. We conclude here, therefore, that the herding behavior effect does exist in Asia but not in Latin America. Herding behavior destabilizes financial markets since international capital tends to buy past winners, sell past losers, i.e. a positive feedback effect and copy others' behavior.

Table 6 shows various capital volatility levels. Interestingly, in the stable Asian and volatile Latin American countries, the coefficients of the exchange rate variations have significantly negative values, meaning the appreciation of the currency induces more capital inflows. Furthermore, capital inflows during the Asian crisis are also considered in this study but not reported. The reason for adopting the Asian crisis as the benchmark is to examine what the reaction is vis-à-vis international capital when one crisis occurs. Also, capital inflows into Asia occupy a greater percentage, over 50% against a mere 20% in Latin America. We find that the positive feedback effect in Asia was not significant before the Asian crisis though it was significant both during and after the crisis. Also, herding behavior only occurred right after the Asian crisis.

Table 6: The Estimated Results for Models (1) and (2)-- By Capital Volatility

	Positive Feedback Effect: Model (1)					Herding Behavior Effect: Model (2)			
	Stable Countries		**Volatile Countries**			**Stable Countries**		**Volatile Countries**	
	Asia	**Latin America**	**Asia**	**Latin America**		**Asia**	**Latin America**	**Asia**	**Latin America**
Stock	2.7440***	0.3931***	7.2795***	25.0987***	KI_{t-1}	0.2429***	0.0863*	0.1396***	0.0006
	(4.233)	(3.497)	(4.715)	(18.575)		(6.537)	(1.774)	(3.113)	(0.010)
$Stock_{t-1}$	3.1455***	0.02411	10.0757***	3.2899**	KI_{t-2}	-0.0387	0.0762	0.0310	0.0050
	(4.831)	(0.217)	(6.379)	(2.441)		(-1.034)	(1.576)	(0.693)	(0.087)
$Stock_{t-2}$	4.1103***	-0.1291	7.7942***	-0.3892	Stock	2.4970***	0.3897***	6.7936***	22.8966***
	(6.385)	(-1.167)	(5.010)	(-0.288)		(3.999)	(3.554)	(4.437)	(16.459)
Exch	-0.0050	-0.0444	-0.2745	-285.0282***	$Stock_{t-1}$	2.3313***	-0.0100	8.8113***	4.4712**
	(-0.206)	(-0.947)	(-0.232)	(-3.424)		(3.681)	(-0.091)	(5.522)	(2.347)
$Exch_{t-1}$	-0.0574**	-0.0345	-1.1335	-198.9913**	$Stock_{t-2}$	3.1777***	-0.1790*	6.1776***	-0.3913
	(-2.336)	(-0.693)	(-0.797)	(-2.333)		(5.019)	(-1.646)	(3.851)	(-0.202)
$Exch_{t-2}$	-0.0475*	-0.0165	-0.1009	32.5664	Exch	0.0020	-0.0388	-0.2009	0.6547
	(-1.943)	(-0.331)	(-0.070)	(0.384)		(0.088)	(-0.851)	(-0.171)	(0.451)
					$Exch_{t-1}$	-0.0541**	-0.0264	-1.1099	0.5565
						(-2.291)	(-0.543)	(-0.788)	(0.349)
					$Exch_{t-2}$	-0.0389	-0.0094	0.1437	0.2889
						(-1.644)	(-0.193)	(0.101)	(0.179)
Hausman Test (P-value	0.933	0.609	0.965	0.397	Hausman Test (*P*-value)	0.791	0.838	0.976	0.565
R^2	0.076	-0.096	0.1123	0.5946	R^2	0.106	-0.068	0.1228	0.4770
Obs	702	408	510	302	Obs	720	408	510	302

Same as in Table 5. In Asia, the volatile countries are China, Hong Kong, Korea, Singapore and Taiwan. In Latin America, the volatile countries are Argentina, Brazil and Mexico

Regarding the augmented Granger causality test, the results for the future, present and past values of KI are reported in Table 7. Since simultaneous-equation estimation is only asymptotically consistent, we only use the larger data sets among the ones available for this paper. For Equation (4), we use the two largest sample sizes for the total sample and Asia.

The results show that there is only weak evidence of a two-way causality: although the present value of capital inflow is significant, the past values and the F test for the joint significance of the past values are both insignificant. Additionally, a Ramsey RESET test reveals that this parsimonious model might have omitted variables (the p-values for the F test of the null hypothesis that that there is no omitted variable are 0.0002 for both samples). This problem might cause the t test and F test to become invalid. Hence, we estimate Equation (5) as an alternative test.

We again use the two largest sample sizes for the total sample and Asia from the EPFR. The results using the total sample shows no evidence of a possible two-way causality. The results using the data for Asia show only weak evidence of a possible two-way causality: the present value of capital flow is significant, whereas the past values and the F test for the joint significance of the past values are both insignificant. The Ramsey RESET tests fail to reject the null hypotheses of no omitted variable (p-values for the F test is from 0.6342 to 0.8351), so the t-test and F test are valid. We therefore conclude that simultaneous-equation estimation is not needed. The above analysis concerning omitted variables implies that the results reported in Tables 5 and 6 are less reliable than those in Tables 8 and 9. As a robust check, we repeat all regressions using the IFS data. The result, which are not reported here, are quite similar to the results using the EPFR data.[5]

Table 7: The Augmented Granger Causality Test —Data for Asia + Latin America

	Asian Volatile	Asian Stable	L.A. Volatile	L.A. Stable
KI_{t-1}	-0.2873***	0.0804	-0.5404***	-0.4867***
	(-3.266)	(1.047)	(-5.375)	(-5.216)
Stock	10827.1208**	1499.2701***	3574.0909***	270.8589**
	(1.980)	(4.487)	(3.226)	(2.505)
$Stock_{t-1}$	-1485.3531	1199.3420***	1639.8361	429.3595***
	(-0.341)	(3.449)	(1.453)	(3.730)
RGDP	8.0703	6.7499	-31.9699*	0.0979
	(0.097)	(1.025)	(-1.696)	(1.561)
$RGDP_{t-1}$	-10.2423	1.5630	-2.8198	0.1917***
	(-0.129)	(0.250)	(-0.150)	(4.170)
INT	2028.3868	10.8909	-45.669**	6.2364***
	(1.607)	(0.203)	(-2.325)	(7.799)
INT_{t-1}	-3007.5987	1.5630	-19.0512	3.8946***
	(-1.592)	(0.250)	(-0.868)	(3.049)
FR/Import	-353.2585	72.9137	-120.7671	1.1823
	(-0.147)	(0.252)	(-0.352)	(0.051)
Exch	0.0434	0.0286	-146.9706	-0.0145
	(0.032)	(0.286)	(-0.224)	(-0.103)
M2/FR	-427.1581	1049.5667**	-487.7206	-7798.2000**
	(-0.245)	(1.990)	(-1.069)	(-2.954)
Hausman Test (P-value)	0.981	0.993	0.781	0.723
R^2	-0.045	0.349	0.351	0.752
Observations	158	216	96	108

Same as in Table 5.

CONCLUSION

The purpose of this paper is to investigate the behavior of international capital flow and its determinants, using the *EPFR* data set. We evidence that international capital has a Positive feed information symmetry does exist in the emerging countries we investigate, and therefore, herding behavior does hold. After the Asian crisis, herding activity is a dominant factor; it could have been an issue that worsened the Asian financial crisis. For this reason, the government, fund managers and even individuals should take these capital characteristics into consideration for policy making or investing decisions.

We believe that the change of regulation will cause further capital movement, even market panic. Taking the Black Tuesday (Feb. 27th , 2007) in China for example, due to widespread rumors of plans by the Chinese government to raise interest rates or institute a capital gains tax, Shanghai and Shenzhen stock markets had been up about 10% for the year before the Tuesday's decline. Also in Thailand, Finance Minister announced the new control on capital which would remain on foreign investments in bonds and commercial paper as part of central bank's measures to stem the surge of speculative investment in the Thai baht. This made investors dump stocks in Hong Kong, India, Indonesia, Malaysia, South Korea and the Philippines amid contagion concerns that the plunge might spread through the region.

Table 8: The Estimated Results for Model (3)—Data for Asia

KI_{t-1}	0.1254**	0.1271**	0.1147**	0.1220**	0.1197**	0.1120**
	(2.495)	(2.513)	(2.262)	(2.399)	(2.350)	(2.180)
Stock	1778.520***	1781.984***	1854.657***	1950.773***	1935.192***	1918.297***
	(7.427)	(7.421)	(7.434)	(7.715)	(7.581)	(7.500)
$Stock_{t-1}$	1140.020***	1124.813***	1190.278***	1243.676***	1238.444***	1240.927***
	(4.954)	(4.731)	(4.911)	(5.131)	(5.080)	(5.091)
RGDP		-0.0938	2.4300	3.0289	2.8340	3.3052
		(-0.026)	(0.619)	(0.751)	(0.645)	(0.734)
$RGDP_{t-1}$		6.7975*	9.2574**	8.2470**	9.1312**	8.7585**
		(1.870)	(2.372)	(2.076)	(2.142)	(2.010)
INT			43.0734	33.9418	26.7783	27.6909
			(0.984)	(0.776)	(0.578)	(0.589)
INT_{t-1}			-14.5016	-6.9677	-2.9473	4.1040
			(-0.186)	(-0.090)	(-0.037)	(0.051)
FR/Import				239.8042	236.9390	212.1334
				(1.583)	(1.557)	(1.362)
Exch					-0.0409	-0.0337
					(-0.420)	(-0.346)
M2/FR						37.4870
						(0.501)
Hausman Test (P-value)	0.813	0.791	0.901	0.948	0.950	0.978
R^2	0.215	0.210	0.226	0.240	0.233	0.232
Observations	380	376	376	376	376	374

Same as in Table 5 The dependent variable is the capital flow (KI) collected by EPFR and IMF's IFS. Stock stands for the stock market return of the specific country. RGDP is real GDP growth rate. INT denotes interest rate differential against LIBOR US dollar lending. FR/Import, foreign reserves divided by import value, is a measure of reserve strength. Exch denotes the exchange rate variation. M2/FR, broad money supply divided by foreign reserves, is a measure of financial liberalization.

Table 9: The Estimated Results for Model (3)—Data for Latin America

KI_{t-1}	-0.3869***	-0.3865***	-0.3976***	-0.4497***	-0.4496***	-0.4723***
	(-5.955)	(-5.921)	(-6.139)	(-6.391)	(-6.356)	(-6.785)
Stock	219.4572	228.8849	404.8820	1791.893***	1806.224***	1897.288***
	(0.679)	(0.682)	(1.191)	(3.012)	(3.023)	(3.225)
$Stock_{t-1}$	-24.7686	-56.5066	-28.3038	638.7563	656.2112	703.9991
	(-0.074)	(-0.155)	(-0.077)	(0.977)	(0.993)	(1.081)
RGDP		0.0027	0.0855	0.4420	0.4445	0.4719
		(0.013)	(0.397)	(1.447)	(1.451)	(1.551)
$RGDP_{t-1}$		-0.0417	-0.0379	0.2333	0.2307	0.2509
		(-0.207)	(-0.175)	(0.757)	(0.746)	(0.809)
INT			6.1034	8.1144	8.1150	8.3570*
			(1.349)	(1.578)	(1.577)	(1.650)
INT_{t-1}			-6.0732	-5.1263	-5.1223	-5.0361
			(-0.980)	(-0.747)	(-0.746)	(-0.757)
FR/Import				54.5753	57.5858	-36.6716
				(0.377)	(0.379)	(-0.226)
Exch					0.0234	0.1588
					(0.379)	(0.115)
M2/FR						-686.8490**
						(-2.482)
Hausman Test (P-value)	0.803	0.800	0.819	0.993	0.992	0.827
R^2	0.107	0.089	0.133	0.218	0.202	0.222
Observations	224	224	224	204	204	204

Same as in Table 5. Dependent variable is the capital flow (KI) collected by the EPFR and IMF's IFS. Stock stands for the stock market return of the specific country. RGDP is real GDP growth rate. INT denotes interest rate differential against LIBOR US dollar lending. FR/Import, foreign reserves divided by import value, is a measure of reserve strength. Exch denotes the exchange rate variation. M2/FR, broad money supply divided by foreign reserves, is a measure of financial liberalization.

Just one after, Thai government lifted controls on foreign investment after market had plunged fifteen percent, rattling regional bourses amid worries about a repeat of the 1997 Asian financial crisis. Thus, regulations do have their impacts on capital movement and deserve for further case study.

Through this study, it is determined that capital flows play a crucial role in the international capital market. Our understanding of the behavior of capital flows and its determinants not only increase our knowledge on returns, but also enable us to manage in such a way that would avoid any potential financial damage.

ENDNOTES

[1] Eichengreen, and Arteta (2000), have provided a significant volume of literature in this area.

[2] Many articles and monographs on the Asian financial crisis are found in the extant literature, such as those by Chang and Velasco (1998), the International Monetary Fund (1998), Goldstein (1998), Kwack (1998, 2000), Letiche (1998), Moreno, Pasadilla and Remolona (1998), Radelet and Sachs (1998), the World Bank (1998) and Corsetti, Pesenti and Roubini (1999a, 1999b).

[3] Existing literature on mutual funds' domestic strategy abounds, such as Grinblatt et al. (1995), Warther (1995) and Wermers (1999) among others that focus on the U.S.; Kim and Wei (2002a, 2002b) study Korean cases.

[4] The Permanent Income Hypothesis implies that portfolio re-allocation should take place instantaneously and lagged information should be irrelevant. However, empirical evidence from existing literature shows that lagged of many variables such as stock prices, interests etc. are relevant.

[5] However, they are available upon request.

REFERENCES

Berg, Andrew, and Catherine Pattillo (1998) "Are Currency Crises Predictable? A Test," *International Monetary Fund Working Paper*, WP/98/154

Bikhchandani, Sushil, and Sunil, Sharma (2001) "Herd Behavior in Financial Markets," *IMF Staff Papers*, vol. 47 (3), p.279-310

Borensztein, Eduardo and R. Gaston Gelos, (2003) "A Panic-Prone Pack? The Behavior of Emerging Market Mutual Funds, " *IMF Staff Papers* 50, p. 43-63

Calvo, G., Leiderman, L., and Reinhart, C. (1993) "Capital Inflows and Real Exchange Rate Appreciation in Latin America: The Role of External Factors,"*IMF Staff Papers*, 1 (40), p. 108-151

Calvo, G. A., (1996) "Capital Flows and Macroeconomic Management: Tequila Lessons," *International Journal of Finance & Economics*, 1, p.207-224

Chang, R. and A. Velasco (1998) "The Asian liquidity crisis," *Federal Reserve Bank of Atlanta Working Paper*, no.98-11

Choe, Hyuk, Bong-Chan Kho and Rene M. Stulz (1999) "Do Foreign Investors Destabilize Stock Markets? The Korean Experience in 1997," *Journal of Financial Economics*, 54, p. 227–264

Corsetti, G., P. Pesenti and N. Roubini (1999a) "What Caused the Asian Currency and Financial Crisis?" *Japan and World Economy*, 11 (3), p.305-373

Corsetti, G., P. Pesenti and N. Roubini (1999b) "Paper tigers? A model of the Asian crisis," *European Economic Review*, 43, p.1211-1236

Dornbusch, R. Y. C. Park, and S. Claessens (2000) "Contagion: Understanding How It Spreads," *The World Bank Research Observer*, vol. 15 (2), p.177–197

Eichengreen, B. and C. Arteta (2000) "Banking Crises in Emerging Markets: Presumptions and Evidence," *University of California, Berkeley*

Fernandez-Arias, Eduardo. (1996.) "The New Wave of Private Capital Inflows: Push or Pull?" *Journal of Development Economics*, 48, p. 389-418

Fernandez-Arias Eduardo and Peter J. Montiel (1996) "The Surge in Capital Inflows to Developing Countries: An Analytical Overview," *The World Bank Economic Review*, vol. 10 (1), p. 51-77

Froot, Kenneth, Paul G. J. O'Connell and Mark Seasholes (2001) "The Portfolio Flows of International Investors," *Journal of International Economics*, vol. 59 (2), p.151–193

Goldstein, M. (1998) *The Asian financial crisis: causes, cures, and systemic implications*, Washington, DC: Institute for International Economics

Grinblatt, Mark, Sheridan Titman and Russ Wermers (1995), "Momentum Investment Strategies, Portfolio Performance, and Herding: A Study of Mutual Fund Behavior," *American Economic Review*, vol. 85 (5), p.1088–1105

International Monetary Fund (1998), "The Asian crisis and the region's long-term growth performance," *World Economic Outlook*, Chapter III, p82-106

Kaminsky, G., and C. Reinhart (1999) "The Twin Crises: The Causes of Banking and Balance of Payment Problems," *American Economic Review*, vol. 89 (3), p.473-500

Kim, W. and S.J., Wei (2002a) "Foreign Portfolio Investors Before and During a Crisis," *Journal of International Economics*, vol. 56 (1), p. 77-96

Kim, Woochan and, S. J. Wei (2002b) "Offshore Investment Funds: Monsters in Emerging Markets?" *Journal of Development Economics*, 68, p. 205-224

Kuo, M. H. and C. C., Chi (2000) "The Investment Behavior and Performance Around the Asian Financial Crisis: Foreign Investment Institutions in Taiwan," *Management Review*, vol. 19 (1), p. 81-118

Kwack, S. Y. (1998) "Factors contributing to the financial crisis in Korea," *Journal of Asian Economics* vol. 9 (4), p. 611-625

Kwack, S. Y. (2000) "An empirical analysis of the factors determining the financial crisis in Asia," *Journal of Asian Economics*, 11, p.195-206

Leitche, J. M. (1998) "Causes of the financial and economic crisis in Southeast Asia and the need for national regional and IMF structural reform," *Journal of Asian Economics*, vol. 9 (2), p. 181-191

McKinnon R. (1999) "The East Asian dollar standard: life after death," *Economics Department Stanford University*, Working Paper no. 99-017

Moreno, R., G. P. and E. Remolona (1998), "Asia's financial crisis: lessons and policy responses," *Pacific Basin Working Paper*, No. PB98-02, Federal Reserve of San Francisco

Obsfeld, Maurice (1998) "The Global Capital Market: Benefactor or Menace?" *Journal of Economic Perspective*, vol. 12 (4), p.9-30

Parkinson, M. (1980) "The Extreme Value Method for Estimating the Variance of the Rate of Return," *Journal of Business*, vol. 53 (1), p. 61-65

Radelet, S., and J. Sachs (1998) "The East Asian Financial Crisis: Diagnosis, Remedies, Prospects," *Brookings Papers on Economic Activity*: Macroeconomic, 1, 1- 90

Rajan R., Siregar R. and Sugema I. (2003) "Why was there a pre-crisis inflow boom in southeast Asia? " *Journal of International Development*,15, p. 265–283

Sachs, Jeffrey, Aaron Tornell, and Andres Velasco. (1996) "Financial Crises in Emerging Markets: The Lessons from 1995," Brookings Papers on Economic Activity 1:147–215

Shen, C. H. and M.F. Hsieh (2000) "Capital inflows and banking vulnerability—a cross-country comparison," *Journal of Financial Studies*, vol. 8 (3), p.1-41 (in Chinese)

Warther, V. A. (1995) "Aggregate Mutual Fund Flows and Security Returns," *Journal of Finance Economics*, 39, p. 209-235

Wermers, Russ (1999) "Mutual Fund Herding and the Impact on Stock Prices," *The Journal of Finance*, vol. LIV (2), p.581-622

World Bank (1998), *East Asia: the road to recovery*, Washington, DC: The World Bank

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A COINTEGRATION TEST TO VERIFY THE HOUSING BUBBLE

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ABSTRACT

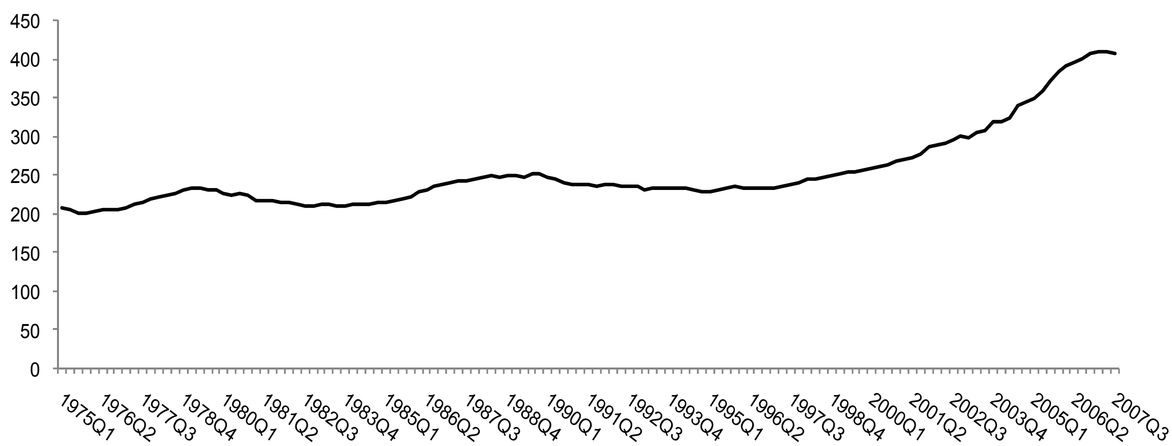
Housing prices in the US rose rapidly from 2000-2007Q3. Based on this evidence, the financial and general press concluded the US experienced a housing bubble. The efficient market theory denies the possibility of a bubble. This paper applies the statistical technique of cointegration to substantiate the presence of a housing bubble. The paper finds the statistical evidence consistent with the presence of a housing bubble in the period 2000-2007Q3 and not the underlying economic conditions.

JEL: C53, G12, G18

INTRODUCTION

Recently housing prices in the US have risen substantially. Figure 1 shows the trend in the US Housing index from 1975 to 2007Q3. Clearly, from 2001 the US has experienced several years of strong home price increases. This sort of evidence has led the financial and the general press to conclude the US housing market has experienced a bubble. Today the press speculates that this bubble has deflated and may soon pop if it has not already popped. Nevertheless, prolonged rapid increases in prices do not imply the presence of a bubble. Nor does an ensuing sharp price decline following a price run-up necessarily imply a bursting bubble. The former event may simply reflect the changes in the fundamental economic factors, such as the low level of mortgage rates in the case of housing prices. Lowenstein (2007) explores this possibility in the general press. The latter event may be a result of a sudden reversal in the underlying fundamental economic factors.

Figure 1: Home Prices in the United States (1975-2007Q3)



US Housing index from 1975 to 2007Q3

The bursting of a real estate bubble has important implications for the US economy. Residential real estate is an important component of householder wealth. In 1996, it represented 39% of household wealth; today it represents 49% (Roubini, 2006). Householder wealth is supporting the high current level of consumption. If a significant portion of real estate wealth disappears, it is likely that consumption will fall considerably. It is important to remember that the housing boom was strongly encouraged by the low interest during the 1990 decade. Many mortgages are adjustable rate. In today's environment of higher interest rates the higher mortgage payments coupled with lower home values will discourage consumer spending particularly on consumer durables. Further, the employment effects of the housing industry are considerable. In the last three years, 30% of the job growth was due directly or indirectly to housing (Roubini, 2006). The construction and real estate mortgage job loses could be considerable if housing prices drop substantially. In sum, a real estate collapse could tip the economy into a recession. We shall attempt statistically to determine whether the US housing market experienced a bubble after 2000. The detection of a bubble is important for two reasons. First, we may prepare for the effects of the bubble bursting on the remainder of the economy. Second, early detection of a bubble allows the regulators to "remove the punch bowl" and avert some of the difficulties of a housing market crash.

LITERATURE REVIEW

The widely accepted efficient market theory claims that financial asset prices reflect all the publicly available information at all times. This denies the possibility of a bubble. While some may believe prices are too high relative to fundamental factors, according to the theory they are wrong. Because investors recognize immediately if the price of anything is too high (or too low) and respond by selling (or purchasing), the asset until the over-(under) pricing is eliminated. A mountainous body of academic research supports this view, Fama (1970).

Nevertheless, the efficient market theory has been subject to much serious criticism, Shiller (2004). Furthermore, much of the research focused on financial assets. The efficient market theory assumes that investors can sell an asset short to eliminate overpricing. Real estate is a real and illiquid asset. During the period of the housing price run-up there was no mechanism known to us for shorting a residential home. A futures market for housing is a relatively recent innovation. These markets do not function well enough to fulfill the assumptions of the efficient market theory. Thus, we should not dismiss the possibility of a housing bubble out of hand.

No single explanation for real estate bubble or bubbles in general has achieved acceptance. Since this literature is too vast for this paper, we will focus mostly but not exclusively on popular explanations and indicate why they do not provide sufficient explanations for bubbles. The greater fool theory purposes that bubbles are driven by optimistic investors (fools) who buy with the expectation of selling to even more optimistic investors (bigger fools) at a profit. The bubble finally bursts when the greater fool fails to find another buyer. Thus, the greatest fool suffers a loss and the bubble unwinds. Many of the popular explanations of the housing bubble rely on the greater fool theory. Consider the following: Supposedly, the American dream of home ownership fuels a frenzy home buying unjustified by economic reality. Americans believe home prices cannot fall, despite the reality that they sometimes do. The proliferation of TV real estate reality programs and other promotions create froth in home prices.

Unfortunately, the evidence is not favorable to this popular theory. Lei et al (2001) find that bubbles occur even in the absence of overconfidence. They attribute bubbles to irrational behavior. Levine and Zajac (2007) show that even without uncertainty or the possibility of speculation bubbles appear even in well-organized experimental markets. They "find no support for the greater fool explanation." They believe the cause lies in "the institutionalization of social norms." People copy the behavior of others.

Another explanation that may be especially relevant for the real estate market is that bubbles result from increases in the money supply. When the money supply grows too rapidly the concomitant lower interest rates discourages investment in savings accounts and encourages investment in financial assets and real estate. The low interest rates reinforce this by encouraging the use of leverage. The bubble ends when the money supply contracts. However, Smith, Suchneck, and Williams (1988) find bubbles even in experimental markets that preclude money supply changes.

METHODOLOGY

The first step in supporting the presence of a bubble is to provide a clear definition of the term. Case and Shiller (2003) define it as:

> …a situation in which excessive public expectations of future price increases cause prices to be temporarily elevated.

They argue that during a bubble homeowners spend more on a home than they otherwise might because they a believe that the house will be worth even more in the future or that the home will soon be even more unaffordable. In sum, expectations of higher housing prices make housing an irresistible investment, high return and virtually risk-free. Since housing prices cannot go up indefinitely, when they stop rising, expectations are revised and the bubble bursts.

According to Case and Shiller's (2003) definition, a bubble is a psychological construct, so it is natural to conduct investor surveys to ascertain the presence of a bubble. They surveyed recent homebuyers in three cities that apparently experienced a rapid price increases along with those in one city that has not. They find that homebuyers in 2003 expect home prices to increase in both the long and short run. Unlike their 1988 survey, this was even true for the city with the slow housing price increases. This clearly supports the presence of a bubble as defined by Case and Shiller (2003). They also found a strong belief that prices are sticky downward. Furthermore, they found buyers in the glamour cities inclined to support simplistic theories that might translate into behavior that supports a bubble. For example, buyers in the glamour cities were more likely to believe "that desirable real estate just naturally appreciates." (p. 325). This is naïve because the argument suggests only that desirable real estate will have a high price but says nothing about how that price will change. Yet the prevalence of this belief might explain how a bubble continues without the participants recognizing it. Their survey data leave the reader with the strong impression of the presence of a housing market bubble.

As Case and Shiller (2003) recognize many economists oppose drawing conclusions on economic behavior based upon consumer's thought processes. Economists prefer to focus on consumer actions. Thus, they also reinforce their case by presenting some statistical evidence. They find that regressions including income and other fundamentals explain housing prices in all but the eight glamour states. In those states, the regressions underestimate home prices during the 2000-2002 period.

In this paper, we will present additional statistical evidence in favor of the plausibility of a bubble in the real estate market. This should shed more light on the subject for those both sympathetic and unsympathetic to survey research. The cointegration methodology developed by Granger and Engle (1987) provides a method of testing for the existence of a bubble. Diba and Grossman (1988) and Campbell and Shiller (1987) apply this methodology to test for the presence of a bubble in the equity market. We use this methodology to check for a bubble in the housing market.

To explore the existence of a housing bubble we examine the stability of the underlying relationship of home prices and the economic forces that determine them. A relationship suddenly becomes unstable during a period of rising home prices is consistent with the presence of a bubble. Cointegration is well suited to test for this. Cointegration tests for the long run relationship among variables. Cointegration implies two variables share a common stochastic trend. A common stochastic trend does not simply mean that they move upward or downward together, but rather that the variables may share both prolonged upward and prolonged downward movements.

Suppose housing prices are cointegrated with an economic variable and a bubble develops in the housing market then housing prices rise without a corresponding rise in the variable. This implies the severing of a long-term relationship between housing prices and the variable. In other words, the cointegration should cease. In summary if there was a bubble in the housing market after 2000, we should be able to find variables, which were cointegrated with housing prices before 2001, and no longer remain cointegrated.

TESTING FOR COINTEGRATION

The first step in establishing cointegration is to test the variables for stationarity. Roughly speaking, a time series is stationary if the series fluctuates with a constant variation about a mean, which remains constant over time. A stationary variable has a propensity to return to and frequently cross the mean. When it is far from the mean, it is more likely to move towards rather than away from the mean. Since a stationary variable shows, a tendency to revert to a constant mean it exhibits no trend. Thus, only variables, which are non-stationary, may be cointegrated. To establish nonstationarity we employ the ADF (Augmented Dickey Fuller) test and the Phillips Peron test. The second step employs Johansen-Juselius procedure (1990) to test for cointegration. This test for cointegration entails estimating the following cointegration regression:

$$Y_T = a + bx_t + u_t \tag{1}$$

Where for our 2 variable case y_t is the housing price index and x_t represents a fundamental variable that is hypothesized to influence housing prices, a and b are regression parameters, and u_t is the error term.
We may rewrite equation (1) as:

$$\Delta y_t = a + (b-1)\, x_{t.-1} + u_t \tag{2}$$

If the variables are cointegrated then b, the slope term, equals one. Thus, the test revolves around the magnitude of the estimated b coefficient. If the null hypothesis b=1 is not rejected, then equation (2) is a stable autoregressive equation and the data are inconsistent with cointegration. If we find the data consistent with b=1, then equation (2) is a random walk, in which case the variables are cointegrated.

The Johansen- Juselius procedure views this equation in a matrix format. It tests the magnitude of the b coefficient indirectly. It tests by measuring the rank of the coefficient matrix. In the two variable case we consider, the matrix is a 1x1 consisting of the b-1 term. The rank can only be zero or one. The magnitude of the eigenvalue of the matrix determines the rank of the matrix. The rank of the matrix equals the number of nonzero eigenvalues. If the eigenvalue is not statistically different from zero, then the rank is zero. Then we reject the null hypothesis (b=1); this leads to the conclusion that the variables lack cointegration. If the eigenvalue is statistically different from zero, then we conclude the rank of the matrix equals one. This implies the variables are cointegrated.

Data

Quarterly data is used and the study covers the period 1975Q1 -2007Q2. We employ the U.S. Office of Federal Housing Enterprise Oversight (OFHEO) Home Price quarterly index to measure housing prices. The index is not seasonally adjusted. A difficulty faced in the construction of a housing index is the extreme lack of uniformity of the commodity. To control for this OFHEO adopts the Case-Shiller geometric-weighted repeat sales procedure. Calhoun (1996) provides the details of the procedure. Next, we consider a series of seven variables, which reflect the fundamental economic forces determining housing prices. The most important of these is income. Case and Shiller (2003) conclude that in non-bubble markets income explains most of the rise in housing prices. We employ two separate measures of income. The first is the mean of the middle quintile of the income distribution, denoted as the Middle Fifth. Second, we use the mean of the highest quintile of the income distribution denoted as Top Fifth. This attempts to account for the possibility that the wealthiest segment of the population influences housing prices disproportionately because of their greater mobility. The U.S. Census Bureau, *Historical Income Tables-Families* (all races), and the National Association of Realtors provided this data.

The Mortgage Rate represents a strong influence on consumer demand for housing. We obtained 30-year conventional mortgage rate (fixed rate, first mortgages) from the Board of Governors of the US Federal Reserve System. The civilian unemployment rate measures the state of the economy. The U.S. Bureau of Labor Statistics provided the seasonally adjusted percent of civilian unemployment. We converted the monthly data for both variables to quarterly data by a simple mean. The Homebuilders Stock Index provides an indication of the state of the housing market. Merrill Lynch supplied a capitalization-weighted, price level index of homebuilding stocks based on stocks included in the S&P 500 stock index.

The final variables measure the ability of consumers to handle mortgage debt. The Household Debt Ratio is the ratio of household credit market debt outstanding to annualized personal disposable income. This data also came from the Board of Governors of the US Federal Reserve System. The Housing Affordability Index for all homebuyers (HAI) measures whether or not a typical family could qualify for a mortgage loan on a typical home, assuming a 20% down payment. We define a typical home as the national median-priced, existing single–family home as calculated by NAR. In its final form used here, the HAI is essentially "median family income divided by qualifying income." The index is interpreted as follows: a value of 100 means that a family with the median family income (from the U.S. Bureau of the Census and NAR) has exactly enough income to qualify for a mortgage on a median-priced home. National Association of Realtors (NAR) provides the data. In this research, the monthly HAI values result from quarterly samples. For more details on the exact calculation, go to www.realtor.org.

PRELIMINARY ANALYSIS

Table 1 shows the simple correlation among the variables during the period before rapidly escalating home prices (1975Q1 – 2000Q4). During this period the fundamental variables exhibited a strong association with the lowest correlation coefficient of .42 between the HAI and the Mortgage Rate and the high of -.96 (both in absolute terms) between the Income of the Top Fifth and the Household Debt Ratio. More importantly, all the variables displayed strong correlation with home prices and the signs were as expected. Mortgage rates have the lowest correlation of -.48 with home prices while the Income of the Middle Fifth has the highest correlation of .88 with home prices.

Table 2 shows similar results for the whole period (1975Q1 – 2007Q3). The Homebuilder Stock Index and The Housing Affordability Index correlation coefficient of .23 was the lowest; and a correlation of -.94 between The Household Debt Ratio (Debt/Income) and Top-Fifth was the highest. The correlations between the variables and Housing prices were strong for the whole period. The Housing price index had the lowest correlation of .26 with HAI and the highest correlation of .92 with both Household Debt Ratio and Homebuilder Stock Index. Unemployment and The Housing Affordability Index were the only variables to suffer a serious decline (in absolute terms) in correlation with housing prices during the period of escalating housing prices (from -.83 to -.56 for the former and .49 to .26 for the latter). The correlation between income variables and housing prices declined minimally (less than .05). The other variables obtained stronger correlations. Based upon the correlation analysis there is little to support the hypothesis that the relationship housing price with fundamental economic variables was substantially altered during the period of housing price rises.. This occurs because correlation analysis is ill suited to detect this shift for non-stationary variables. Cointegration offers a better alternative to which we turn to in the next section.

Table 1: Correlation Matrix 1975-2000

	Mortgage Rates	Unemployment	Debt/Income	Housing Affordability Index	Home Builder Stock Index	Middle Fifth	Top Fifth	Housing Price Index
Mortgage Rates	1							
Unemployment	0.59	1						
Debt/Income	-0.644	-0.77	1					
Housing Affordability Index	-0.96	-0.6	0.71	1				
Home Builder Stock Index-Builder	-0.42	-0.48	0.71	0.46	1			
Middle Fifth	-0.64	-0.9	0.83	0.65	0.65	1		
Top Fifth	-0.66	-0.79	0.96	0.73	0.76	0.91	1	
Housing Price Index	-0.48	-0.83	0.84	0.49	0.58	0.88	0.82	1

Table 2: Correlation Matrix 1975-2005Q2

	Mortgage Rates	Unemployment	Debt/Income	Housing Affordability Index	Home Builder Stock Index	Middle Fifth	Top Fifth	Housing Price Index
Mortgage Rates	1							
Unemployment	0.65	1						
Debt/Income	-0.74	-0.69	1					
Housing Affordability Index	-0.9	-0.58	0.52	1				
Home Builder Stock Index-Builder	-0.55	-0.42	0.85	0.23	1			
Middle Fifth	-0.76	-0.84	0.91	0.59	0.71	1		
Top Fifth	-0.77	-0.78	0.94	0.66	0.7	0.95	1	
Housing Price Index	-0.59	-0.56	0.92	0.26	0.92	0.83	0.79	1

EMPIRICAL RESULTS

For all the variables, both the ADF tests and the Phillips-Perron statistic support non-stationarity at a five per cent level of significance. (The results are not shown but available upon request). We conclude that all the variables are non-stationary. Next, we examine whether home prices and the seven fundamental variables are cointegrated. We accomplish this by examining a cointegrating regression for each of the seven variables with home prices. Table 3 presents the results of these cointegration tests for the1975Q1-2000Q4 period. The Trace Statistic Test shows that for three of the seven variables, Top Fifth, Middle Fifth and the Unemployment Rate, we may reject the null hypothesis of no cointegration at a five per cent level of significance. Furthermore, we may reject the null hypothesis of no cointegration at the 10% level of statistical significance for one additional variable the Homebuilders Stock Index. Thus for the period preceding the run up in home prices there appears to have been a strong link between home prices and both the income variables and the Unemployment Rate and a marginal link with the Home Builders Stock Index.

Table 4 presents the results of these cointegration tests for the period 1975-2007Q3. The trace tests indicate the eigenvalues are not statistically distinguishable from zero in any equation at the 5%. However, the P-value for the middle fifth of income was .0502. Recognizing the belief that the bubble burst in late 2005 we did the cointegration test for the period 1975-2005Q2. The P-value (hypothesized no. of CE(s) = none) for Home prices vs. middle fifth of income was 11%. Although we did not display the results, we cannot reject the hypothesis of no cointegration for any of the other fundamental variables during this period. (The results are available upon request.) This suggests that in the post 2005 period the normal relationship between Home Prices and income was reasserting itself. This result suggests that the linkage between Home Prices and the fundamental variables weakened substantially after 2000. The evidence is consistent with a real estate bubble.

Table 3: Cointegration Tests 1975-2000

Cointegration between	Hypothesized No. of CE(s)	Eigenvalue	Trace Statistic	0.05 Critical Value	Prob. **
Home Prices vs. Household Debt Ratio	None	0.09	14.72	25.87	0.60
	At most 1	0.05	5.28	12.52	0.56
Home Prices vs. Housing Affordability Index	None	0.13	20.34	25.87	0.21
	At most 1	0.07	7.06	12.52	0.34
Home Prices vs. Mortgage Rate	None	0.10	18.16	25.87	0.33
	At most 1	0.08	7.73	12.52	0.27
Home Prices vs. Homebuilders stock index	None	0.15	25.69	25.87	0.05
	At most 1	0.09	9.49	12.52	0.15
Home Prices vs. Unemployment Rate	None *	0.15	25.90	25.87	0.05
	At most 1	0.09	9.66	12.52	0.14
Home Prices vs. mean of Middle fifth of Income	None *	0.20	32.58	25.87	0.01
	At most 1	0.10	9.90	12.52	0.13
Home Prices vs. mean of Top fifth of Income	None *	0.21	32.29	25.87	0.01
	At most 1	0.08	8.70	12.52	0.20

The table shows the eigenvalues of a regression of housing prices on the seven fundamental economic variables. For each eigenvalue we can find an associated characteristic root. The natural log of the characteristic root times the times the number of quarters (156) yields the trace statistic. This statistic tests the null hypothesis of no cointegration. We may reject this hypothesis if the p-value is less than .05 for the case of the no. of CE(s) [co integrating errors] equals none. * denotes rejection of the null hypothesis of no Cointegration at the 0.05 level
** denotes the p-value

Table 4: Cointegration Tests for the Whole Period 1975-2007Q3

Cointegration between	Hypothesized No. of CE(s)	Eigenvalue	Trace Statistic	0.05 Critical Value	Prob. **
Home Prices vs. Household Debt Ratio	None	0.07	7.68	15.49	0.5
	At most 1	0	0.29	3.84	0.59
Home Prices vs. Housing Affordability Index	None	0.12	13.64	15.49	0.09
	At most 1	0.01	0.89	3.84	0.35
Home Prices vs. Mortgage Rate	None	0.1	10.63	15.49	0.24
	At most 1	0	0.42	3.84	0.52
Home Prices vs. Homebuilders stock index	None	0.1	12.28	15.49	0.14
	At most 1	0.02	1.78	3.84	0.18
Home Prices vs. Unemployment Rate	None	0.1	12.13	15.49	0.15
	At most 1	0.02	1.81	3.84	1.18
Home Prices vs. mean of Middle fifth of Income	None	0.13	15.48	15.49	0.05
	At most 1	0.02	2.02	3.84	0.16
Home Prices vs. mean of Top fifth of Income	None	0.09	8.85	15.49	0.38
	At most 1	0	0.03	3.84	0.87

The table shows the eigenvalues of a regression of housing prices on the seven fundamental economic variables. For each Eigenvalue we can find an associated characteristic root. The natural log of the characteristic root times the times the number of quarters (156) yields the trace statistic. This statistic tests the null hypothesis of no cointegration. We may reject the hypothesis if the p-value is less than .05 for the case of the no. of CE(s) [co integrating errors] equals none. Trace test indicates no Cointegration at the 0.05 level for any variable
*** denotes the p-value*

CONCLUSION

The statistical evidence we presented supports the Case and Shiller (2003) survey conclusions of the presence of a real estate bubble in the housing market. Our study uses a national housing index. Case and Shiller (2003) use local housing indices. They find the bubble limited to a few cities. They find that income has the strongest influence on housing prices. We also find a strong relationship between income and housing prices during the pre-bubble period. This relationship severed during after 2000. Case and

Shiller (2003) speculate that unemployment rates are a key precipitating factor of the housing bubble. Thus, upward (downward) movements in the unemployment rate will decelerate (accelerate) home prices. In contrast, we find a link between home prices and unemployment during the pre-bubble phase. After 2000, the linkage has disappeared. We were somewhat surprised to find the lack of a cointegration relationship between housing prices and mortgage rates.

Since we were able to find the cointegration vanished as early as 2005Q2, we believe cointegration can serve as an early bubble detection system. Alan Greenspan's (former Fed Chairman) statement, "I really didn't get it" (the housing bubble and subprime crisis) until late 2005 or 2006" supports the usefulness of this analysis Perhaps, with cointegration methodology, the Fed could have "removed the punch bowl" in a timely manner.

In the future cointegration analysis should help to identify local housing market bubbles. Since there are housing indices for local markets, this extension should prove feasible. In addition, the technique could help to verify bubbles in other asset markets.

BIBLIOGRAPHY

Calhoun, Charles, 1996, OFHEO House Price Indexes: HPI Technical Description, Office of Federal Housing Enterprise Oversight, Washington D.C.

Campbell, John and Robert Shiller, 1987, "Cointegration and Tests of Present Value Models," *Journal of Political Economy*, 95 (October), p. 1062-88.

Case, Karl and Robert Shiller, 2003, Is there a Bubble in the Housing Market," Brooking Papers on Economic Activities, 2, p. 299-362.

Diba, Behzad and Herschel Grossman, 1988, "Explosive Rational Bubbles in Stock Prices," American Economic Review, 78, p. 520-630.

Dickey, David and Wayne Fuller, 1981, "Likelihood Ratio Statistics for Autoregressive Time Series with a Unit Root," Econometrica, 49, (July), p. 1057-72.

Granger, Clive and Engle, Robert, 1987, "Cointergration and the Error-Correction Representation: Estimation and Testing," Econometrica, 55, p. 251-76.

Engle, Robert and B. Yoo, 1987, "Forecasting and Testing in Co-integrated Systems," *Journal of Econometrics*, 35, p. 143-159.

Fama, Eugene, 1970, "Efficient Capital Markets: A Review of Theory and Empirical Work," *Journal of Finance*, 25, p. 383-417.

Greenspan, Alan, Interview with Jim Lehrer (http: "youtube.com/watch?v=yu4wjhijomy>)". The news show with Jim Lehrer, 18 September 2007.
Johansen, Soren and Katrina Juseluis, 1990, "Maximum Likelihood Estimation and Inference on Cointegration with Application to the Demand for Money," Oxford Bulletin of Economics and Statistics, 52, 169-209.

Lei, Vivian, Charles Noussair, and Charles Plott, 2001, "Nonspeculatative Bubbles in Experimental Asset Markets: Lack of Common Knowledge of Rationality vs. Actual Irrationality", *Econometrica*, 69, p. 831-859.

Levine, Sheen, and Edward Zajac, 2007, "The Institutionalization of Price Bubbles," Available at SSRN http://ssrn.com/abstract=960178.

Lowenstein, Roger, 2007, "On the Home Front; Pop Psychology," New York Times Magazine, March 18, 2007.

Phillips, Peter and Pierre Perron, 1988, "Testing for a Unit Root in Time Series Regression," Biometrika, 75 (June), p. 335-46.

Roubini, Nouriel, 2006, "The Biggest Slump in the Last 40 Years…or 53 Years," Nouriel Roubini's Blog, August 23, 2006.

Shiller, Robert, 2004, "From Efficient Markets to Behavioral Finance," *Journal of Economic Perspectives*, 17 (1), p. 83-104

Smith, Vernon, Gerry Suchenek, and Arlington Williams, 1988, "Bubbles, Crashes, and Endogenous Expectation in Experimental Spot Asset Markets," *Econometrica,* 69, p. 831-59.

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IS THREE A CROWD? CONSIDERING THE VALUE OF MANAGER DIVERSIFICATION FOR ADDING ALPHA

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ABSTRACT

Creating a portfolio that consistently generates alpha—market-adjusted abnormal returns—is the holy grail of active management. Given that excess returns can come both from manager skill and from luck, some advocates of active management suggest that active funds should be combined into diversified portfolios, eliminating all but "pure" active risk and thereby optimizing the risk/return trade-off. In this paper, we present a simple model of such a diversified portfolio, and show that under certain conditions a portfolio manager actually would be better off by not diversifying.

JEL: G11

INTRODUCTION

Active fund managers are paid to deliver excess returns. However, even when they succeed, it is difficult for the portfolio manager who employs them to determine whether those returns were borne of luck or skill. It is obviously not the goal of an portfolio manager to pay extra for luck. Instead, her goal is to identify the active fund managers who earn their fees through skill, consistently delivering "pure alpha."

However, given that even skilled fund managers can be unlucky sometimes, is it not a good idea to create a diversified portfolio of these managers, reducing risk while protecting pure alpha? Should we not run portfolios of funds the same way we run portfolios of stocks, deriving the benefits of diversification?

If a portfolio manager chooses to concentrate, she is exposed to the full risk of one fund manager, eliminating the potential for risk reduction. However, if she diversifies, she must choose *many* managers who will outperform, not just one. Successful active management requires two levels of active skill: the first at the fund level (skillful active fund managers must exist), and the second at the portfolio level (the portfolio manager must be able to identify those skillful fund managers). Proponents of diversified active portfolios, such as Waring and Siegel (2003), suggest that if the portfolio manager is not skilled, she should index. However, as we will show, even if she is skilled, she may nonetheless be better off by choosing not to diversify her active portfolio.

In this paper, we present a model of the tradeoff between a diversified portfolio of active managers and a concentrated one. Using the information ratio (the ratio of alpha to active risk) as the evaluation criterion, we find that a portfolio manager's choice between a concentrated and a diversified portfolio depends upon the proportion of truly skillful managers in the active pool, the correlations among active managers' returns, and the number of fund managers chosen for the portfolio. If there are too few skilled fund managers or if correlations among the merely lucky are too high, a portfolio manager is better off by not diversifying her portfolio.

The paper proceeds as follows. In the next section, we briefly review related literature. We then discuss various expressions for alpha that are common in the institutional literature, and define the decision metric, the information ratio. In the fourth section, we present and discuss the model of the portfolio manager's decision, and present a numerical example. After a brief discussion of the model's implications in the fifth section, we summarize and conclude.

RELATED LITERATURE

Two strands of literature inform this study. The first addresses squarely the question at hand: how many funds should a portfolio manager employ? The answers provided tend to be either in the range of 20 to 30 funds (perhaps an extension of earlier work on the optimal number of stocks in a portfolio; see, for example, Statman, 1987) or, at the other end of the spectrum, no more than ten. Both answers come primarily from experience and simulation, rather than from theory. However, the more recent work, which moves beyond the traditional focus on mean and variance to incorporate higher moments of the return distribution, tends to suggest that smaller is better. We will survey this literature in this section. The second strand of related literature concerns the variables underlying our model, the most important of which are the correlations among funds. We will discuss this literature in section five, after we present the model.

The conventional wisdom is that diversification of funds is good. For example, Ross (2003), in a paper promoting an allocation to hedge funds in institutional portfolios, emphasizes the need for diversification within the added hedge fund portfolio. He stresses that "[t]he returns on a *diversified* portfolio of hedge funds are relatively uncorrelated with other assets" and that "a *diversified* portfolio of hedge funds offers stable returns that are significantly higher than those for bonds, exhibiting very low volatility" (emphasis added). Thus, hedge funds can offer superior risk/return characteristics to bonds, but only if their own unique risks are properly diversified. This requires not only diversification across hedge fund strategies, but also within each strategy: "*a well managed portfolio of hedge funds should have adequate diversification across managers as well as across strategies*" (page 7; emphasis original).

Park and Staum also advocate for significant diversification in funds of funds. They note that, while the average number of funds in portfolios is about five, "…with hundreds… of hedge funds in existence, it is hard for a fund of funds manager to claim that he can only find five good ones" (page 3). Asserting that under-diversifying means foregoing profits, they urge fund managers to "embrace diversification more fully." However, like Ross (2003), Park and Staum do not suggest a specific number of funds that would be required to effect this embrace.

Nesbitt, *et al.* (2003) do provide a number. Based on samples from historical returns (which we will discuss in more detail later), they conclude that "… an optimal portfolio of hedge funds includes between 20 and 30 individual funds. This provides substantial diversification without substantially reducing the potential for adding alpha" (reported in Bonafede, *et al.*, 2004, page 22). This conclusion is based on the behavior of portfolio mean and variance as n, the number of included funds, increases. However, expanding the set of relevant moments to include skewness and kurtosis, as in recent academic work, suggests that using even 20 funds may lead to portfolio overdiversification.

For example, Brands and Gallagher (2005) use simulations to study the effects of diversification on Australian funds of funds (FoFs). The authors find that return stays almost constant as the number of funds increases (in contrast to Park and Staum's foregone profits argument). Variance decreases at a decreasing rate, with most of the benefits accruing by n=6. However, the risk reduction is "slight," decreasing variance by at most 7% for a portfolio of 30 funds. The authors explain this result by noting "a diminishing increase in the number of unique securities added to the FoF as the number of funds in the FoF rises, given increasing levels of common stock holdings across funds" (page 190). (In fact, as the number of funds employed rises, portfolios run increasing risk of becoming closet index funds—forcing investors to pay active management fees for benchmark-like performance.) Combining reward and risk measures into the Sharpe ratio (a common performance metric, discussed more fully in the next section), Brands and Gallagher note that diversification can increase the ratio by 3.57% as n increases, but that "beyond a portfolio of 4 funds, it is not possible to significantly improve the fund's Sharpe Ratio" (page

193). Thus, these authors suggest that diversification benefits, while possible, are limited, and that they can be fully captured using only a very few funds.

On the other hand, they also note that diversification may hurt higher-level moments of portfolio performance. For example, investors are assumed to prefer positive skewness and lower kurtosis. However, as *n* increases, skewness in their funds of funds becomes increasingly negative—the chance of including poor funds increases as more funds are chosen. (This is exactly the effect that we will discuss below.) There is a comparable effect on kurtosis: as *n* rises to about ten, distributions become more peaked, and less attractive to investors. Amin and Kat (2002) find similar results: using data from 2,183 hedge funds and funds of funds from 1994 through 2001, they show that increasing the number of funds in a portfolio decreases skewness while increasing both kurtosis and correlation with the market. They summarize this as follows: "when adding more funds, the probability of a relatively large loss rises while the diversification potential within the context of a larger stock-bond portfolio drops," implying that, "with hedge funds, diversification is not necessarily a good thing" (page 5). Lhabitant and Learned (2004) concur; given their similar empirical results, they summarize their findings by saying that, "for some strategies, too much diversification results in undesirable side effects in the higher moments of the return distribution. Thus, while a fund of hedge funds may mitigate the negative effects of a hedge fund failure through diversification, too much diversification is also likely to result in deworsification" (page 2).

For fund portfolios, then, diversification is not a free lunch. Even where it seems to help—mean and variance—its benefits play out very quickly. Our model, which focuses on mean and variance, may help explain why. We will present the model after a brief review of the performance metrics we use.

ALPHA AND THE INFORMATION RATIO

Searching for alpha is often complicated by its multiple definitions—it is hard to find something you cannot describe. In this section, we briefly review some of the common definitions for alpha. We then justify using the ratio of alpha to active risk, the information ratio, as our portfolio manager's decision criterion.

Alpha is a portfolio's return above a benchmark. Benchmark specification therefore is central to the identification of superior returns. Also critical is the recognition that variations from the benchmark can occur both from luck and from manager skill. Thus, differences in the benchmarks chosen and varying acknowledgments of the effects of luck are the central points of contention among alpha definitions.

The earliest of the alpha measures was Jensen's (1967). Jensen defined this risk-adjusted return as the difference between a portfolio's average return and its predicted return using the Capital Asset Pricing Model (CAPM):

$$\alpha_p = \overline{R_p} - \{r_f + \beta_p * [\overline{R_M} - r_f]\}, \tag{1}$$

where r_f is the risk-free rate, $\overline{R_p}$ is the average return on the portfolio, $\overline{R_M}$ is the average return on the "market" portfolio, and β_p is the portfolio's beta coefficient, defined as $\sigma_{pM}/\sigma_M{}^2$ (see Bodie, Kane, and Marcus, 1993, page 804).

Since the CAPM is an ex ante model, composed of market and risk-free benchmarks that exist only in theory, making Jensen's alpha operational requires some heroic assumptions. We need a proxy for the market portfolio (a perfectly diversified portfolio that lives on the efficient set of risky assets), a proxy for the risk-free rate, and a way to measure beta (specifying return measurement intervals, length of measurement period, and any accommodation for potential mismeasurement).

For example, Morningstar suggests that investors use the S&P500 for the market and the 90-day Treasury bill for the risk-free rate. Thus, we are to substitute a 500-stock, domestic large-cap equity index for the ex ante efficient portfolio of all existing risky assets, and an inflation-vulnerable, domestic money-market rate for the riskless benchmark. Even if we accept that these proxies are reasonable, we have no guidance as to how specifically to use them to calculate an asset's predicted return, since we are not told specifically how to measure beta. Since our identification of excess return is dependent upon our calculation of the predicted return, these measurement problems complicate our application of Jensen's alpha.

Academic theoretical and empirical work on alpha has addressed these sorts of issues by increasing the complexity of the models used to define alpha and the econometric techniques used to identify it. For example, Avramov and Chordia (2006) model a K-factor specification for stock returns, utilizing a set of macroeconomic variables that may help predict them. In their empirical model, alpha may vary over time. Similarly, Kosowski, Naik, and Teo (2007) use a seven-factor return model to define alpha. They then use a seemingly unrelated regression approach to create Bayesian posterior alphas. These Bayesian alphas measure performance more accurately and are better at predicting future superior performance. Ordinary least squares (OLS) alphas, on the other hand, tend to overstate excess return while simultaneously prohibiting that excess performance from easily being distinguished from luck. Lo (2007), criticizing current performance metrics for being static measures of dynamic processes, proposes an "active/passive" approach that defines active risk using the covariance between portfolio weights and returns. He breaks the active component of returns into two parts: that coming from security selection (related to the intercept of a K-factor model, alpha, "identifying untapped sources of expected return" [page 14]) and that coming from factor timing (related to exposure to the model's factors). Nielsen and Vassalou (2004) also extend traditional performance measures from discrete to dynamic processes. They develop both the Sharpe ratio (discussed below) and Jensen's alpha in continuous time, under conditions that allow a fund manager to incorporate active strategies, including the use of options. They show that alpha in this case equals the discrete version, (1), minus half of the variance of the portfolio, plus half of its covariance with the market benchmark. Managers identifying a fund with positive alpha can improve their performance by adding some of the fund to their core portfolio.

Although academic characterizations of excess return have become increasingly sophisticated, these models have yet to be fully incorporated into practice. Instead, institutional work often relies on problematic conceptions of alpha. For example, statements that merely refer to alpha in passing with phrases such as "returns above and beyond those of the market" (Bonafede, *et al*., 2004, page 2) may lull us into thinking that alpha is simply ($\overline{R_p}$ - $\overline{R_M}$), increasing the risk that we will ignore beta effects when evaluating managers' above-market returns. Lo (2007) puts it this way: "active management is not simply adding value in excess of a passive benchmark, which can be done passively by taking on non-benchmark factor exposures in a multi-factor world" (page 9). (Thus, there is a third reason that a manager's returns can vary from a benchmark: not just luck, not just skill, but also exposures to non-market [priced] risk factors.)

More serious than this type of obviously simplistic, throwaway reference are the models that attribute all ex-market returns to alpha, ignoring error. For example, Bonafede *et al*. (2004) present the following single-factor model:

$$R_p = r_f + \beta^*(R_M - r_f) + \alpha. \qquad (2)$$

Viewed as an ex ante measure, or as an ex post relationship among averages, this correlates nicely with Jensen's measure, (1). However, (2) is not a fully specified empirical model. This distinction is clouded when the authors state that "[i]nvestment returns can be broken into three distinct parts," as shown in (2), and that this tripartite description implies that portfolio risk must therefore be:

$$\sigma^2_p = \beta^2_p{*}\sigma^2_M + \sigma_e^2, \tag{3}$$

where σ_e "is the variance [standard deviation] of the alpha component of return" (Bonafede, *et al*., 2004, page 4). σ_e^2 is actually active risk, and is composed not only of variation attributable to "pure" active risk, but also of random error. Waring and Siegel's (2003) work is similar, as is Foresti's (2005) and Foresti and Toth's (2006). For example, although Waring and Siegel explicitly recognize the role of luck, using the term "pure active risk" to distinguish "the portion of the return that is not explained by market or beta risk exposures of the portfolio" (page 2), they also present versions of equations (1) and (2). They also echo Bonafede *et al.*'s dichotomy when they state that risk can be broken into systematic and "pure active risk," and when they describe alpha as the residual from a return/benchmark regression. Where is luck here? Attributing all nonmarket risk to the variance of alpha ignores error. The whole point of performance measurement is to be able to identify those managers who beat the benchmark consistently—because of skill. Just beating the benchmark is insufficient. Expressions such as (2), when considered as an empirical model, ignore the distinction between luck and skill, missing the whole point of the quest for alpha.

While common institutional definitions of alpha may obscure the distinction between skill and luck, these types of definitions are unlikely to be the measure of an active manager's performance. Instead, managers are usually evaluated using reward-to-risk measures. Two commonly reported by portfolio consultants are the Sharpe measure and the information ratio (see, for example, Bonafede, *et al*., 2004, Appendix B). The Sharpe measure is defined as:

$$S_p = (\overline{R_p} - \overline{r_f})/\sigma_p, \tag{4}$$

(where expressing the risk-free rate as an average, $\overline{r_f}$, recognizes that this rate may vary over the sample period). This ex post measure relates excess return to total risk; the higher the ratio, the better the performance. The Sharpe ratio is the appropriate measure of performance when plunging into a single portfolio (see Bodie, Kane, and Marcus, 1993). This case is similar to identifying the Capital Market Line in (expected return, standard deviation) space: the goal is to choose the efficient risky asset that, when combined with the risk-free asset, produces the steepest opportunity set. In the theoretical derivation of the CAPM, this linear opportunity set defines the "market" portfolio, which is chosen by all market participants. In the manager-evaluation application, the steepest line identifies the best single risky portfolio to choose.

However, when an active portfolio is to be added to a diversified core portfolio, the total risk of the addition is not the appropriate risk measure. In this case, we use the information ratio (IR):

$$IR_p = \alpha_p/\sigma_{ep}. \tag{5}$$

The IR compares a portfolio's active excess return to its total nonsystematic risk. Bodie, Kane, and Marcus (1993) note that the square of the Sharpe measure for overall portfolio—the combination of the core and the active "satellite"—equals the square of the market's Sharpe ratio plus the square of the added portfolio's information ratio. (See also Nielsen and Vassalou, 2004.) The IR therefore measures the contribution of the addition to the performance of the complete portfolio.

Kosowski, Naik, and Teo (2007) use a measure similar to the IR, the t-statistic of alpha, to evaluate the performance of hedge fund managers. (This statistic is approximated by the IR multiplied by the square root of the number of observations, N: [IR* $\sqrt{N}$]. Note that the standard error of the alpha estimate is $\sigma_e*[\Sigma X^2/(n*S_{xx})]^{1/2}$, where S_{xx} equals $[\Sigma X^2 - (\Sigma X)^2/n]$; see Ott, 1984, page 284.) They find that this measure, compared to alpha alone, allows them to better distinguish skillful from lucky performance. They report that these results "empirically validate the industry practice of ranking fund managers on their information ratios as opposed to their Jensen's alpha." We will follow that practice.

In this paper, we use the IR to consider a combined core/satellite portfolio. This strategy is consistent with Wilshire Associates' suggested use of hedge funds in institutional portfolios (see Bonafede *et al.*, 2004), and with the exposition in Waring and Siegel (2003). Specifically, we consider the contention that maximizing the IR of an added portfolio of active managers is best achieved by diversifying across many active managers: by treating the satellite as we would any other portfolio, perhaps we can decrease active risk, σ_e^2, by diversifying across multiple alpha sources. If we can, the lowering of active risk through diversification would translate directly into higher information ratios, and therefore into superior performance.

However, the problem with diversifying across active managers is that you may actually be diversifying away from skill and toward luck. Not all managers are equally skillful. In fact, some are just lucky. As Waring and Siegel (2003) note, a portfolio manager who wants to hire active managers "must believe that he or she can identify which ones will be the winners" (page 20). If she cannot, she may be better off by not diversifying.

THE MODEL

Assume that there are three types of fund managers: passive managers (P), active managers who have no special skill (U), and active managers who are skilled (S). Thus, denoting the type of fund manager by T, we have T $\in$ (U, P, S). The proportions of the three types of fund managers in the market, w_T, must sum to one. We will assume that, among active managers, there are fewer who are skilled, so that $w_U > w_S$.

Regardless of the return model we adopt, the average risk-adjusted abnormal performance, $\bar{\alpha}$, must equal zero across the full market. Thus, we must have:

$$\bar{\alpha} = w_S*\bar{\alpha}_S + w_U*\bar{\alpha}_U + w_P*\bar{\alpha}_P = 0 \tag{6}$$

If we assume that the average excess return to the passive managers, $\bar{\alpha}_P$, itself equals zero, then (6) implies that:

$$\bar{\alpha}_U = -\bar{\alpha}_S*(w_S/w_U). \tag{7}$$

Since we assume that the skilled managers deliver a positive alpha, (7) says that $\bar{\alpha}_U$ will be negative. In the zero-sum game of delivering excess performance, the skilled managers are taking their excess from the unskilled.

To determine the variance of any portfolio of fund managers that a portfolio manager chooses, we must describe the covariance matrix for the returns for the various fund manager types. First, we will assume that all active managers' excess returns are distributed normally with a common variance σ_e^2, so that excess returns are distributed N($\bar{\alpha}_S$, σ_e) for skilled managers and N($-\bar{\alpha}_S*(w_S/w_U)$, σ_e) for unskilled.

(Throughout the remaining exposition of the model, we will drop the "e" subscript from σ_e for notational simplicity. Also, we will refer to managers' σ_e^2 simply as the "variance.")

These assumptions about the diagonal elements of the covariance matrix warrant some comment. Normality is a simplifying assumption: as Kosowski, Naik, and Teo (2007) note, hedge fund alphas will probably not be normally distributed, given the complex and often option-like strategies they employ. However, we are being consistent with institutional work, which also assumes normality. On the other hand, the institutions may not assume that the variances for the skilled and unskilled managers would be equal. For example, Waring and Sigel (2003) suggest that we should "prefer skillful lower active risk managers to higher risk, concentrated managers," implying that active managers' variances will depend upon their strategies. However, they do note that the "luck component" of risk is comparable for skillful and unskillful managers (page 6). This comparability is sufficient to motivate our simplifying assumption: we do not distinguish among strategies, and allowing for unequal variances would unnecessarily complicate the model.

For the off-diagonal elements of the covariance matrix, we will assume that skilled managers' returns are uncorrelated with each other and with those of the other two types of managers, so that $\rho_{SS} = \rho_{SU} = \rho_{SP} = 0$. This models the unique contribution of a manager who is truly skilled. For simplicity, we will also assume that unskilled active managers' returns are uncorrelated with the passive, $\rho_{PU} = 0$. However, they are positively correlated with each other: $\rho_{UU} > 0$. This models the observed behavior of active managers' following similar strategies (we discuss this point more fully below). Given our assumption that only ρ_{UU} is nonzero, we will refer to this correlation from now on simply as ρ.

We want to compare the active risk/return tradeoff between a concentrated, skilled portfolio and a diversified portfolio of active managers. For the skilled portfolio, we will use only one manager, denying us the diversification potential that would undeniably come from a portfolio of skilled, uncorrelated managers. (If it is difficult to distinguish luck from skill, and if skilled managers are scarce, we would be fortunate to identify our one skilled manager.) The variance of this skilled portfolio will therefore be σ^2, and we expect that the delivered information ratio will be $\bar{\alpha}_S/\sigma$.

For the diversified portfolio, a portfolio manager's expected alpha and variance will depend upon the proportion of unskilled managers she chooses. Calling this proportion w, and calling the total number of managers used n, we can describe the portfolio characteristics as:

$$\alpha_p = w^*(-w_S/w_U)^*\bar{\alpha}_S + (1-w)^*\bar{\alpha}_S \tag{8}$$

$$\sigma_p^2 = n^*(1/n)^{2*}\sigma^2 + n^*w^*(n^*w - 1)^*(1/n)^{2*}\sigma^{2*}\rho_{UU} \tag{9}$$

The first term in the variance expression (9) represents the contributions to variance from the n individual managers. The second term represents the contributions from the covariances among the returns of the n*w unskilled managers. All of the other portfolio covariance terms are zero, given our assumptions of independence.

How a portfolio manager performs depends upon her ability to identify skilled managers. The smaller is her w (the smaller her proportion of unskilled managers), the higher is her α_p, the lower is her σ_p^2, and therefore the higher is her information ratio. Focusing our attention on w allows us to model the second level of superior ability that is necessary for positive active returns: the ability of the portfolio manager to identify fund managers who can consistently deliver alpha.

If a portfolio manager has no special skill at identifying these fund managers, then we would expect that w would equal $w_U/(w_U + w_S)$, which makes α_p zero. (Thus, as Waring and Siegel, 2003, note, the portfolio manager should stick to passive investments if she has no skill at picking fund managers.) Values of w higher than this will clearly make the situation even worse, and α_p will be negative. However, as w approaches zero, α_p approaches $\overline{\alpha}_S$, and more importantly, σ_p^2 approaches σ^2/n. This is the dominating case: a diversified portfolio of skillful managers (the case to which Waring and Siegel undoubtedly refer).

The key to the choice between our concentrated portfolio and a diversified portfolio therefore lies in a portfolio manager's expectation about her realization of w, our proxy for her ability to identify skillful fund managers. A consideration of the factors that make a portfolio manager inherently successful in this identification is beyond the scope of this paper. However, by evaluating our model's implications about w, we can consider the market factors that imply a payoff to this identification.

We will begin our characterization of w by finding its break-even value, the value of w at which we are indifferent between a diversified portfolio and a concentrated one. This is most easily accomplished by equating the squares of the two strategies' information ratios:

$$(\overline{\alpha}_S/\sigma)^2 = [w*(-w_S/w_U)*\overline{\alpha}_S + (1-w)*\overline{\alpha}_S]^2/[n*(1/n)^2*\sigma^2 + n*w*(n*w - 1)*(1/n)^2*\sigma^2*\rho_{UU}] \quad (10)$$

Rearranging (10) leads to the following quadratic equation:

$$w^2*[n*\rho - n*r^2] + w*[2*n*r - \rho] + (1-n) = 0, \quad (11)$$

where r equals $(w_U + w_S)/w_U$, or the inverse of the proportion of active managers who are unskilled. Solving for w, we find the following unwieldy expression for w_U^{max}, the break-even proportion:

$$w_U^{max} = \{\rho - 2*n*r + \{\rho^2 - 4*n*[\rho*(r + 1 - n) - r^2]\}^{1/2}\}/2*n*(\rho - r^2) \quad (12)$$

By evaluating the comparative statics of this expression, we can investigate the environmental factors that influence the relative payoff to a diversified portfolio of fund managers. If these factors result in a break-even w that is too low, the portfolio manager may conclude that her skill level is insufficient to expect success in creating a superior, diversified portfolio. Before examining these relationships, however, we will first consider a simple numerical example.

<u>A Basic Example</u>

The values we will use for this example are presented in Table 1.

Table 1: Numerical Example

Variable	Example Value
n	10
σ	0.04
ρ	0.2
w_S	0.1
w_U	0.3
$\overline{\alpha}_S$	0.02

The table presents the variable values that we will use to demonstrate the influences from equation (12).

We should motivate these choices briefly. First, we assume that the portfolio manager will use 10 fund managers. As a "satellite" strategy to a passive core, the active portfolio should not be so large as to be unmanageable. Ten active funds is within the usual range for small- to medium-sized pension funds, and it is the number of funds at which Nesbitt *et al.* (2003) show the strongest benefit to diversification. However, the n variable, more than being an influence on the outcome of equation (12), is primarily a reflection of the active opportunities available to the portfolio manager. (Given that it might then be a function of the active strategy employed, there may be a functional relationship between n and ρ; we will ignore these types of relationships.)

The correlation between unskilled active managers' returns, ρ_{UU}, is positive. Similar trading strategies or holdings can induce correlations among fund returns, which will reduce the value of portfolio-level diversification. For the universe of thousands of hedge funds, Kosowski *et al.* (2007) assert that most of their return distributions are likely to be "sparsely overlapping" (page 230); similarly, Ross (2003) notes that, from 1993 to 2002, hedge fund portfolio strategies had risk that was "markedly lower than that for any individual strategy, confirming that the returns on the strategies are relatively uncorrelated" (page 7). However, this evidence does not preclude significant overlap among subsets of these returns. To quantify this relationship, then, we turn to Waring and Siegel (2003), who note that historical alphas have correlations falling usually between -0.2 and +0.2 (page 14; see also Ross, 2003, Table 3). This informs our choice of 0.2 for ρ_{UU}. These authors also note that "pure" alphas are often assumed to be independent; this is what we have done for all correlations involving skilled managers.

The standard deviation of 4% comes from a broad characterization used by pension consultants: 1% to 3% active risk implies a strategy like enhanced indexation; 4% and up means a truly active strategy (see, for example, Foresti *et al*., 2006, page 17). This is comparable both to Waring and Siegel's (2003) assertion that active risk usually falls in the range from 4% to 25%, averaging 5-6%, and to Ross's (2003) evidence that hedge fund risks tend to fall between 3% to 10.5% (see his Table 2). Similarly, the 2% $\overline{\alpha}_S$ was suggested by the premium above index that is a common benchmark for active managers. Together, these choices of $\overline{\alpha}_S$ and σ imply an information ratio of 0.5, the value Waring and Siegel (2003) say is the lower threshold for top-quartile active managers.

Finally, the .1 and .3 values for w_S and w_U, respectively, imply that the market in our example has 60% passive managers and 40% active. In the active space, there are three times as many unskilled managers as skilled, so that a random choice of active manager has a 75% chance of drawing someone unskilled. This is the ratio that the portfolio manager has to beat: she must choose her fund managers so that her realized proportion of unskilled active managers, her w, is lower than 75%.

How much lower? Given the values assumed in Table 1, equation (12) gives a break-even value for w of 47%. In a market where unskilled managers outweigh skilled by 3-to-1, our portfolio manager must be able to create her active satellite with at least 53% skilled managers. If she cannot, she is better off with a concentrated portfolio with one skilled manager. (Of course, this requires that she be able to identify at least *one* skilled manager. Given the distribution of fund managers' skills, even a random grab gives her a 25% chance of choosing one skilled manager. However, she would have only a 7.8% chance of choosing *five* or more of them.)

Having considered this basic example, we now consider the general implications of the model.

<u>Comparative Statics and Implications</u>

In this section we will review the effects on w_U^{max} of three variables: r, n, and ρ. To facilitate the discussion, we will simplify equation (12) by making the following substitution:

$$z \equiv \{\rho^2 - 4*n*[\rho*(r + 1 - n) - r^2]\} \quad (13)$$

This allows us to express w_U^{max} as:

$$w_U^{max} = \{\rho - 2*n*r + z^{1/2}\}/2*n*(\rho - r^2) \quad (14)$$

Expressing z as $\rho^2 + 4*n*\rho*(n-1) + 4*n*r*(r-\rho)$, and noting that $\rho \leq 1$ while $r \geq 1$, we can see that z is positive. In fact, $z > 1$, as is its square root. These relationships will help us sign the partial derivatives of w_U^{max}.

We begin by considering the effect of r on w_U^{max}. Remember that r is defined as $(w_S + w_U)/w_U$, so that higher values of r imply that a larger relative proportion of active fund managers is skilled. With higher values of r, it is easier to choose a skilled fund manager by chance, making it more likely that a portfolio manager would prefer a diversified portfolio to a concentrated one.

But what of the effect of r on w_U^{max}? Using (14), we see that:

$$\delta w_U^{max}/\delta r = [.5*z^{1/2}*(\delta z/\delta r) - 2*n]/(2*n*\rho - 2*n*r^2) + (\rho - 2*n*r + z^{1/2})*4*(n*r)/(2*n*\rho - 2*n*r^2)^2 \quad (15)$$

The denominator of the first term is negative. The first term will therefore be positive if $.5*z^{1/2}*(\delta z/\delta r)$ is greater than 2*n. (Note that $\delta z/\delta r$, which equals $[8*n*r - 4*n*\rho]$, is positive.) Simplifying, this requires that $(2*r - \rho)*z^{1/2} > 1$. We can verify easily that this inequality holds by substituting 1 as the minimum for r and the maximum for ρ. Thus, the first term in equation (15) is (+/-), or negative.

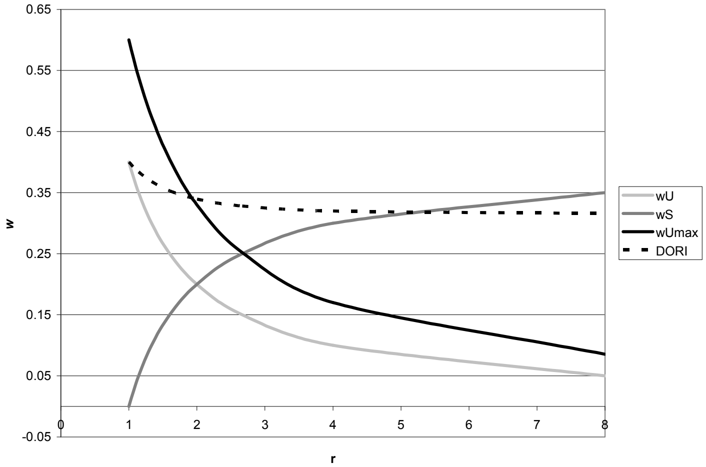
In the second term, the denominator is positive. Therefore, this term will be negative if $(2*n*r - \rho) > z^{1/2}$. Squaring both sides and simplifying, this leads to the requirement that $n*(r^2 - \rho) - n^2*(r^2 - \rho) < 0$. As this is obviously true, the second term in equation (15) is negative. Thus, $\delta w_U^{max}/\delta r$ is negative. This relationship between r and w_U^{max} is illustrated in Figure 1, using the values from Table 1.

Intuition suggests that, as the proportion of skillful managers in the active pool increases, a fund manager can more confidently diversify. On the other hand, we have just shown that as r rises, a portfolio manager has a *lower maximum* proportion for the number of unskilled managers that she can choose for a successful diversified portfolio. However, this negative relationship between r and w_U^{max} does not mean that she is less likely to diversify when r is high. Instead, it reflects the fact that the proportion of unskilled managers in the active pool is lower when r is higher. We can see this by looking at the "degree of required improvement" (DORI) which is also plotted in Figure 1. The degree of required improvement is the (absolute value of) the percentage difference between w_U^{max}, the maximum possible proportion of unskilled managers in a successful diversified portfolio, and the actual proportion of unskilled managers in the active pool. Thus,

$$\text{DORI} \equiv |[w_U^{max} - (1/r)]/(1/r)| \quad (16)$$

The lower is the DORI, the less skill a portfolio manager needs in order to benefit from diversification. If a manager's expected deliverable degree of improvement is less than DORI, she should not attempt to diversify. DORI falls as r increases. This means that having more skilled managers in the active pool puts less pressure on a portfolio manager to choose well—her job is easier, requiring less skill. It is, as we would expect, easier to create a successful diversified portfolio when the average skill level of active fund managers rises.

Figure 1: Effect of r on w_U^{max}



A higher value of r implies a larger proportion of skilled managers in the active pool (w_U falls and w_S rises). The maximum proportion of unskilled managers in a successful diversified portfolio (w_U^{max}) falls, but only because the total proportion of unskilled managers falls. Successful diversification is actually easier, as illustrated by the falling level of the degree of required improvement (DORI).

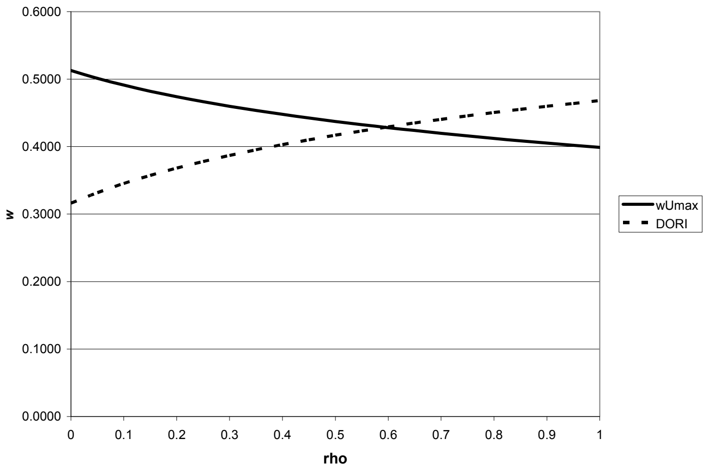
We now consider the effects of ρ and n on w_U^{max}. Both $\delta w_U^{max}/\delta\rho$ and $\delta w_U^{max}/\delta n$ are quite convoluted, so signing them directly is difficult. Fortunately for us, the intuition in these cases is straightforward: it is easier to diversify successfully when n is higher and ρ is lower. We will consider this intuition more fully after a brief discussion of the mechanics.

Looking at equation (14), we can see how ρ affects w_U^{max}. In the ratio for w_U^{max}, both the numerator and the denominator are negative. The derivative of the numerator with respect to ρ is $[1 + ½*z^{-1/2}*(\delta z/\delta\rho)]$. Given that n > (r+1), δz/δρ is positive. Thus, the numerator of w_U^{max} rises with ρ. The derivative of the denominator with respect to ρ is simply 2*n, obviously also positive. However, it is not immediately clear is how the ratio *w* behaves. We know that the denominator is larger in absolute value and that it increases linearly with ρ. However, the second derivative of the numerator with respect to ρ, $[4*z - (\delta z/\delta p)^2]/(4*z^{3/2})$, is negative. Over the relevant range, this causes the difference (relative to the denominator) between the two components to rise. The ratio w_U^{max} therefore falls.

We can observe this behavior using the numbers given above in Table 1. Setting all inputs to their Table 1 values except for ρ, we find that when ρ is zero, w_U^{max} is 0.5128. In this case, z is 71.11, the numerator of w_U^{max} is -18.23, and the denominator is -35.56. Raising ρ to 1, w_U^{max} falls to 0.3989. z is now 378.78, the numerator is -6.2, and the denominator is -15.56. The numerator rose by 66%, while the denominator only rose by 56%. This decreased the absolute difference between numerator and denominator, but increased the *relative* difference. The ratio, w_U^{max}, therefore fell. This behavior is clear from the graph in Figure 2.

The intuition informing the relationship between ρ and *w* is much more straightforward than the previous discussion would suggest. Increasing ρ means that diversification is more difficult, since unskilled active managers' returns are more closely related. A fund manager must have more skill in choosing active managers in this case—her realized w_U^{max} must be smaller. Thus, the degree of required improvement increases with ρ, as $\delta w_U^{max}/\delta\rho$ decreases.

Figure 2: Effect of ρ on w_U^{max}



If unskilled managers' returns are more closely correlated, it is harder to create a successful diversified portfolio. As ρ rises, the portfolio manager must choose a lower proportion of unskilled managers (lower w_U^{max}*)—delivering a higher degree of relative improvement—to succeed at diversification*

Now let us consider the effect of n on w_U^{max}. The partial derivative of w_U^{max} with respect to n is:

$$\delta w_U^{max}/\delta n = \{[-2*r + .5*z^{-1/2}*(\delta z/\delta n)]*(2*n*\rho - 2*n*r^2) - (\rho - 2*n*\rho + z^{1/2})*(2*\rho - 2*r^2)\}/(2*n*\rho - 2*n*r^2)^2 \quad (17)$$

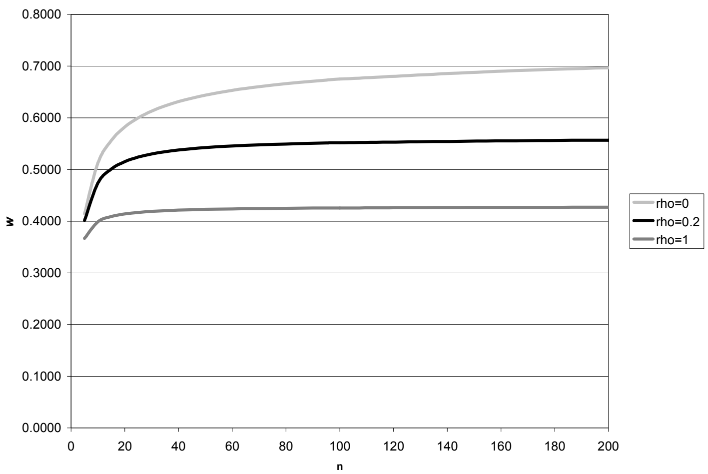
The partial of z with respect to n, δz/δn, equals $4*\{r^2 + \rho*[2*n - (r+1)]\}$. Since n > (r+1), δz/δn is positive, as are the z terms. However, every other bracketed term in the numerator of $\delta w/\delta n$ is negative, making this derivative hard to sign. Using numerical methods, we can determine that $\delta w_U^{max}/\delta n > 0$, while $\delta(w_U^{max})^2/\delta n^2 < 0$: w_U^{max} rises with n, but at a decreasing rate. The vast majority of n's effect on w_U^{max} comes at fairly low levels of n. In fact, once n exceeds about ten, there is very little marginal effect of increasing n on w_U^{max}, and w_U^{max} rapidly approaches its limit of:

$$\lim_{n \to \infty} w_U^{max} = (\rho^{1/2} - r)/(\rho - r^2) \quad (18)$$

However, as figure 3.A. shows, the value of having more managers depends upon the correlation among their returns. As we would expect, the lower the correlation, the greater the potential benefits of diversification, and the greater the effect of n on w_U^{max}. Considering the degree of required improvement from increasing n confirms these effects, as shown in Figure 3.B. The DORI falls with n, meaning that portfolio managers can more easily diversify successfully. However, this opportunity dissipates quickly as ρ increases.

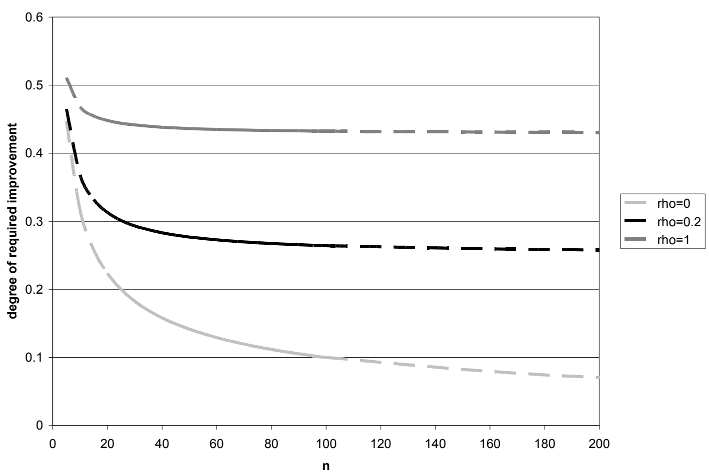
Figure 4 summarizes the major points of this section by plotting w_U^{max} against both ρ and r. The portfolio manager should diversify if she can keep her realized *w*, the proportion of unskilled managers she chooses, below the "ceiling" defined by the figure in the graph. If her realized *w* lies above this barrier, she would have been better off with a concentrated portfolio. Diversification is potentially most fruitful when the proportion of skillful managers in the active pool is relatively large, and when the returns of the unskilled managers are correlated less closely. Note that the correlation effect is particularly pronounced when r is low—when there are relatively more unskilled managers.

Figure 3.A : Effect of n on w_U^{max}



Successful diversification is easier (w_U^{max} is higher) with more funds, but this effect dissipates very quickly once n>10. Again we see that ρ is the crucial variable: when unskilled managers' returns are too closely correlated, increasing n does little to allow successful diversification.

Figure 3.B: Degree of Required Improvement with n

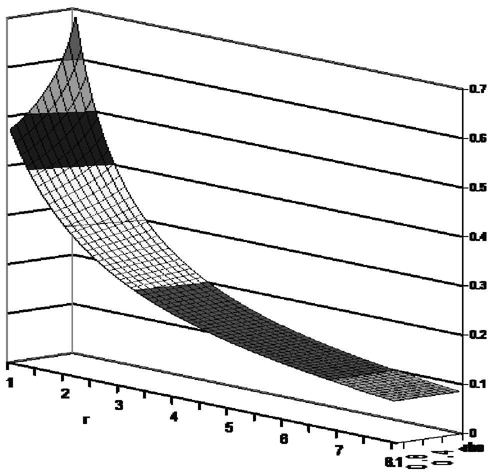


Successful diversification is easier (DORI is lower) as n rises, but higher correlations among unskilled managers' returns severely weaken this effect.

DISCUSSION

In a paper directly related to our model, Nesbitt *et al.* (2003) attempt to empirically determine the optimal number of hedge funds for a diversified portfolio. Using historical returns for 612 hedge funds, they form 1,000 equally weighted portfolios at each size *n*, where n ranges from 1 to 50 funds. They then compare the risk and return quartiles across sizes. Beginning with a one-fund standard deviation of 11.7% (the median annualized standard deviation for their sample), they show that portfolio risk falls to 6.5% when *n*=10. Further expanding the portfolios to 20 funds lowers risk marginally more, to 5.9%. Looking again at Figure 3.A., we see that, in our example, the beneficial effect of increasing the number of managers in a portfolio rapidly decreases above *n*=20, the same result that Nesbitt *et al.* find empirically. As noted earlier, however, their results lead them to the broad conclusion that it takes 20 to 30 individual funds to diversify optimally a hedge fund portfolio.

Figure 4: Effects of Both ρ and r on w_U^{max}



Successful diversification requires that the portfolio manager's proportion of unskilled managers falls below the illustrated "ceiling" (w_U^{max}). This will be easier when ρ is low and when there are more skillful managers in the active pool.

Here, then, is an institutional argument in favor of diversification. How do these empirical results inform our model, especially at the lower levels of *n*, where our conclusions may vary? Let us try to reconcile their results with ours.

Nesbitt *et al*. do not distinguish between skilled and unskilled managers. Assume, then, that the proportion of unskilled managers in the active pool is .75, while the skilled proportion is .25 (this will give us the same r value as we used earlier). Assume also that σ = .117. A portfolio manager randomly selecting ten funds from this universe would achieve a median portfolio standard deviation of 6.5% if ρ_{UU} = .5 (using equation (9), and assuming a binomial distribution from which fund managers are drawn). We will therefore take this ρ value as the model input implied by the authors' results.

Nesbitt *et al*. suggest that a skillful portfolio manager should be able to deliver top-quartile returns; in this case, such a result implies that σ = 5.85%. Using the binomial distribution, we see that a manager achieving this σ could choose at most six unskilled managers (a w_U^{max} of .6), meaning that she would need a degree of improvement of only .2. On the contrary, our model implies that this manager should concentrate: the implied w_U^{max} is only .44, with a DORI of .42. Given the high implied correlation among unskilled fund managers' returns, the model predicts little improvement from diversification.

The different conclusions of our model and Nesbitt *et al.*'s empirical results therefore hinge on the assumed correlations among active managers' returns. They acknowledge the importance of this input when they repeat their analysis on subsamples. Nesbitt *et al.*'s primary sample included all types of hedge funds, maximizing the potential for diversification (so that the actual ρ was almost certainly lower than our implied .5). However, when they separated funds by type, this potential fell. Macro manager benefited the most from diversification, "probably due to the non-correlated positions within their individual portfolios" (page 3). These authors summarize their findings by saying that "a good deal of diversification comes from mixing hedge fund styles, more so than from simply adding managers from within any one hedge fund style" (page 3). In the subsamples, then, their results are broadly consistent with ours—diversification is less beneficial when correlations are high.

Khandani and Lo (2007) echo this emphasis on correlations, and warn that increasingly correlated returns may decrease the potential for diversification across hedge funds in the future. Their study examines the

August, 2007 behavior of statistical arbitrage, quantitative equity market neutral, long/short equity, and 130/30 ("active extension") hedge-fund strategies. They show that the returns of these long/short strategy funds over this volatile period suggest that these types of funds are becoming much more highly correlated. For example, they note that "the necessarily quantitative nature of 130/30 strategies creates an unavoidable commonality between them and quantitative equity market-neutral strategies" (page 23). They note that if funds…

> …use similar quantitative portfolio construction techniques, then more often than not, they will make the same kind of bets because these techniques are based on the same historical data, which will point to the same empirical anomalies to be exploited… Moreover, the widespread use of standardized factor risk models… by many quantitative managers will almost certainly create common exposures among those managers to the risk factors contained in such platforms…. But even more significant is the fact that many of these empirical regularities have been incorporated into non-quantitative equity investment processes, including fundamental "bottom-up" valuation approaches like value/growth characteristics, earnings quality, and financial ratio analysis. (page 27)

Similarly, Lhabitant and Learned (2004), in explaining the significant increase in kurtosis and decrease in skewness from diversifying within fixed-income arbitrage and event-driven strategies, note that, "[i]t is our assumption that many of these managers have heavily invested in the same underlying assets, and are, therefore exposed to the same underlying risks…By diversifying among them, we are, in a sense, sure to capture these risks" (page 8). Correlation again is the enemy of successful fund-of-fund diversification.

Even the popular press acknowledges this increasing correlation. Describing the same period as Khandani and Lo, the *Wall Street Journal* noted that, "The reliance on models can be especially problematic because many quant hedge funds have very similar models. That means they are often doing the same trades and buying the same shares" (8/9/07). Again, the next day, the paper reported that, "Since market-neutral funds often are guided by similar computer models and share similar holdings, the actions magnified moves in asset prices. The last week has been the worst on record for many large hedge funds focusing on this strategy, worrying traders across Wall Street, many of whom look to these firms for signs of stability in difficult markets" (8/10/07; see also Cox *et al*., 9/4/07).

Khandani and Lo (2007) help quantify this correlation. In a summary comparison of the correlations among 13 hedge-fund indexes, 49 of these authors' coefficients were above .25, while 23 were above .50. Mean and median 36-month rolling-window correlations were between .50 and .60. (See their Figures 7 and 8.) Perhaps, then, our model's implied ρ of .5 was not so far off; perhaps diversification is not as efficacious at the fund level as it is at the security level.

But what of return? Nesbitt *et al.* suggest that, while expected return declines with n, it does so more slowly than risk. Diversification therefore benefits the portfolio manager.

However, their results track return and risk separately. Combining the median return measure and the median risk measure does not necessarily give the median information ratio. If we assume that a skilled manager delivers an average alpha of 2% and an unskilled manager delivers 0%, while continuing other assumptions above, we find that IR declines significantly as n increases: at the 25th percentile, IR is approximately .14, down from .54 when n=1 (however, the probability of achieving the IR of .54 is 0.0000009537). Considering both risk and return, diversification can impose a significant cost on a naïve portfolio manager.

In a consistent finding, Kosowski *et al*. (2007) report that the best and worst funds in their hedge fund sample are Directional Trader funds, which make aggressive bets on the market's direction. Either the big bets pay off, or they do not. On the other hand, the overall worst type of investment strategy is the fund-of-funds: even their best performances are less statistically significant than those of other strategies. In fact, these authors assert that "it is well known that Fund of Funds have lower average returns than individual hedge funds"—and not just because of the added layer of fees—a conventional wisdom that their empirical work verifies (page 250; see also Bondafede *et al*., 2004). Concentration may be better than diversification.

Nesbitt *et al*. admit this possibility when they discuss "slippage," which they define as the additional decrease in expected return that comes from combining managers in the face of decreasing portfolio manager skill. They explain that slippage could occur for many reasons: "...One is that it becomes much more difficult to identify good managers… Another could be the cost associated with researching and monitoring more managers. Fund sponsors' experiences with long-only managers suggest that slippage is more likely to appear as the number of managers increases" (page 5). Similarly, Foresti and Rush (2007) note that a manager's success at choosing active weights should fall as the number of these bets rises (page 12). These acknowledgments support the basic premise of our model: sometimes diversification attempts can do more harm than good.

CONCLUSION

Alpha is an enticing concept. Portfolio managers want to believe that they can add "juice" to their portfolios by adding active managers, just as active managers want to believe they can deliver it. However, the average excess return in the market must be zero. If there are fund managers who can add juice, there must also be those who drain it.

The job of the portfolio manager, should she chose to leave behind the relative safety of indexing, is to identify those fund managers who can deliver alpha because they are skillful, while avoiding those who may simply get lucky. Only the skillful fund managers will be worth what they cost. If a portfolio manager cannot distinguish luck from skill, she should stick with a passive portfolio.

However, even if she does believe that she can identify skillful fund managers, she may not be able to diversify profitably. Waring and Siegel (2003) prescribe that active fund managers should "[b]ias toward diversified portfolios, away from concentration," and that active portfolio managers should do the same: portfolio managers should "build efficient, or optimized, portfolios of managers just as they should of asset classes." However, these authors also assume that manager skill levels are "generally equal" (pages 19 and 20). What if they aren't?

According to Bonafede *et al.* (2004), the "Fundamental Law of Active Management" states that the information ratio is "approximately equal to skill times the square root of breadth." That is, current institutional practice suggests that we should be able to add breadth through diversification. However, if by doing so we sacrifice skill for size, we lose: skill, not size, is the input that enters the performance expression at full strength.

In this paper, we present a model of the choice between diversification and concentration. Three variables influence this choice: the number of funds used, the relative proportion of skilled managers in the active pool, and the degree of correlation in unskilled managers' returns. Of the three, correlation is the most potent. As fund managers' returns become more highly correlated—and recent evidence suggests this is happening—portfolio managers will need to exhibit more skill in selecting funds. If they cannot reliably separate the good from the lucky, portfolio managers may learn to prefer concentration to "deworsification."

REFERENCES

"Alpha Defined," Morningstar Investing Classroom, Course 205: Gauging Risk and Return Together, Part 1, (http://news.morningstar.com/classroom2/course.asp?docId=2932&page=2&CN=COM)

Amin, Guarav S. and Harry M. Kat, "Portfolios of Hedge Funds," Alternative Investment Research Centre, working paper #0003, 7/11/02

Avramov, Doron, and Tarun Chordia, "Predicting Stock Returns," *Journal of Financial Economics*, Vol. 82, 2006, page 387-415

Bodie, Zvi, Alex Kane, and Alan J. Marcus, *Investments*, 2nd ed., Irwin, Burr Ridge, IL, 1993

Bonafede, Julia K., Steven J. Foresti, and Thomas E. Toth, "Alpha: A Right or a Privilege? An Examination of the Challenges and Risks of Hedge Fund Investing," Wilshire Associates Incorporated, 10/14/04

Brands, Simone and David R. Gallagher, "Portfolio Selection, Diversification, and Fund of Finds: A Note," *Accounting and Finance*, Vol. 45(2), 2005, p.185-197

Cox, Rob, Jeff Segal, and Richard Beales, "Hedge Fund Correlation," breakingviews.com Financial Insight, *Wall Street Journal*, 9/4/07

Foresti, Steven J. and Michael E. Rush, "Short-Extension Strategies: To 130/30 and Beyond!" Wilshire Associates Incorporated, 8/16/07

Foresti, Steven J. and Thomas E. Toth, "Incorporating Active Management… Over Active Imagination," Wilshire Associates Incorporated, 6/28/06

Foresti, Steven J., "Institutional Use of Hedge Funds: Penetrating the Darkness on the Hedge of Town," Wilshire Associates Incorporated, 7/26/05

Jensen, Michael, "The Performance of Mutual Funds in the Period 1945-1964," *Journal of Finance*, Vol. 23(2), p. 389-416

Khandani, Amir E. and Andrew W. Lo, "What Happened to the Quants in August 2007?", working paper, MIT Sloan School of Management, 9/20/07

Kosowski, Robert, Narayan Y. Naik, and Melvyn Teo, "Do Hedge Funds Deliver Alpha? A Bayesian and Bootstrap Analysis," *Journal of Financial Economics*, Vol. 84, 2007, p. 229-264

Lhabitant, Francois-Serge and Michelle Learned De Piante Vicin, "Finding the Sweet Spot of Hedge Fund Diversification," EDHEC Risk and Asset Management Research Center, working paper, 4/04

Lo, Andrew W., "Where Do Alphas Come From? A New Measure of the Value of Active Investment Management," working paper, MIT Sloan School of Management, 5/15/07

Mikhail, Michael B., Beverly R. Walther, and Richard H. Willis, "Do Security Analysts Exhibit Persistent Differences in Stock Picking Ability?" *Journal of Financial Economics*, Vol. 74, 2004, p. 67-91

Nesbitt, Stephen L., Kevin O'Donnell, and Tyler Dritz, "The Optimal Number of Hedge Funds," Wilshire Associates Incorporated, 5/6/03

Nielsen, Lars Tyge and Maria Vassalou, "Sharpe Ratios and Alphas in Continuous Time," *Journal of Financial and Quantitative Analysis*, Vol. 39(1), March 2004, p. 103-114

Ott, Lyman, *An Introduction to Statistical Methods and Data Analysis*, 2nd ed., Duxbury Press, Boston, MA, 1984

Park, James M. and Jeremy C. Staum, "Diversification: How Much is Enough?" Social Science Research Network working paper, 3/13/98

Ross, Stephen A., "The Role of Hedge Funds in Institutional Portfolios," January 2003 (http://huangcapital.com/INST.pdf)

Sender, Henny and Kate Kelly, "Blind to Trend, Quant Funds Pay a Heavy Price," *Wall Street Journal*, 8/9/07

Statman, Meir, "How Many Stocks Make a Diversified Portfolio?" *Journal of Financial and Quantitative Analysis*, Vol. 22(3), p. 353-363

Waring, Barton and Laurence Siegel, "Understanding Active Management," *InvestmentInsights*, Barclays Global Investors, Vol. 6(1), April, 2003

Zuckerman, Gregory, James R. Hagerts, and David Gauthier-Villars, "Impact of Mortgage Crisis Spreads; Dow Tumbles 2.8% as Fallout Intensifies," *Wall Street Journal*, 8/10/07

CAUSALITY BETWEEN TAX REVENUE AND GOVERNMENT SPENDING IN MALAYSIA

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ABSTRACT

The trend of tax collection in Malaysia is inconsistent, changing upward and downward depending upon economic conditions. However, over a 30 year period, most years show an increasing increment in total collection. The exceptions are when there is an abnormal economic condition such as financial crisis, war or increase in world oil prices. Total tax revenue has always been a major contribution to Malaysia's federal government revenue. Income tax is one of the surest ways to fund the government. The main objective of this study is empirically tests the causality between tax revenues and government spending in Malaysia for the past 36 years by applying an econometrics model. The results provide evidence for the existence of a long-run relationship between tax revenues and government spending with unidirectional and bidirectional causality in VAR models for the sample period 1970-2006.

JEL: C01, H20, H59

INTRODUCTION

Taxes in Malaysia have been, and still are an important source of government revenue and the most dependable source of government funding. In many countries, tax relief has become a significant tool to boost the economic growth. In fact, taxation policy itself is a fundamental element for economic policies, ensuring that countries are able to maintain and improve its global competitiveness and to expand. This applies to developed, developing and countries in transition. The attractiveness of the tax system structure is important to ensure that it able to attract domestic and foreign investors. Hence, the decision of the Malaysia government to change its system of indirect tax from sales and service tax to Good Service Tax (GST) is an interesting economic issue.

In Malaysia the dependency on tax as a source of income is unquestioned. Taxation has been used as the main policy instrument for transferring resources to the public sector. This was demonstrated with the expanded role of the Inland Revenue Board as a tax collection agent for the government. This agency was transmitted from department to board on 1 March 1996, and was established in accordance with the Inland Revenue Board of Malaysia Act 1995. This act gives the Board more autonomy especially in financial and personnel management to improve the quality and effectiveness of tax administration.

Total government revenue in Malaysia is derived from two sources, which can be classified as tax revenue, and non-tax revenue. The responsibility for collecting tax revenue falls on IRB itself and Royal Customs and Excise Department. The responsibility for the collection of non-tax revenue, is based on the type of income. Direct tax revenue consists of income tax from individuals, companies, and other persons as well as petroleum, stamp duty, estate duty and real property gains. Indirect taxes are collected by the Royal Customs and Excise Department consists of tax revenue, and are not imposed directly on the taxpayer. Since the 1960's, indirect tax has became the major contribution to government revenue. Indirect taxes consist of import duties, export duties, excise duties, sales tax and service tax. Non-tax revenues consist of fees for issue of licenses and permits, fees for specific services, proceeds from the sale of government assets, rental of government property, bank interest, returns from government investments fines and forfeitures.

Spending increases every year in Malaysia as well as throughout most of the world. However, the question here is whether sufficient resources are available to fund these expenditures. Careful budgeting is critical and, a good fiscal policy is vital to stimulate a stable economy. Fiscal policy in Malaysia can be described as expansionary fiscal policy where there is always an increase in spending and lower taxes. The government always provides better incentives to both individual and company taxpayers. The Malaysian government spends public money to provide a wide variety of facilities and benefits to the public. From federal government reports we can classified the spending to two major categories, current expenditure and development expenditure. Current expenditure consists of emoluments, supplies and services, asset acquisition and routine expenditures. Development expenditure varied from economic services, social services, security and general administration.

Based on the neoclassical and endogenous growth model, growth or changes in government spending have driven revenue collection from both tax sources and non-tax sources. This implies that there is a unidirectional relationship between the dependent and independent variable. It suggests that every increase in revenue collection should lead to an increase in government spending especially in the short term. While during long-term frameworks, there is inconsistency in the relationship between the variables. In this case the level of dependency varies based on the situation.

TAXATION AND ECONOMIC POLICIES: THE CASE OF MALAYSIA

The Malaysian government has focused on development since the First Malaysia Plan (1966-1970). The Plan's objectives were to promote the welfare of all citizens, and improve the living conditions in rural areas, particularly among low-income groups. Currently the Ninth Malaysia Plan (2006-2010) is in place which highlights issues of current importance specifically infrastructure, health, environment, agriculture, education, culture and arts and heritage. The transition from plan to plan shows an expanded growth where every year that revenue was increased together with the raise of development boosting expenditures.

The Inland Revenue Board (IRB) carries out tax collection as well as policy implementation, monitoring and evaluation. This board plays a significant role to ensure the successfulness of any tax policy. Average percentage of increment in total collection starting from 1970 to 2006 for tax-revenue and non tax-revenue is 9.29% and 9.09% respectively (Federal Government Financial Position, 2007). The trend of collection is inconsistent with upward and downward movements depending upon the economic realities of a particular year. However over a period of 30 years most years show an increment in total collection except for where there is an abnormal economic condition. These increases are attributable primarily to strong of economic growth especially in the business sector.

During the period of the Seventh Malaysia Plan (7MP), the federal government revenue registered a moderate increase of 4.0 per cent, amounting to RM301.3 billion surpassing the revised target of 3.0 per cent (8MP). This was primarily credited to higher collection from tax and non-tax revenue, in tandem with impressive economic performance and the recovery in aggregate demand in 1999-2000 from a severe contraction that occurred in 1998. However, because of the financial crisis in 1999 there is a downward trend in tax collection from all categories of direct tax. Overall, the performance of the economy was commendable during the Plan period where total government revenue showed a moderate increase allowing the government to implement programmed projects on schedule.

In order to foster economic efficiency the Eighth Malaysia Plan (8MP) was design to focus on achieving sustainable growth with resilience. However, 8MP began with the slowdown in the global economy caused by the downturn of the US economy and collapse in world electronics demand. This was aggravated by the September 11 incident in the US in the same year. After a short breather, the events in the first half of 2002, particularly the invasion of Iraq and the outbreak of the Severe Acute Respiratory

Syndrome (SARS) again negatively affected economic recovery. Despite world economy uncertainty, revenue collection growth was projected with cumulative for five-year collection around RM470,000 million. As revenue collection increased, government spending also showed an increase. However unexpected results show the overall deficit had dropped from a high of 5.8% of GDP in 2000 to 3.8% by the end of 2005 (Economic Report, 2005/2006). Again, it is expected that in 2007 economic growth will remain strong with GDP estimated to grow again at a 5.8% rate.

Table 1 draws on public sector accounts to illustrate the financial position of the Malaysia government over a period of 2001 to 2007. Total revenue ranged from RM79.5 billion in 2003 to an expected RM134.8 in 2007. These increases are important to ensure that government is able to control the deficit. Development expenditures reported a reduction of 20.9% (Economic Report, 2004) in line with government objective to reduce the deficit. In fact, the continued weakening in the external environment and its adverse impact on domestic private demand necessitated the Government to pursue a stronger expansionary fiscal stimulus to revive domestic economic activities.

Table 1: Federal Government Finances, 2001-2007 (in billions)

	2001	2002	2003	2004	2005	2006	2007[a]
Total Revenue	79.5	83.5	92.6	99.4	106.3	123.5	134.8
Current Expenditure	63.7	68.6	75.2	91.3	97.7	107.7	113.0
Development Expenditure	34.2	35.0	93.4	28.9	30.5	35.8	44.5
Overall	-18.4	-20.2	-21.0	-19.4	-18.7	-19.1	-20.2

This table shows statistics of Malaysian Government finance from 2001-2007. Figures are in billions. [a] Forecast Source: Department of Statistics Malaysia, 2007

For detailed breakdown of government expenditure, Table 2 and Table 3 presents exact spending figures. For every year the government reported an increase in spending there was a corresponding increase in collection. In 2007 a total of RM157,496 million was budgeted for both operating and development expenditures. The increased expenditure was aimed at sustaining growth momentum given the more challenging external environment (Economic Report, 2006). Operating expenditures receive high allocation due to the compulsory nature of the expenditure. Whereby, the development expenditure is not based on necessity.

Table 2: Federal Government Development Expenditures by Sector, 2001-2007 (in millions)

	2001	2002	2003	2004	2005	2006	2007[a]
Agriculture and rural development	1,366	1,470	1,620	2,671	2,482	3,681	4,157
Trade and industry	1,870	2,084	3,456	1,629	3,221	3,791	5,102
Transport	1,671	1,887	7,354	7,014	7,660	6,198	7,298
Education and training	14,422	16,111	10,193	4,494	3,736	5,175	7,941,
Health	4,680	4,742	2,681	2,646	1,220	1,297	1,629
Housing	397	131	1,928	1,320	1,082	1,895	2,153
Security	3,531	3,587	6,029	4,338	4,803	5,781	6,817
General administration	8,635	7,021	1,824	2,670	3,324	3,556	2,648

This table shows Malaysian government expenditures. [a] Forecast Source: Department of Statistics Malaysia, 2007

LITERATURE REVIEW

Discussion and studies on the determination of growth remain one of the most significant and popular topics for economists and policy makers. Different studies indicate that different factors contribute to growth. However many believe that tax levels are one of the most significant factors that contribute to country growth. One tenet of taxation is its distorting effect on economic behavior either directly or indirectly. Moreover most developed and developing countries depend on the revenue collection for development purposes. The discussion concerning the relationship between tax and growth in either the

Table 3: Federal Government Operating Expenditures by Object, 2001-2007 (in millions)

	2001	2002	2003	2004	2005	2006	2 007[a]
Emolument	17,443	18,232	21,721	21,814	25,587	25,089	25,815
Debt service charges	9,634	9,710	10,543	11,655	11,604	12,726	13,127
Grant to state governments	2,012	2,189	2,125	2,960	2,614	2,907	3,645
Pension and gratuities	4,711	4,486	5,870	6,174	6,809	6,638	7,049
Supplies and services	10,703	11,854	13,968	18,133	17,984	21,608	23,147
Subsidies	4,552	4,646	2,679	6,250	13,387	11,251	11,908
Grant to statutory bodies	5,312	5,828	6,844	7,153	8,289	9,183	9,854
Refunds and write-off	1,776	2,376	2,657	5,113	288	357	1,815
Others	4,575	4,224	8,814	12,272	11,182	15,615	16,626

This table shows Malaysian government operating expenditures. [a] Forecast Source: Department of Statistics Malaysia, 2007

short run or long run has been widely discussed by economist with the finding that there is significant relationship between tax revenue and spending growth in the long run. Loganathan and Taha (2007) findings have demonstrated a consistent relation between revenue and spending. In fact government expenditures namely those that enter as input into the production function for final output and those that enter as inputs in investment technologies may well have large impact on long-run growth (Glomm and Ravikumar, 1997). On the other hand, they believed that changes in policy have significant implications since government expenditures in dynamic general equilibrium models may influence long run-growth rates and welfare.

Regardless of country size, tax has become a dominant factor in country endogenous growth, either from direct or indirect sources. From the revenue view, Wang (2007), Padovano and Galli (2002), and Brown (2002) argued that tax has a significant impact on economic growth. In fact Padovano and Galli verify the robustness of the correlation between tax variables and growth by progressively include additional policy and control variables in the growth regression. Supporting this contention, Erbil (2001) shows that trade tax has become a significant source of government revenue. Meanwhile, Hsing (2005) believed that economic growth is positive related with the growth in civilian employment, investment spending, technological progress, and human capital. Contrary to this finding Cantley (2004) argues that action to control spending only spurs the economy at least in the short run, but no evidence is provided for a long run relationship. The implication for changes in tax policy to either increase or decrease the taxpayer burden cannot be observed in the short run.

Researchers have conducted extensive tests on the relationship between economic growth and the structure of taxes. For instance, Koch et al. (2005) by using time series analysis for the period of 1960-2002 examined the implications of tax policy on economic growth by using a two-stage modeling technique. The authors found that decreased tax burdens are strongly associated with increased economic growth potential. In addition, contrary to most theoretical research, decreased indirect taxation relative to direct taxation is strongly correlated with increased economic growth potential. Koch et al. (2005) exploit the exogenous variation of United States enterprise zone policies to estimate the impact of geographically targeted tax incentives on a number of dimensions of local economic growth by using economic analysis. The results offer empirical evidence that incentives have a more complex dynamic in net growth performance. Similarly Yamarik (2000) includes the role of tax distortions in explaining state-level economic growth through the estimation of disaggregated personal income, general sales and property tax rates. As a result, disaggregated tax rates generate predictions more consistent with growth theory.

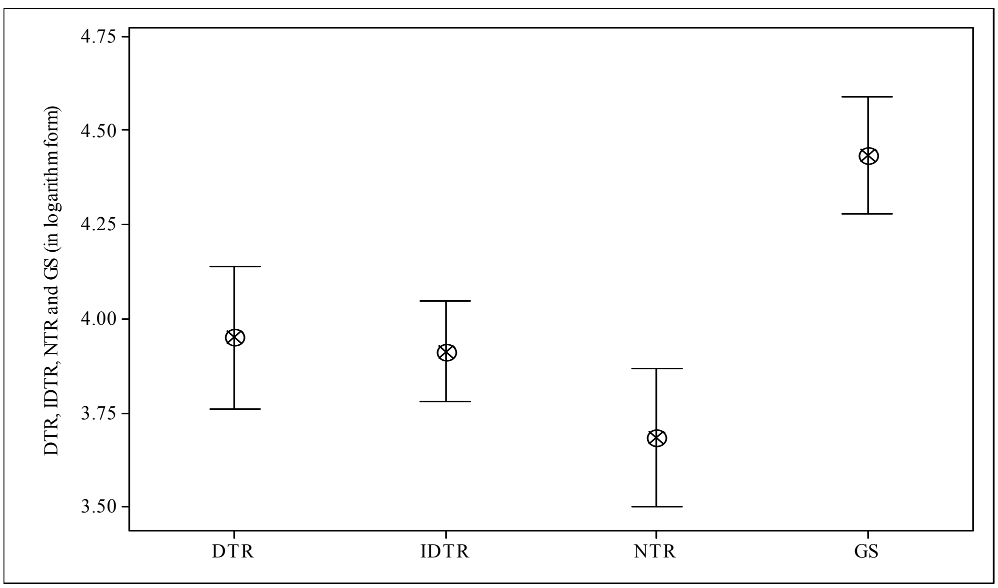
The determination of the causal ordering between these two macroeconomic aggregates has important implications for fiscal policy and the concomitant determination of budgetary balances. This is particularly true for countries that participate in the euro zone and thus fall under the provisions of the stability and growth pact (Kollias and Paleologou, 2006). Besides, an increase in monetary growth associated with an increase in government consumption results in crowding out both private consumption and private investment. This in turn reduces the growth rate of the economy (Ghosh and Mourmouras, 2001). Specifically the authors demonstrate that capital projects, financed through money creation namely; shortage of private savings and inadequate tax systems in many developing countries will give positive impact on economic growth. Furthermore, changes in policy will give significant impact on revenue elasticity besides changes in real income growth and inflation (Creedy and Gemmell 2004).

Various other realized and unrealized considerations mitigate growth. The standard static trade model cannot explain the large effect on growth in the long run. However this relationship could be explained by using more than one type of analysis (Cunat and Maffezzoli, 2007). Farmer and Lahiri (2006) drew attention to the fact that investment ratios are strongly correlated with growth across countries as well as saving ratios within countries. Similar with our method of study, Sinha, (1998) found the relationship between GDP and government expenditure in Malaysia by using Penn World Table annual data for 1950-1992. He uses two types of analysis with one methodology finding that there is a long run relationship between both variables namely GDP as independent and government expenditure as dependent variables. However, by using Granger-causality no evidence of a relationship is found. This is consistent with our study where taxation remains a causality effect in the long run, and therefore taxation policies always reflect on government spending.

THE DATA AND MODEL SPECIFICATIONS

This study utilizes yearly direct tax revenues, indirect tax revenues, non-tax revenues and government spending of Malaysia, covering the sample period of 1970 to 2006, with 36 observations on each of the variables. The International Monetary Funds (IMF) obtains the data from Malaysia's Department of Statistics (DOS) and World Development Indicators Database. Prior to the analysis, all variables are transformed into logarithm form. A graphical depiction of the data is provided in Figure 1:

Figure 1: Interval Plot of DTR, IDTR, NTR and GS, 1970-2006



Unit Root Tests

It is important to determine the characteristics of the individual series before conducting the cointegration analysis. Many studies have shown that majority of macroeconomics time series are not stationary, rather they are stationary with a deterministic trend. This creates a problem for econometricians as the normal properties t-statistics and Durbin Watson statistics (DW) and measures such as R-square results are biased when data is non-stationary. To test the order of integration, we used Augmented Dickey-Fuller test (ADF), Phillip-Perron test (PP) and Kwiatkowski et al. test (KPSS). It is widely acknowledged that ADF and PP tests are not very efficient in distinguishing between a unit root and a near unit root case. To complement ADF and PP tests, we employ the KPSS test proposed by Kwiatkowski et al. (1992). The KPSS test assumes that the null is stationary against the alternative that the variable does have a unit root.

Johansen and Juselius Cointegration Test

Granger (1969) proposes the concept of cointegration and, Engel and Granger (1987) provide further in depth discussion of the technique. The components of the vector X_t are said to be cointegrated of order *d*, *b*, and denoted by $X_t \sim CI(d,b)$ if (i) X_t is $I(d)$ and (ii) there exists a non-zero vector α such that $\alpha' X_t \sim I(d-b)$, $d \geq b \geq 0$. The vector α is called the cointegrating vector. Cointegration suggests that there exists a long-run equilibrium relationship linking these variables, or they tend to move together over time. Therefore, cointegration reveals long-run effects between time series variables. In this study, we employ Johansen and Juselius (JJ) cointegration test. The JJ cointegration approach suggests an alternative method to perform the cointegration test. Basically, the JJ method is presently used and takes the form of following equation:

$$\Delta Y_t = \prod Y_{t-1} + \sum_{i=1}^{p-1} \Gamma_i \Delta Y_{t-i} + BX_t + \varepsilon_t \qquad (1)$$

Where, $\prod = \sum_{i=1}^{p} A_i - I$, $\Gamma_i = -\sum_{j=i+1}^{p} A_j$, Y_t is a *k*-vector of non-stationary $I(1)$ variables, X_t is a *d*-vector of deterministic variables, and ε_t is vector of white noises with zero mean and finite variance. The number of cointegrating vectors is represented by the rank of the coefficient matrix Π. Johansen's method is to estimate the Π matrix in an unrestricted form, then test whether one can reject the restrictions implied by the reduced rank of Π. The likelihood ratio (LR) test for the hypothesis that there are at most *r* cointegration vectors is called the trace test statistic. It is to be noted that the variables under consideration should have identical orders, and in particular are integrated of order one (Engle and Granger, 1987). Testing for cointegration of the type CI(*d*,*b*) for $b<d$ are not of primary interest, since for $b<d$ the cointegrating vector is not stationary and does not have a straightforward economic interpretation (Charemza and Deadman, 1997).

Granger Causality Tests (VAR Approaches)

When in a regression equation we say that the explanatory variable X_t affects the dependent variable Y_t we indirectly accept that variable X_t causes variable Y_t, in the sense that changes in variable X_t induce changes in variable Y_t. This is in simple terms the concept of causality. With respect to the direction of causality, we can distinguish the following cases:

a) Unidirectional causality: This is the case when X_t causes Y_t, but Y_t does not cause X_t.
b) Bidirectional causality: This is the case when variables X_t and Y_t, are jointly determined.

In most cases, the direction of causality is not known and various tests have been suggested to identify the directions. The most well known test is the one proposed by Granger (1969). This test being based on the premise that the future cannot cause the present or the past utilizes the concept of the Vector Autoregressive model (VAR). Let us therefore consider the two variables, X_t and Y_t VAR (k) model:

$$Y_t = \alpha_{10} + \sum_{j=1}^{k} \alpha_{1j} X_{t-j} + \sum_{j=1}^{k} \beta_{1j} Y_{t-j} + \varepsilon_{1t} \qquad (2)$$

$$X_t = \alpha_{20} + \sum_{j=1}^{k} \alpha_{2j} X_{t-j} + \sum_{j=1}^{k} \beta_{2j} Y_{t-j} + \varepsilon_{2t} \quad (3)$$

With respect to this model, we can distinguish the following cases:

a) If $\{\alpha_{11}, \alpha_{12}, ... \alpha_{1k}\} \neq 0$ and $\{\beta_{21}, \beta_{22}, ... \beta_{2k}\} = 0$, there exists a unidirectional causality from X_t to Y_t, denoted as X→Y.
b) If $\{\alpha_{11}, \alpha_{12}, ... \alpha_{1k}\} = 0$ and $\{\beta_{21}, \beta_{22}, ... \beta_{2k}\} \neq 0$, there exists a unidirectional causality from Y_t to X_t, denoted as Y→X.
c) If $\{\alpha_{11}, \alpha_{12}, ... \alpha_{1k}\} \neq 0$ and $\{\beta_{21}, \beta_{22}, ... \beta_{2k}\} \neq 0$, there exists a bidirectional causality between X_t to Y_t, denoted as X ◄–► Y

In order to test the hypotheses referring to the significance or lack thereof, the sets of the coefficients of the VAR model equations (2) and (3) the usual Wald F-statistics could be utilized. The hypotheses in this test may be formed as follows:

H_0: X does not Granger cause Y, i.e $\{\alpha_{11}, ... \alpha_{1k}\} = 0$, if F_c<critical value of F-statistics
H_1: X does Granger cause Y, i.e $\{\alpha_{11}, ... \alpha_{1k}\} \neq 0$, if F_c>critical value of F-statistics

EMPIRICAL RESULTS AND DISCUSSIONS

Table 4 summarizes the outcome of the ADF, PP and KPSS tests on all four variables in this study. The null hypothesis tested is that the variable under investigation has a unit root against the alternative that it does not. In each case, the lag-length is chosen using the Akaike Information Criteria (AIC) and Kwiatkowski et al. (1992) after testing for first and higher order serial correlation in the residuals. In the first half of Table 4, the null hypothesis that each variable has a unit root cannot be rejected by both ADF and PP tests, but the KPSS test rejected the unit root hypothesis:

Table 4: Unit Root Tests

Variables	ADF Test (τ)		PP Test (Z_τ)		KPSS Test (η)	
	Level	First Differences	Level	First Differences	Level	First Differences
GS	-1.53(2)	-3.77(1)*	-2.27[0]	-4.22[0]*	0.14[4]	0.11[1]
DTR	-1.68(0)	-4.78(0)*	-1.64[2]	-4.79[1]*	0.13[4]	0.06[1]
IDTR	-2.18(0)	-5.43(0)*	-2.34[3]	-5.43[1]*	0.15[4]**	0.05[3]
NTR	-1.38(1)	-8.50(0)*	-0.82[2]	-8.47[1]*	0.15[5]**	0.08[2]

This table shows the results of three different unit root tests. Figures in parentheses indicate the lag length. Asterisks (*) and (**) denote statistically significant at 1% and 5% significance levels, respectively.

However, after applying the first difference, both ADF and PP tests reject the null hypothesis, but the KPSS test does not. Since the data appear to be stationary by applying the ADF and PP tests in first differences, no further tests are performed. We, therefore, maintain the null hypothesis that each variable is integrated of order one *I*(1). Given the results of unit roots, we now apply the Johansen techniques to test for cointegration between the variables within a VAR model specification. The results of testing for the number of cointegrating vectors are reported in Table 5, which presents both the maximum eigenvalue $(\lambda_{\text{Max-Eigen}})$ and the trace statistics (λ_{Trace}).

The cointegration results in Table 5 are obtained using a VAR specification where the variables and the cointegration space contain linear trends and the results does not indicates any cointegrating vectors. Therefore, the results in Table 5 indicate that the results is now completely identified a long-run relation equilibrium relationship and, indicating the speed of adjustment of each variable to the long-run equilibrium states. In order to examine the long-run causal relationship, we test for Granger-causality using block exogeneity Wald tests and report the results in Table 6.

Table 5: Johansen's Cointegration Tests

Null Hypotheses	Maximum Eigenvalue	$\lambda_{Max-Eigen}$ [k=1,r=0]	Critical Value (5%)	Critical Value (1%)
r=0	0.45	21.36	27.07	32.24
r≤1	0.29	12.35	20.97	25.52
r≤2	0.18	7.11	14.07	18.63
r≤3	0.09	3.38	3.76	6.65
Null Hypotheses	**Maximum Eigenvalue**	λ_{Trace} **[k=1,r=0]**	**Critical Value (5%)**	**Critical Value (1%)**
r=0	0.45	44.21	47.21	54.46
r≤1	0.29	22.84	29.68	35.65
r≤2	0.18	10.49	15.41	20.04
r≤3	0.09	3.38	3.76	6.65

This table shows the Johansen's Cointegration test results. Asterisks (*) and (**) denote statistically significant at 1% and 5% significance levels, respectively.

Table 6: VAR Granger-Causality (Block Exogeneity Wald Tests)

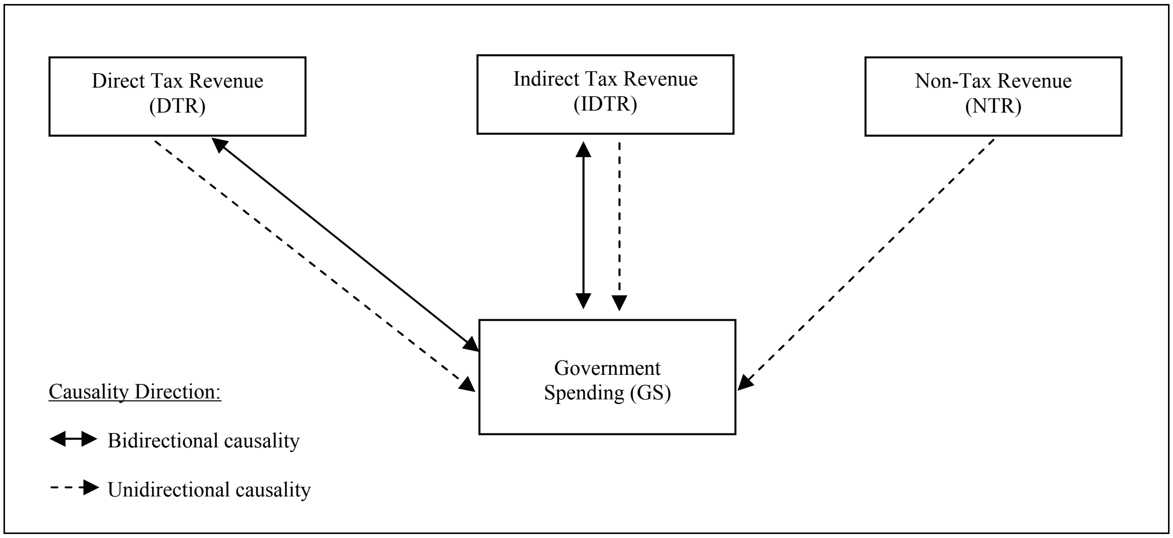
	GS	DTR	IDTR	NTR
GS	_	0.51	4.77*	1.52
		[0.47]	[0.02]	[0.21]
DTR	8.83*	_	38.53*	1.64
	[0.00]		[0.00]	[0.20]
IDTR	0.24	0.45	_	0.69
	[0.62]	[0.50]		[0.40]
NTR	1.57	0.00	0.05	_
	[0.21]	[0.98]	[0.81]	

*This table shows the VAR Granger-Causality test. *(**) indicates statistical significance at the 1% ,(5%). Figure in [ ] stands for probability value.*

The Granger-causality tests conducted above are conducted using a joint F-statistic for the exclusion of variable from one equation as illustrated above in a simple matrix form. The results of these tests indicate that Granger-causality is running in both directions between government spending and tax revenues. Thus, in contrast with the neo-classical argument that tax revenues is neutral to growth, our results for Malaysia are consistent with the view that direct and indirect tax revenues does have a causal impact on government spending. Our results are also in line with findings by Cunat and Maffezzoli (2007), Creedy and Gemmell (2004), Koch et al. (2005), and Sinha (1998) who obtained similar results on other countries. Figure 2 clearly shows the Granger-causality directions between government spending and tax revenues in Malaysia.

Therefore, any changes in government spending may lead to the changes in both tax revenue and vice versa. The reason for this is when the government is able to effectively collect tax revenue, the tendency to increase the government spending is very high. While in the case of spending, a good spending will project a good growth and influence the economy. Therefore, when the economy is in an upward trend the tendency to collect more tax is very high. When government issues any new policy that can increase or decrease direct tax collection especially in individual and companies' collection, this will also influence the indirect tax collection. For example, when the government raises public worker salaries the purchasing power of this party will also increases, indirectly affecting the sales tax or service tax collection. Therefore, there is bidirectional causality relationship between government spending, direct tax revenue and indirect tax revenues.

Figure 2: Granger-Causality Effects



CONCLUSIONS REMARKS

Since in the 1970s, the Malaysian government played a key role in the economy. The government ventured beyond its traditional functions and took on a more direct and active role in the country's overall social and economic development process. The establishment of collector body (IRB) as an agent of government collector has contributed to the successfulness of revenue collection. In addition, year-to-year the government has been able to increase the revenue collection as well as reduce the budget deficit. However, there is a need for empirical tests regarding the relationship between revenue collection and growth in order to see how both components of government finance interact. The determination of the causal ordering between these two macroeconomic aggregates is vital to ensure enactment of appropriate tax policies and the effectiveness of fund management. Most of the results show that changes in tax policy or structure has positive effect with growth and spending.

The purpose of this study was to test for Granger-causality between government spending and tax revenues for Malaysia. This study finds that there was bidirectional Granger-causality running from direct tax revenues, indirect tax revenues to government spending, but no unidirectional Granger-causality running between non-tax revenues and government spending. This supported the results of financial analysis in Malaysia where direct tax has become the major portion of government revenue followed by indirect and non-direct tax revenue. These results indicate that reducing direct and indirect tax rates may lead to a fall in government spending in the future, but these results may suffer from the omission of other relevant variables. In addition, non-tax revenues seem to be a less important contributor to the successfulness of country's growth as compared to direct and indirect tax. Therefore, future research should attempt to incorporate more variables in the analysis.

REFERENCES

Brown, D. (2002). "Canada's Conundrum: Tax and Spending," *Canadian Tax Journal*, vol. 50(4) p.1322-1332.

Cantley, G.B. (2004). "The Dividend Tax Cut: Bush's Prescription for Rescuing the Economy," *Akron Tax Journal*, vol.19, p.25-48.

Creedy, J. & Gemmell, N. (2004). “The Income Elasticity of Tax Revenue; Estimates for Income and Consumption Taxes in the United Kingdom,” *Fiscal Studies*, vol. 25(1), p.55-77.

Cunat, A. & Maffezzoli, M. (2007). “Can Comparative Advantage Explain the Growth of US Trade,” *The Economic Journal*, vol. 117, p.583-602.

Engel, R. and Granger, C. (1987). “Cointegration and Error Correction: Representation, Estimations and Testing,” *Econometrica*, 55(2), p.251-276.

Erbil, C. (2001). “Trade Taxes Are Better?!?” *Brandeis University Boston College and EcoMod*. Available online at: ecomod.net/conferences/ecomod2001

Farmer, R.A. & Lahiri, A. (2006). “Economic Growth in an Interdependent World Economy,” *The Economic Journal*, 116, p.969-990.

Ghosh, S & Mourmouras, A. (2001). “Endogenous Inflationary Finance and Long-run Growth,” *Journal of Economic Studies*, vol. 28(1), p.43-54.

Granger, C. (1969). “Investigating Causal Relations by Econometric Models and Cross-spectoral Methods,” *Econometrica*, 37, p.424-438.

Glomm, G & Ravikumar, B. (1997). “Productive Government Expenditures and Long-run Growth,” *Journal of Economic Dynamics and Control*, 2, p.183-204.

Hsing, Y. (2005). “Economic Growth and Income Inequality: The Case of the US,” *International Journal of Social Economics*, vol.32, p.639-647.

Koch, S., Schoeman, N. and Tonder, V. (2005). “Economic Growth and the Structure of Taxes in South Africa: 1960-2002,” *South African Journal of Economics,* vol. 72:3, p.190-210.

Kollias, C & Paleologou, S. (2006). ”Fiscal policy in the European Union: Tax and Spend, Spend and Tax, Fiscal Synchronization or Institutional Separation?,” *Journal of Economic Studies*, vol.33(2), p.108-120.

Kwiatkowski et, D., Phillips, P. Schmidt, P. and Shin, Y. (1992). “Testing the Null Hypothesis of Stationary Against the Alternative of a Unit Root. How Sure are We that Economic Time Series Have a Unit Root?,” *Journal of Econometrics*, 54, p.159-178.

Loganathan, N. & Taha, Roshaiza. (2007). “Have Taxes Led Government Expenditure in Malaysia?,” *Journal of International Management Studies*, vol.2(2), p.99-113.

Malaysia. (2004). Economic Report 2004/2005 Available online at http://www.treasury.gov.my

Malaysia. (2005). Economic Report 2005/2006. Available online at http://www.treasury.gov.my

Malaysia. (2006). Economic Report 2006/2007 Available online at http://www.treasury.gov.my

Malaysia. (2006). *Inland Revenue Board: Annual Report 2000*. Available online at http://www.treasury.gov.my

Malaysia. (2006). *Ninth Malaysia Plan 2006-2010.* Percetakan Nasional Berhad: Kuala Lumpur.

Malaysia. (2006). *First Malaysia Plan 1966-1970.* Percetakan Nasional Berhad: Kuala Lumpur.

Malaysia. (2007). *Federal Government Financial Position 2007*. Available online at http://www.treasury.com

Malaysia. (2007). *Federal Government Finance 2007*. Available online at http://www.epu.jpm.my

Malaysia. (2007). Tax Analysis Division, 2007, Will the Implementation of the "Current Year Assessment" System Affect the Government's Cash Flow? Available online at http://www.treasury.com

Malaysia. (2006). The 2006 Budget Speech. Available online at http://www.treasury.com

Padovano, F. & Galli, E. (2002) "Comparing the Growth Effects of Marginal vs. Average Tax Rates and Progressivity," *European Journal of Political Economy*, vol.18, p.529-544

Sinha.D. (1998) "Government Expenditure and Economic Growth in Malaysia," *Journal of Economic Development*, vol.23, p.71-80.

Yamarik, S. (2000) "Can Tax Policy Help Explain State-Level Macroeconomic Growth?," *Economics Letters*, 68, p.211-215.

Wang, D. (2007) "Convergence Tests on Tax Burden and Economic Growth Among China, Taiwan and the OECD Countries," *Physica*, A380, p.278-286.

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THE DETERMINANTS OF CAPITAL STRUCTURE OF LARGE NON-FINANCIAL LISTED FIRMS IN NIGERIA

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ABSTRACT

This paper examines capital structure determinants of non-financial firms in Nigeria using a panel of 33 large firms. Statistical tests are performed for the period 1990-2004. The results reveal that profitability, tangibility and company size are positively related to total debt and long-term debt, and growth opportunities are negatively associated with total debt. The empirical results indicate that the financing decisions of large firms in Nigeria can be explained by the determinants suggested by trade-off theory.

JEL: G31, G32

INTRODUCTION

The move towards a free market, coupled with the widening and deepening of various financial markets has provided the basis for the corporate sectors to optimally determine their capital structure. This environment has also encouraged more meaningful research of the capital structure issue. The corporate sector in Nigeria is characterized by a large number of firms operating in a largely deregulated and increasingly competitive environment. Since 1987, financial liberalization has changed the operating environment of firms, by giving more flexibility to the Nigerian financial managers in choosing the firm's capital structure.

There are only a limited number of studies that examine factors which influence the capital structure of Nigerian firms. Although the capital structure issue has received substantial attention in developed countries, it has remained neglected in the developing countries. The reasons for this neglect are discussed by Bhaduri (2002). He notes that until recently, development economics have placed little importance to the role of firms in economic development. Second, until the eighties, the corporate sectors in many lesser developed countries (LDCs) faced several constraints on their choices regarding sources of funds. Access to equity markets was either regulated, or limited due to the underdeveloped stock market (Bhaduri, 2002).

Planning capital structure involves the consideration of shareholders interest and other groups. Upon firm initiation, a company should plan its capital structure. Subsequently, whenever funds have to be raised to finance investment, a capital structure decision is also involved (Salawu, 2007). It is clear that capital structure is an important management decision as it greatly influences the owner's equity return, the owners risks as well as the market value of the shares. It is therefore incumbent on management of a company to develop an appropriate capital structure. In doing this, all factors that are relevant to the company's capital decision should be properly analyzed and balanced. The remainder of the paper is organized as follows: Section II discussed the relevant literature. Section III describes the methodology and data used. Section IV presents the results, and some concluding comments are provided in Section V.

LITERATURE REVIEW

The empirical literature suggests a number of factors that may influence the financial structure of companies. Salawu, (2006) examined the considerable factors involved in deciding on the appropriate amount of equity and debt in the Nigerian banking industry and the factors influencing banks' capital structure. His study revealed that ownership structure and management control, growth and opportunity, profitability, issuing cost, and tax issues associated with debt are the major factors influencing bank's capital structure. Following Rajan and Zingales (1995); Banerages, Heshmati and Wihlborg (2000) and Bevan and Danbolt (2001), the following variables shall be considered in this study: company size, profitability, tangibility, growth opportunities, non-debt tax shields and dividend as possible determinants of the capital structure choice. We discuss the relevant literature for each of these variables in turn.

Company Size

The trade-off theory predicts an inverse relationship between size and the probability of bankruptcy, i.e., a positive relationship between size and leverage. Berger et. al., (1997) find the positive relationship between leverage and company size. The results hold regardless of whether the regressions are estimated using OLS, random effects or fixed effects panel estimation. Rajan and Zingales (1995) argued that larger firms tend to be more diversified and fail less often, so size may be an inverse proxy for the probability of bankruptcy. Large firms are also expected to incur lower costs in issuing debt or equity. Thus, large firms are expected to hold more debt in their capital structure than small firms.

Ozkan (2000, 2001) – who control for firm heterogeneity through random effect and generalized method of moments (GMM) estimation respectively – obtain results similar to prior studies, which have failed to control for such effects. Barclay and Smith (1996), Stohs and Mauer (1996), Demirgue Kunt and Maksimovic (1999) all find debt maturity to be positively correlated with company size.

According to Drobetz and Fix (2003), the effect of size on leverage is ambiguous. Some studies reveal a positive relationship between size and the debt maturity structure of companies (Michaelas et. al. 1999). Accordingly, the pecking order theory of the capital structure predicts a negative relationship between leverage and size, with larger firms exhibiting increasing preference for equity relative to debt. Despite some contradictory evidence, the weight of available empirical evidence finds debt maturity to be positively correlated with company size.

Profitability

In the trade-off theory, agency costs, taxes, and bankruptcy costs push more profitable firms toward higher book leverage. First expected bankruptcy costs decline when profitability increases. Second, the deductibility of corporate interest payments induces more profitable firms to finance with debt. In a trade-off theory framework, when firms are profitable, they prefer debt to benefit from the tax shield. In addition, if past profitability is a good proxy for future profitability, profitable firms can borrow more, as the likelihood of paying back the loans is greater. In the agency models of Jensen and Meckhing (1976), Easterbook (1984), and Jesen (1986), higher leverage helps control agency problems by forcing managers to pay out more of the firm's excess cash. Accordingly, the pecking-order model predicts a negative relationship between book leverage and profitability.

Again, the empirical evidence on the issue is mixed. For instance, Toy et. al., (1974); Kester (1986); Titman and Wessels (1988); Harris and Raviv (1991); Bennett and Donnelly (1993); Rajan and Zingales (1995), and Michaeles et. al. (1999); Booth et al. (2001); Bevan and Danbolt (2001) all find leverage to be negatively related to the level of profitability (supporting the pecking-order theory), while Jensen, Solberg and Zorn (1992) find a positive one (supporting the trade-off theory).

Tangibility

The nature of a firms assets impact capital structure choice. Tangible assets are less subject to informational asymmetries and usually they have a greater value than intangible assets in the event of bankruptcy. In addition, moral hazard risks are reduced when the firm offers tangible assets as collateral, because this constitutes a positive signal to the creditors. Creditors can sell off these assets in the event of default.

The trade-off theory predicts a positive relationship between measures of leverage and the proportion of tangible assets. Relative to this theory, Bradley et. al., (1984); Rajan and Zingales (1995); Kremp et al., (1999) and Frank and Goyal (2002) find leverage to be positively related to the level of tangibility. However, Chittenden et. al., (1996) and Bevan and Danbolt (2001) find the relationship between tangibility and leverage to depend on the measure of debt applied.

Alternatively, Grossman and Hart (1982) argue that the agency costs of managers consuming more than the optimal level of perquisites is higher for firms with lower levels of assets that can be used as collateral. The monitoring costs of the agency relationship are higher for firms with less collateralizable assets. Therefore, firms with less collateralizable assets might voluntarily choose higher debt levels to limit consumption of perquisites (Drobetz and Fix, 2003). This agency model predicts a negative relationship between tangibility of assets and leverage. Firms with more tangible assets have a greater ability to secure debt. Consequently, collateral value is found to be a major determinant of the level of debt financing (Omet and Mashharance, 2002).

From a pecking order theory perspective, firms with few tangible assets are more sensitive to informational asymmetries. These firms will thus issue debt rather than equity when they need external financing (Harris and Raviv, 1991), leading to an expected negative relation between the importance of intangible assets and leverage.

Growth Opportunities

The trade-off model predicts that firms with more investment opportunities have less leverage because they have stronger incentives to avoid under-investment and asset substitution that can arise from stockholder-bondholder agency conflicts. This theory predicts a negative relationship between leverage and investment opportunities.

The empirical evidence regarding the relationship between leverage and growth opportunities is, at best, mixed. Titman and Wessles (1988); Barclay and Smith (1996) and Chen et. al., (1997) all find a negative relationship between growth opportunities and the level of either long-term or total debt. Rajan and Zingales (1995) find a negative relationship between growth opportunities and leverage. They suggest that this may be due to firms issuing equity when stock prices are high. As mentioned by Hovakimian et al. (2001), large stock price increases are usually associated with improved growth opportunities, leading to a lower debt ratio. However, Bevan and Danbolt (2001) find a negative correlation between growth and long-term debt, but find total leverage to be positively related to the level of growth opportunities. On the other hand, Beran and Danbolt (2001) find short-term debt to be positively related to growth opportunities. In fact, the simple version of the pecking order theory supports a positive relationship. Debt typically grows when investment exceeds retained earnings and falls when investment is less than retained earnings.

Non-debt Tax Shields

The effective tax rate has been used as a possible determinant of the capital structure choice. According to Modigliani and Miller (1958), if interest payments on debt are tax-deductible, firms with positive taxable income have an incentive to issue more debt. That is, the main incentive for borrowing is to take advantage of interest tax shields.

Accordingly, in the framework of the trade-off theory, one hypothesizes a negative relationship between leverage and non-debt tax shields. DeAngelo and Masulis (1980) argue that the marginal corporate savings from an additional unit of debt decreases with increasing non-debt tax shields. This is because of the likehood of bankruptcy increases with leverage. The empirical evidence is mixed. According to Graham (2000), the tax shield accounts on average to 4.3% of the firm value when both corporate and personal taxes are considered. In this study, dividend is included as a supplementary indicator of firm liquidity.

METHODOLOGY AND DATA

This study covers only non-financial quoted companies on the first and second tiers of Nigerian Stock Exchange. Thirty-three firms with market capitalization of five hundred million naira and above were regarded as large firms and included in the sample. Data were obtained from the annual reports of the sampled firms and publications of the Nigerian Stock Exchange. The study excludes the financial and securities sector companies for two reasons. First, these firms tend to have substantially different financial characteristics and use of leverage than other companies. In addition, the balance sheets of the firms in the financial sector (banks, insurance companies, and investments trust) have a significantly different structure from those of non-financial firms.

The selection of the variables (regressand and regressor) is primarily guided by the results of the previous empirical studies and the available data. The dependent and independent variables are defined so that they are consistent with those of Rajan and Zingales (1995). The analysis utilizes the following variables.

Leverage (LEV1)	= Total Debt/Total Assets
Leverage (LEV2)	= Long-term debt/Total Assets
Leverage (LEV3)	= Short-term debt/Total assets
Profitability	= Earnings after interest and tax to book value of total assets
Tangibility	= Book value of fixed assets to total assets
Size	= Natural logarithm of sales
Growth Opportunities	= Total Assets in year (t)/Total assets in year (t-1)
Non-debt Tax Shields (NDTS)	= Depreciation divided by Total Assets
Dividend (DIV)	= Dividend paid/Book value of equity

Using the above defined variables, the following model is estimated for the sample

$$\text{Leverage}_{it} = \beta_1 + \beta_2\text{Profitability} + \beta_3\text{Tangibility} + \beta_4\text{Growth} + \beta_5\text{Size} + \beta_6\text{NDTS} + \beta_7\text{Dividend} + \mu_i + \varepsilon_{i,\,t} \quad (1)$$

Where μ is used to capture the unobserved individual effects (either fixed or random), and ε is the error term, which represents measurement errors in the independent variables, and any other explanatory variables that have been omitted, as well as in the measurement of the independent variables. In order to estimate the panel regression model, three alternative methods were used: pooled ordinary least squares, the fixed effects model, and random effects model.

RESULTS AND DISCUSSION

Based on data availability, six potential determinants of capital structure are analyzed in this study-profitability, tangibility, growth opportunity, size, non-debt tax shields and dividend. The regression results for the large firms are presented in Tables 1 to 3 respectively.

In the case of large firms, the pooled OLS, fixed effect and random effects results in Tables 1 to 3 reveal a positive correlation between profitability and total and long-term debt. These results indicate that large firms in Nigeria are profitable and they are expected to prefer debt in order to benefit from the tax shield. The positive effect might be due to the tax advantage of debt with profitable firms having a high demand. Also, debt holders may see more profitable firms as less risky (i.e. probability of bankruptcy is low). As a result, these firms can get debt financing relatively easily. However, despite the fact that the relationship between the profitability and leverage (LEV1 and LEV2) are positive, they are not statistically significant. The relationship between profitability and short-term debt (LEV3) are negative under pooled OLS and fixed effect estimation with coefficients of –0.0013 and –0.0031 respectively. This suggests that large firms in Nigeria prefer short-term debt to long-term debt. The positive coefficients of profitability for large firms provide evidence supporting the trade-off theory.

In the case of large firms, tangibility (TANG) is positively correlated with both total debt and long-term debt. In Table 1, the coefficient of tangibility is positive (0.0743, 0.0685 and 0.0976 respectively) and significant at the 1%, 5% and 1% levels respectively in the case of long-term liabilities/total asset (LEV2). Firms with more tangible assets have a greater ability to secure debt. Consequently, collateral value is found to be a major determinant of the level of debt financing. This finding shows that a large firm in Nigeria has the potential to obtain external financing, especially equity and short-term debt.

For the large firms in Nigeria, tangible assets are likely to have an impact on the borrowing decisions of a firm because they are less subject to informational asymmetries and usually have a greater value than intangible assets in the event of bankruptcy. The trade-off theory predicts a positive relationship between measures of leverage and the proportion of tangible assets. Most empirical studies conclude to a positive relation between tangibility and the level of debt (Rajan and Zingales, 1995, Kremp et al, 1999; Frank and Goyal, 2002).

The pooled OLS and random results in Table 1 uncover a negative correlation between growth opportunities (GROW) and total liabilities for large firms. Similarly, the results in Table 2 reveal negative correlation between growth and long-term liabilities. The results are significant at the 5% level using the pooled OLS and random effect methodologies. This result is consistent with trade-off theory. The theory predicts that firms with more investment opportunities have less leverage because they have stronger incentives to avoid under-investment and asset substitution that can arise from stockholder-bond holder agency conflicts. Moreover, the costs associated with agency problems are likely to be higher for growing firms, since they may have more flexibility in the choice of future investments. Therefore, one would expect a negative association between long-term debt and growth of firm. Table 3 shows that large firms in Nigeria substitute short-term debt for long-term debt. This is because growth is positively correlated with short-term debts (LEV3) especially under pooled OLS and fixed effect estimation. However, the growth factor coefficients are not significant in the short-term model.

The size (SIZ) of firms (measured by the logarithm of sales) is positive related with all debt types except pooled OLS result for long-term debt (LEV2), which is negative but significant at 5% level. The signs of the coefficients of the firm size are consistent with trade-off theory. These coefficients are significant for total debt in Table 1 (both OLS and fixed) and short-term debt in Table 3 (under fixed effect at 5% level).

These finding suggest that large firms have the capacity to employ more debt, because they can hold a greater bargaining power towards creditors. In other words, larger firms might be more diversified and fail less often. To the extent that this is the case, small firms are expected to borrow less than large firms. Moreover, the informational asymmetries tend to be less severe for larger firms than for smaller firms and hence, large firms find it easier to raise debt financing. Therefore, a significant and positive debt level indicates that large firms in Nigeria depend more on short-term borrowing. This is probably because firms with poor financial health solve the risk premium problem by issuing short-term borrowing as it involves less risk for creditors.

In the case of large firms, non-debt tax shield (NDTS) is positively related to both total debt and short-term debt (LEV3) in Tables 1 and 3. However, NDTS is negatively correlated with long-term debts in Table 2. The inverse relationship between NDTS and long-term debt suggests that tax deductions for depreciation, losses and investment tax credits are substitutes for the tax benefits of debt financing. Therefore a firm with a large non-debt tax shield is likely to be less leveraged.

Table 1: Regression Model Estimates: Total Liabilities (LEV1)

	OLS	**Fixed Effect Result**	**Random Effect Result**
Constant	0.3856	0.3057	-0.0518
	(3.4893)	(2.4703)	(-0.2308)
PROF	0.0133	0.0006	0.0043
	(1.0999)	(0.0519)	(0.2604)
TANG	0.0650	0.3240	0.8917
	(1.0379)	(3.7794)	(7.8658)
GROW	-0.0007	0.0069	-0.0142
	(-0.0277)	(0.3481)	(-0.5813)
SIZ	0.0152	0.0171	0.0243
	(2.1433)**	(2.2011)**	(1.6438)
NDTS	0.5927	0.6737	3.6525
	(1.1999)	(1.3776)	(4.7389)*
DIV	0.0116	-0.0005	-0.0020
	(3.7352)*	(-0.1472)	(-0.2523)
Adjusted R^2	0.1013	0.5241	0.2621
F – statistic	9.720	14.44	28.47
	(0.0000)	(0.0000)	(0.0000)
D-Watson Stat	1.075	1.346	
Hausman Test	-	-	33.34
			(0.0000)
Cross-sections included	33	33	33
Number of observations	465	465	465

*Profitability (PROF) refers to earning after interest and tax/ net assets, tangibility (TANG) is defined as fixed assets/total assets, growth prospect (GROW) refers to the ratio of total assets in year t to total assets in year t-1. Size (SIZ) is the natural logarithm of sales. Non-debt tax shield (NDTS) is defined as the ratio of depreciation to total assets and dividends (DIV) refers to dividend paid /total equity. Numbers in parentheses appearing below the coefficients are t-values. *, ** and *** indicates the coefficient is significant at the 1, 5 and 10 percent levels respectively.*

The large firms results in Table 2 indicate that dividend (DIV) is positively correlated with long-term debt under each of the three estimation models. However, only the fixed effect result is statistically significant at the 5% confidence level. Both total debt and short-term debt are negatively correlated with dividend except in the pooled OLS results. In the case of short-term debt, none of the coefficients are significant. These results indicate that dividend payment does not represent a better financial approach for large firms in Nigeria.

CONCLUSION

The results of this work further confirm some prior findings and extend the capital structure analysis by analyzing capital structure in Nigerian firms using additional firm characteristics such as non-debt tax shields, dividend and a decomposition analysis of firm leverage. The findings revealed that profitability has positive impact on leverage of large firms in Nigeria, confirming that the tax advantage of debt financing has relevance in these firms. The results indicate that large Nigerian firms are profitable and

Table 2: Regression Model Estimates: Long Term Liabilities (LEV2)

	OLS	Fixed Effect Result	Random Effect Result
Constant	0.1909 (6.0894)	0.0489 (1.6180)	0.0624 (1.2166)
PROF	0.0039 (1.5367)	0.0020 (0.7729)	0.0021 (0.5336)
TANG	0.0743 (4.2927)*	0.0685 (2.9396)**	0.0976 (3.7773)*
GROW	-0.0122 (-2.0078)**	-0.0084 (-1.7982)	-0.0129 (-2.2576)**
SIZ	-0.0095 (-4.5597)*	0.0007 (0.3648)	0.0005 (0.1335)
NDTS	-0.0144 (-0.1110)	-0.0790 (-0.6808)	-0.4013 (-2.2455)**
DIV	0.0018 (1.7937)	0.0029 (3.1037)**	0.0010 (0.5676)
Adjusted R^2	0.0785	0.4469	0.0339
F – statistic	7.591 (0.0000)	10.867 (0.0000)	3.712 (0.0013)
D-Watson Stat	0.7167	0.9656	0.897
Hausman Test	-	-	13.774 (0.0323)
Cross-section included	33	33	33
Number of observations	465	465	465

*Profitability (PROF) refers to earning after interest and tax/ net assets, tangibility (TANG) is defined as fixed assets/total assets, growth prospect (GROW) refers to the ratio of total assets in year t to total assets in year t-1. Size (SIZ) is the natural logarithm of sales. Non-debt tax shield (NDTS) is defined as the ratio of depreciation to total assets and dividends (DIV) refers to dividend paid /total equity. Numbers in parentheses appearing below the coefficients are t-values. *, ** and *** indicates the coefficient is significant at the 1, 5 and 10 percent levels respectively.*

they are expected to prefer debt in order to benefit from the tax shield. However, the results reveal that large firms in Nigeria prefer short-term debt to long-term debt financing. The study shows that there was a significant positive relationship between asset structure (tangibility) and long-term debt ratios. Therefore, collateral value is found to be a major determinant of the level of debt finance. The size of the company was found to have a statistically significant positive relationship with both total debt and short-term debt ratios for the sample.

The results reveal that dividend payment does not represent a better financial approach for large firms in Nigeria. In addition, non-debt tax shields are positively and significantly correlated with capital structure. This suggests that large Nigerian firms that have large non-debt tax shields are less leveraged. The evidence of the behavior of large firms in Nigeria is consistent with the trade-off theory.

In conclusion, management should strive to identify and maintain an optimal capital structure of the firm since it represents the point where the market value of the firm is maximized. Furthermore, the top echelon of company management should take interest in the issue of capital structure and constantly monitor its form and adaptability. Further study of this issue might involve taking a dynamic look at the issue and formulating dynamic models of debt policy with instrumental variables. Such an approach could enrich the analysis here. Dynamic models enable researchers to discriminate between the various factors that impact the capital structure and those that impact on the speed of adjustments.

Table 3: Regression Model Estimates: Short Term Liabilities (LEV 3)

	OLS	Fixed Effect	Random Effect
Constant	0.4179 (3.6602)	0.3046 (2.5175)	-0.0436 (-0.1975)
PROF	-0.0013 (-0.0988)	-0.0031 (-0.2459)	0.0011 (0.0691)
TANG	-0.0665 (-1.0551)	0.0762 (0.8978)	0.7718 (6.9140)*
GROW	0.0228 (0.9269)	0.0134 (0.6884)	-0.0007 (-0.0306)
SIZ	0.0068 (0.9333)	0.01616 (2.1516)**	0.0185 (1.2788)
NDTS	0.9941 (1.8822)	0.9382 (1.9774)	4.2496 (5.5897)*
DIV	0.0163 (4.8639)	-0.0017 (-0.4928)	-0.0005 (-0.0764)
Adjusted R^2	0.0853	0.5855	0.2498
F – statistic	8.211 (0.0000)	18.250 (0.0000)	26.752 (0.0000)
D-Watson Stat	0.733	1.281	0.949
Hausman Test	-	-	31.898 (0.0000)
Cross-section included	33	33	33
Number of observations	465	465	465

*Profitability (PROF) refers to earning after interest and tax/ net assets, tangibility (TANG) is defined as fixed assets/total assets, growth prospect (GROW) refers to the ratio of total assets in year t to total assets in year t-1. Size (SIZ) is the natural logarithm of sales. Non-debt tax shield (NDTS) is defined as the ratio of depreciation to total assets and dividends (DIV) refers to dividend paid /total equity. Numbers in parentheses appearing below the coefficients are t-values. *, ** and *** indicates the coefficient is significant at the 1, 5 and 10 percent levels respectively.*

REFERENCES

Banerjee, S., Heshmati, A. and Wihlborg, C. (2000), "The Dynamics of Capital Structure". SSE/EFI *Working Paper Series in Economics and Finance* No. 333, May 12.

Barclay, M.J. and Smith, C.W. (1996), On Financial Architecture: Leverage, Maturity and Priority, *Journal of Applied Corporate Finance,* 8(4), p.4-17.

Bennett, M. and Donnelly, R. (1993), "The Determinants of Capital Structure: Some UK Evidence", *British Accounting Review*, 25(1), p. 43-59.

Berger, P. G., Ofek, E and Yermack, D.L. (1997), "Managerial Entrenchment and Capital Structure Decisions" *The Journal of Finance*, Vol.LII, No. 4, p.1411-38. September.

Bevan, A. A. and Daubolt, J. (2001), "Testing for Inconsistencies in the Estimation of UK Capital Structure Determinants", Working Paper 2001/4, Department of Accounting and Finance, University of Glasgow, Glasgow G 12 *LE.

Bhaduri, S.N. (2002), "Determinants of Capital Structure Choice: A Strudy of the Indian Corporate Sector". *Applied Financial Economics*, Vol. 12, p. 655-665.

Booth, L;, Aivazian, Demirguc – Junt, A. and Maksimolic, V. (2001), "Corporate Structures in Developing Countries" to be Published in the *Journal of Finance*, Vol. 56, p.87-130.

Bradley, M., Jarrell, G. and Kim, E. H. (1984), "On the Existence of an Optimal Capital Structure: Theory and Evidence," *Journal of Finance,* 39, p. 857 – 878.

Chen, C.J.P., Cheng, C.S.A., He, J. and Kim. J. (1997), "An Investigation of the Relationship Between International Activities and Capital Structure", *Journal of International Business Studies,* (Third Quarter), p. 563-577.

Chittenden, F., Hall, G. and Hutchinson, P. (1996), "Small Firm Growth, Access to Capital Markets and Financial Structure: Review of Issues and An Empirical Investigation". *Small Business Economics*, Vol. 8, p. 59-67.

De Angelo, H. and Masulis, R. (1980), "Optimal Capital Structure under Corporate and Personal Taxation", *Journal of Financial Economics*, 8(1), p.3-29.

Demirgue – Kunt, A. and Maksimovic, V. (1996), "Stock Market Development and Corporate Finance Decisions". *Finance and Development* 33 (2), p. 47 – 49.

Drobertz, W and Fix, R. (2003), "What are the Determinants of the Capital Structure? Some Evidence for Switzerland", WXYZ / Department of Finance, Working Paper No. 4103, University of Basel.

Easterbrook, F. (1984), "Two-Agency Cost Explanations of Dividends, *American Economic Review,* 74, p. 650-659.

Frank,M.Z. and Goyal, V.K. (2003), "Testing the pecking order theory of capital structure", *Journal of Financial Economics*, Vol.67, p. 217-248.

Graham, J.R. (2000), "How big are the Tax Benefits of Debt?" *Journal of Finance* 55 (No. 5, October), pp. 1901-1941.

Grossman, S. and Hart, O. (1982), "Corporate Financial Structure and Managerial Incentives", in the Economics of Information and Uncertainty, (ed.) J. McCall, Chicago: University of Chicago Press.

Harris, M. and Raviv, A. (1991), "The Theory of the Capital Structure", *Journal of Finance* 46, p. 297-355.

Jensen, M.C. and Meckling, W.H. (1976), "Theory of the Firm: Managerial Behaviour, Agency Costs and Ownership Structure", *Journal of Financial Economics*, 3, p. 305-360.

Jensen, M. (1986), "Agency Costs of Free Cash Flozo, Corporate Finance and Takeovers, *American Economic Review*, 76.

Jensen, M.; Solberg, D. and Zorn, T. (1992), "Simultaneous Determination of Insider Ownership, Debt and Dividend Policies", *Journal of Financial and Quantitative Analysis* 27, p. 247-261.

Kester, V. (1986), "Capital and Ownership Structure: A Comparison of United States and Japanese Manufacturing Corporations" *Financial Management*, 15, p.5-16.

Kremp, E., Stoss, E., and Gerrdesmeier, D. (1999), 'Estimation of a debt function: evidence from French and German firm panel data', in Sauve, A., Scheuer, M. (ed.) Corporate finance in Germany and France. A joint research project of Deutsche Bundesbank and the Banque de France, SSRN working paper.

Michaelas, N., Chittenden, F. and Poutziouris, P. (1999), "Financial Policy and Capital Structure Choice in UK. SMEs: Empirical Evidence from Company Panel Data, *Small Business Economics* 12, p. 113-130.

Modighiani, F. and Miller, M.H. (58), "The Cost of Capital, Corporation Finance and the Theory of Investment", *American Economic Review*, 48, p. 261-277.

Omet, G. and Mashharawe, F. (2002), "The Capital Structure Choice in Tax Contrasting Environments: Evidence from the Jordanian, Kuwanti, Omani and Sandi Corporate Sectors, The Economic Research Forum 10th Annual Conference, December (Marrakesh, Morocco).

Ozkan, A. (2000), "An Empirical Analysis of Corporate Debt Maturity Structure", *European Financial Management*, 6(1), p.197-212.

Ozkan, A. (2001), "Determinants of Capital Structure and Adjustment to Long Run Target: Evidence from UK Company Panel Data" *Journal of Business Finance and Accounting* 28, p.175-198.

Rajan, R. and Zingales, L. (1995), "What do We Know about Capital Structure? - Some Evidence from International Data", *Journal of Finance*, Vol.50, p. 1421-1460.

Salawu, R.O. (2006), "The Determinants of the Capital Structure of Financial Firms in Nigeria: The Financial Managers' Perspectives" Proceeding of the 2006 Global Conference on Business and Finance, organized by the Institute for Business and Finance Research between 31 May and 3 June, at Costa Rica. p. 74-77.

Salawu, R.O. (2007), "An Empirical Analysis of the Capital Structure of Selected Quoted Companies in Nigeria". An Unpublished M.Phil Thesis submitted to the Department of Management and Accounting, Obafemi Awolowo University, Ile-Ife, Nigeria.

Stohs, M. H and Mauer, D. C. (1996), "The Determinants of Corporate Debt Maturity Structure, *Journal of Business*, 69 (3), p. 279 – 312.

Timan, S. and Wessels, R. (1988), "The Determinants of Capital Structure choice", *Journal of Finance*, 43, p.1-19.

Toy, N, Stonehill, A. , Remmers, L. , Wright, R. and Beekhuisen, T. (1974), "A Comparative International Study of Growth, profitability and risk as determinants of corporate debt ratios in the manufacturing sector", *Journal of Financial and Quantitative Analysis*, 9, p. 875 – 886.

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REEXAMINING THE EXPIRATION DAY EFFECTS OF STOCK INDEX DERIVATIVES: EVIDENCE FROM TAIWAN

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ABSTRACT

This study examines whether the expiration of derivative contracts affects the underlying spot assets in Taiwan. The expiration effect refers to abnormal return, price reversal, abnormal return volatility, and abnormal volume in underlying spot stock markets as derivatives contracts expire. Due to the unique settlement procedure in the Taiwan Futures Exchange, this study also examines if the expiration effects occur on the settlement day which is the next business day after the expiration day. Our empirical results indicate that expiration day effects do exist in Taiwan. However, the more pronounced expiration day effect occurs on the settlement day due to the unique settlement mechanism in Taiwan. This paper also investigates the expiration effect of MSCI Taiwan Stock Index Futures traded on the Singapore Exchange, which also uses Taiwan's stock market as the underlying asset. The results indicate that as SGXTW expires, there are also expiration effects such as abnormal return, abnormal return volatility, and abnormal volume in the Taiwan spot market.

JEL: G14, G18, G19

INTRODUCTION

In the new century, the globalization and liberalization of financial markets have been the trend. Hot money runs all over the world and beats the financial market. Stock index derivatives are one of the most successful and important innovations in the last century financial markets. The markets for these products first emerged in the U.S. in the 1980s and rapidly spread to other financial markets of Europe and Pacific Rim. Trading volume of index derivatives has grown dramatically and even exceeds that of the underlying assets. This metric is commonly used to measure the success of a financial contract market. The primary reason for the popularity of stock index futures and option markets are that they provide a fast and inexpensive way to change exposure to a stock market.

Despite their success, index derivatives are often criticized because of their effects on the market for the underlying securities. The positive effects of these new securities are due to new risk sharing opportunity and more complete financial markets. On the negative side, derivative securities, whose payoffs are a function of some other assets prices, also offer new opportunities for price manipulation. Expiration day effect is one of these negative effects.

Expiration day effects may result from a combination of several factors, including the existence of arbitrage opportunities, the cash settlement feature, the stock market procedures for accommodating the unwinding of arbitrage positions in the stocks, and attempts to purposely manipulate prices. Arbitrage positions are often unwound at the expiration of the index derivative contract. If index derivatives expire at the close, the derivatives self-liquidate through cash settlement at the closing price level. The stock position, on the other hand, must be liquidated through trades in the marketplace. An arbitrageur who longs the underlying stocks and shorts the derivative contract must sell the underlying stocks at their closing prices. If many arbitrageurs liquidate positions at the same time and in the same direction, price effects may happen.

The price effects on expiration day depends in part on the stock market procedures for accommodating order imbalances that may result when arbitrage positions are unwound. If the underlying market for the stocks is not deep and if suppliers of liquidity are not quick enough to respond to selling or buying pressure, the price effects of larger arbitrage unwinding will be large. If unjustified price effects were known to occur, informed investors would attempt to buy under priced stocks and sell overpriced stocks, thereby limiting price effects to fall within the bounds of transaction costs. If market mechanisms are not well designed to offset sudden imbalances, however, the price effects may be substantial.

Expiration day price effects may also result from attempts to manipulate stock prices. Such attempts may occur directly in the way an arbitrage position is unwound or indirectly through arbitrage unwinding that benefit other positions. An arbitrageur might engage in indirect manipulation, not to benefit the arbitrage account but to benefit another account.

The "triple witching hour" quarterly expiration of stock index futures, stock index options, and individual stock options in the U.S. has received vast attention from both academics and regulators. The settlement procedure has been blamed for generating volatile price swings in the final hour of trade. In June 1987, the settlement price was changed from the closing price to the opening price on the third quarterly Friday for some of the existing index futures and options contracts, aiming to reduce the volatility. However, a variety of other solutions for limiting expiration day effects were proposed. One of them was to use the expiration day average rather than the closing price as the settlement price.

Stock index expiration effects have been studied in the past. Widely known is a series of studies by Stoll and Whaley (1986, 1987, and 1991) that examined expiration day effects of U.S. index derivatives. Across all contract expirations since the inception of index futures, they found that the effects were remarkably consistent: index stock trading volume was abnormally high and observed price movements were small and within the bounds of transaction costs. Karolyi (1996) examined Nikkei 225 futures contract expirations, and, like Stoll and Whaley, concluded that the expiration of the Nikkei 225 futures induced abnormal trading volume but economically insignificant price effects. Stoll and Whaley (1997) found similar results for Australian All Ordinaries Share Price Index futures and option expiration.

As the Futures Trading Law was published in March 1997, it led to the establishment of Taiwan Futures Exchange (TAIFEX) in September in the same year. TAIFEX opened for business and launched its first product: the Taiwan Stock Exchange Capitalization Weighted Stock Index (TAIEX) futures on July 21, 1998. Over the years, TAIFEX has earned significant attention and recognition. According to the statistics from the Futures Industry Association (FIA), TAIFEX's global ranking based on total trading volume rose to the 18th in 2005, from the 57th in 1998. In 2007, the trading volume of TAIEX Options (TXO) is ranked the 4th globally.

Taiwan's fast growing economy has also attracted international fund managers to invest in its vibrant stock market. SGX MSCI Taiwan Index Futures traded on the Singapore Exchange (SGX) is the first contract that traded Taiwan equity derivative in the world. Since its launch in 1997, trading volume and open interest in SGX MSCI Taiwan Index Futures have grown significantly.

It is uncommon to see that different contracts with the similar underlying can develop like these. Moreover, there are few studies that compare the relationship between different expiration days in the same month. Of interest is to know what kind of role the different features of the settlement system play in expiration day effects. This study examines the expiration day effect of derivative contracts (index futures, stock futures, index options, and stock options) on the abnormal return, price reversal, abnormal price volatility, and abnormal volume of the underlying in the Taiwan stock market. The Taiwan experience can shed new light on the existing literature and can also provide some evidence on emerging markets.

AN OVERVIEW OF THE TAIWAN DERIVATIVES MARKETS

From its establishment in July 1998 through December 2007, twelve financial derivatives products had been launched on the TAIFEX. The Taiwan Stock Exchange Capitalization Weighted Stock Index futures (TX) were the index product first introduced on July 21, 1998, followed by the Electronics Sector Index Futures (TE) and the Finance Sector Index Futures (TF) on July 21, 1999. In 2001, two other new products were launched: the Mini-TAIEX futures (MTX) and the TAIEX options (TXO) on April 9 and December 24. The former was designed to attract more small investors to participate, by asking a margin of one quarter of that for the TAIEX futures. The debut of the TAIFEX options added the first option products to the domestic futures market, which offered investors more vehicles to hedge. On June 30, 2003, the TAIFEX launched Taiwan 50 Futures (T5F) that consisted of the fifty biggest market value stocks in TSEC. Two more option products were launched on March 28, 2005: the Electronic Sector Index option (TEO) and the Finance Sector Index option (TFO), supporting more hedging tools to the investors. In order to attract more foreign investors, on March 27, 2006, TAIFEX introduced TAIFEX MSCI Taiwan Index Futures (MSF) and TAIFEX MSCI Taiwan Index Option (MSO). Their underlying index, the MSCI Taiwan Index^{SM}, is widely used by institutional investors worldwide to track the movement of Taiwan's stock market. The contacts are priced in U.S. dollar. On October 8, 2007, NonFinance NonElectronics Sub-index Futures (XIF) and NonFinance NonElectronics Sub-Index Options were launched to meet the requirement of investors who wished to hedge.

In the early years, the trading volumes of the index derivatives were small. Because of the diversified product line and growing investor familiarity, volume greatly increased, attracting the interest of arbitrageurs and speculators. As shown in Table 1, the trading volumes of each index derivative accelerate year by year. In 2004, the TAIFEX was awarded the honor "Derivates Exchange of the Year 2004" by Asia Risk. Moreover, in 2005, the TAIFEX expanded to become the 18th largest market in the world and the trading volume of TXO is the forth largest in single product in the world.

Table 1: Trading Volume Growth of Index Derivatives on the Taiwan Futures Exchange

Contracts	Year					
	2002	2003	2004	2005	2006	2007
TAIFEX Futures (TX)	4,132,040 (45%)	6,514,691 (58%)	8,861,278 (36%)	6,917,375 (-21%)	9,914,999 (43%)	11,813,150 (19%)
Electronics Sector Index Futures (TE)	834,920 (22%)	990,752 (19%)	1,568,391 (58%)	1,179,643 (-25%)	1,459,821 (24%)	1,004,603 (-31%)
Finance Sector Index Futures (TF)	366,790 (-5%)	1,126,895 (207%)	2,255,478 (100%)	909,621 (-60%)	786,477 (-14%)	909,383 (16%)
Mini-TAIFEX Futures (MTX)	1,044,058 (183%)	1,316,712 (26%)	1,943,269 (48%)	1,088,523 (-44%)	1,760,583 (62%)	2,964,042 (68%)
TAIFEX Options (TXO)	1,566,446	21,720,083 (1287%)	43,824,511 (102%)	80,096,506 (83%)	96,929,940 (21%)	92,585,637 (-4%)
Taiwan 50 Futures (T5F)		4,068	6,157 (51%)	9,483 (54%)	332 (-96%)	506 (52%)
Electronics Sector Index Options (TEO)				680,026	773,353 (14%)	1,066,141 (38%)
Finance Sector Index Options (TFO)				756,570	937,044 (24%)	1,203,084 (28%)
TAIFEX MSCI Taiwan Index Futures (MSF)					8,333	1,132 (-86%)
TAIFEX MSCI Taiwan Index Options (MSO)					867,597	1,634,117 (88%)
Non-Finance Non-Electronics Sub-index Futures (XIF)						37,197
Non-Finance Non-Electronics Sub-index Options (XIO)						186,161

This table lists the trading volumes of index derivatives on the TAIFEX from 2002 to 2007. Numbers in parentheses indicate the average growth rate if trading volume per year for the current year versus the previous year.

The Singapore Exchange (SGX) is the first integrated securities and derivatives exchange in Asia Pacific. It was established on December 1, 1999 as a result of the merger of two well-established and respected financial institutions: the Stock Exchange of Singapore (SES) and the Singapore International Monetary Exchange Limited (SIMEX). The merger signifies the shift from a member-only club that existed largely to serve the interests of brokers - to a commercial, customer-focused organization. SGX MSCI Taiwan Index Futures is the first exchange traded Taiwan equity derivative in the world. Since its launch in 1997, trading volume and open interest in SGX MSCI Taiwan Index Futures have grown significantly. SGX MSCI Taiwan Index Options contracts offer another dimension of Taiwan equity investment and trading.

There are six main differences between TAIFEX and SGX Taiwan index futures contracts. First, the contracts traded in TAIFEX are quoted in New Taiwan Dollar, while the SGXTW traded in SGX is quoted in US dollar. Second, the TAIFEX index futures benefits from home market advantages in terms of local knowledge, proximity to news sources, and lack of language barriers, while the SGXTW obtains the first mover-advantage. Third, the trading mechanisms are different. Regular trading hours in both markets are the same, 8:45 to 13:45, in the morning, but SGX employs an Electronic Trading System from 14:45 to 19:00 in the afternoon. Fourth, the futures trading in Taiwan is levied a 0.025% transaction tax, while no transaction tax is applied to the Singapore Exchange. Fifth, the expiration day of TAIFEX is the third Wednesday of the month, but the expiration day of SGXTW is the day before the last trading day of the month. Finally, TAIFEX and SGXTS utilize different settlement procedures. The settlement price of TAIFEX is computed based on the first fifteen-minute volume-weighted average price on the next business day of expiration day. The settlement price of SGXTW is the closing price on the expiration day. The features of the derivative product of TAIFEX are showed in Appendix 1.

Because the unique settlement mechanism of TAIFEX distinguishes it from those in the U.S. markets or in the Hong Kong market we expected the expiration day effects in TAIFEX to behave somewhat in between those of the U.S. markets and the Hong Kong market. On the other hand, one of the characteristics of the Taiwanese futures market is that all the twelve index derivatives use the same settlement procedure. Thus, trading activities resulting from arbitrage and speculation should be more obvious when more derivatives products entered the market. If arbitrageurs attempt to liquidate their positions, or speculators are eager to change the final settlement price after the last trading day, then expiration effects in relation to the cash market should became more significant.

As discussed earlier, in the early and mid-1980s in the U.S., researchers expressed concern about the so-called "triple witching hour", the last hour of trading on the third Friday of the quarterly month when stock index futures, stock index options, and equity options all expired simultaneously. The index derivatives traded on the TAIFEX expire once every third Wednesday of the delivery months, and settlement price for each contract is computed based on the first fifteen-minute volume-weighted average of each component stock's prices in the index on the final settlement day. Therefore, the Taiwanese market could be used as a special case to examine the impact of expiration effects on the stock market. We will also analyze whether the effects will become more significant as more contracts expire at the same time.

LITERATURE REVIEW

There is extensive empirical evidence on the expiration effect of U.S. derivatives markets. Stoll and Whaley (1987) studied eight expiration days of the Standard and Poor's Composite 500 (S&P 500) futures contracts in 1984 and 1985, and examined the S&P 500 and S&P 100 price changes in the last hour of expiration days as compared with those of non-expiration days. By measuring price reversal, they observed that the spot market volatility significantly increased on the futures expiration days and stated that the expiration-day price reversal appeared to be associated only with the expiration of index futures contracts, but the index options expiration did not cause the abnormal price movement. They inferred that the additional cost of liquidity frequently increasing on expiration days lead to the price effect, and pointed out the efforts by the regulators to find effective ways for handling unexpected order imbalances in the stock market were desirable.

Nevertheless, Stoll and Whaley (1991) arrived at a different conclusion in another study by investigating the volatility and price effects on quarterly and monthly expiration days in the two and one-half year period before and the two and one-half year period after June 1987. The Chicago Mercantile Exchange, New York Stock Exchange, and New York Futures Exchange had changed the settlement time of their derivative contracts from the closing price to the opening price on expiration day beginning with the June 1987 quarterly expiration in order to lessen concern about abnormal stock price movements in the triple witching hour. Thus, from June 1987 forward, the last trading day of the S&P 500 futures contracts was moved from Friday to the Thursday on the third Friday of the contract month, and the final settlement price was based on the Friday opening for the underlying index. However, the expiration mechanism for the S&P 100 was not changed. Stoll and Whaley (1991) found that the return volatility is smaller at the expiration days during the period after June 1987 than at expiration days during the period before 1987. They also stated that switching the settlement from close to open would not decrease volatility meaningfully, but would only shift the location of volatility. They further concluded that regulators have overreacted to the expiration effect.

By using daily data from October 1, 1981 to June 30, 1987, Aggarwal (1988) examined the impact of index futures on the volatility of the S&P 500 and the Dow Jones Industrial Average (DJIA) indices with the over-the-counter (OTC) Composite index as a control. He also concluded that the high levels of the intra-day stock market volatility may be a result of futures-related activity. However, the inter-day price volatility of the underlying spot market should be nothing related to the index futures trading.

Herbst and Maberly (1990, 1991) adopted two different methodologies to investigate whether the change of the expiration settlement procedures for the S&P 500 in June 1987 reduces spot expiration volatility. They observed the change in triple witching hour volatility and expected price reversals before and after the new settlement mechanism. Their results were consistent with the findings of Stoll and Whaley (1991). They found that the volatility on the expiration-days had been significantly reduced, but the volatility shifted with a significant increase in the first hour volatility.

On the contrary, Bessembinder and Seguin (1992) used daily data on the S&P 500 index from January 1978 to September 1989 to investigate the volatility effect on expiration day and found different results. Adopting a regression of daily S&P 500 return standard deviation with the independent variables of spot-trading volume and a futures-trading, and days to expiration dummy variables, they concluded that although the equity volatility was slightly higher on the expiration days, it is not related to the futures.

Hancock (1993) investigated the triple witching hour effect by using minute-by-minute S&P 500 index values for the period from April 30, 1987 to July 24, 1989. He observed that the expiration day effect appeared in the underlying spot market. However, the noticeable increases or decreases in the spot market volatility around the triple witching hour in the U.S. can be attributed to the responses of investors to new information, especially important economic news often declared on Fridays.

Besides the U.S. stock markets, Chamberlain et al. (1989) studied the Toronto Stock Exchange (TSE) 300 during the last half-hour of trading on expiration and non-expiration days in the Canadian market from November 1985 to May 1987. They inferred that the rate of return during the last half hour of trading is significantly more volatile on expiration days than that on non-expiration Fridays. Moreover, the expiration returns also tended to be reversed during the first half-hour of trading on the following Monday. With these results, Chamberlain et al. (1989) concluded that index arbitrage activities really exist in the

Canadian market. Illueca and Lafuente (2005) investigated the IBEX 35 in the Spanish market from January 17, 2000 to December 20, 2002. They detect a slight downward trend in the spot index, but the behavior of index reversals suggests that this pattern is not attributable to the maturity of the futures index derivative contract.

Gannon (1994) studied the Share Price Index (SPI) futures and All Ordinaries Share Price Index in the Australian market with 15-minute intra-day data. The observations are sampled for the three months to expiration for the March 1992 SPI futures contracts, and came to a different conclusion. They conclude that the expiration-day volatility effect is not at all significant on the spot market. Bacha and Vila (1994) studied the Nikkei 225 Stock Average with its futures contracts traded on Singapore International Monetary Exchange (SIMEX), Osaka Securities Exchange (OSE) and Chicago Mercantile Exchange (CME) for the period from September 3, 1986 to August 31, 1991. They also find that the futures expiration days on all three types of futures contracts cause no higher cash market volatility than ordinary. Chow, Tung, and Zhang (2003) observed the 5-minute interval of the Hang Seng Index from 1990 to 1999 to examine the impact of the expiration of HSI derivatives on the underlying spot market in Hong Kong. They indicated that expiration days in Hong Kong may be associated with a negative price effect and some return volatility on the underlying stock market but there is no evidence of abnormal trading volume on the expiration day or price reversal after expiration, so the existence of expiration day effects cannot be confirmed in the Hong Kong market. They inferred that the HSI derivative market is different from most other markets because the settlement price is computed by taking the average of 5-minute quotations of the HSI on the last trading day.

DATA AND METHODOLOGY

Data

Following the previous studies, we assess the expiration day effects of index derivatives by focusing on abnormal price movements and trading volumes. To examine whether the stock market produces abnormal changes when the index derivatives expire, we employ the comparison-period approach (CPA) to compare the average returns, stock price volatilities, and trading volumes on expiration days with those for non-expiration days, as did Masulis (1980). The CPA is used to compare the estimated parameters, such as the mean and the standard deviation of returns, on expiration days with those on non-expiration days.

The data for this study is from TSEC (Taiwan Stock Exchange) and Taiwan Economic Journal, a local data vendor. The sample period is drawn from January 1, 2002 to December 31, 2007, including 72 expiration days. The data includes both the Taiwan Stock Exchange Capitalization Weighted Stock Index and MSCI Taiwan Index^{SM} collected at 5-minute intervals respectively during the entire sample period. Besides the whole period, we also analyze the sub-period to gain additional insights related to this event. Specifically, we investigate the expiration day effects using the full sample and the following four sub-periods: (1) January 2002 to June 2003, (2) July 2003 to March 2005, (3) April 2005 to March 2006, and (4) April 2006 to December 2007.

It is customary to choose the trading day before and after the expiration day as comparison days. To avoid any potential weekly effect, we also examine 5 trading days before and 5 trading days after the expiration day for comparison. Furthermore, based on the empirical finding of Stoll and Whaley (1991) and Hancock (1993) that the changing of settlement system may affect the reaction time of expiration effects, we also set the first 15-minute period and the whole day of the settlement day as the comparison base for examining the impacts of index derivatives.

Abnormal Return Effect

If various arbitrageurs and speculators create price pressure on the expiration day for the index derivatives, the returns of the stock market will abnormally move. In order to detect whether the abnormal price movement of the stock market on the expiration day exists or not, we compare the average return for the expiration day with that of the non-expiration day. We use the logarithm of the return to calculate the stock index return. For the *t*th 5-minute interval on the expiration day or non-expiration day in the *i*th month, the return ($r_{i,t}$) is:

$$r_{i,t}=\ln(\frac{P_{i,t+1}}{P_{i,t}}) \tag{1}$$

where $P_{i,t}$ is the opening stock index for the *t*th 5-minute interval on the expiration day or non-expiration day of the *i*th month. The average returns on the expiration days and the non-expiration days during the overall sample period are defined as:

$$AR_e=\frac{1}{N_e}\sum_{i=1}^{N_e} r_{e,i}\ , \tag{2}$$

and

$$AR_c=\frac{1}{N_c}\sum_{i=1}^{N_c} r_{c,i} \tag{3}$$

respectively, where N_e and N_c signify the sample numbers for the expiration days and non-expiration days throughout the overall sample period. From the literature we know that financial returns are neither independent, nor distributed normally, so we use the Modified Levene's paired t-test to examine whether the average returns of the stock market for the expiration days are different from those for the non-expiration days. The null hypothesis is $H_0{:}AR_{e,i}\text{-}AR_{c,i}=0$, and the alternative hypothesis is $H_1{:}AR_{e,i}\text{-}AR_{c,i}\neq 0$. The testing statistic t_L^* is:

$$t_L^*=\frac{\bar{d}_e\text{-}\bar{d}_c}{S_p\sqrt{\frac{1}{T_e}+\frac{1}{T_c}}} \tag{4}$$

where $\bar{d}_e$ and $\bar{d}_c$ are the mean of d_{ei} and d_{ci}, respectively, and $T_e=\sum_{i=1}^{N_e} N_{e,i}$ and $T_c=\sum_{i=1}^{N_c} N_{c,i}$, respectively, are the numbers of expiration days and non-expiration days with standard deviation S_p as follows:

$$S_p=\sqrt{\frac{\sum(d_{ei}\text{-}\bar{d}_e)^2+\sum(d_{ci}\text{-}\bar{d}_c)^2}{T_e+T_c\text{-}2}}\ . \tag{5}$$

We use the last hour period of the expiration day and the whole expiration day of TAIFEX and SGXTW to investigate the abnormal returns. Furthermore, we also set the first 15-minute period of the settlement day and the whole settlement day of TAFEX as the second comparison base for comparing the average stock market returns. This approach allows us to examine the change of the reaction time of expiration effects.

Index Reversal Effect

The prices in the spot markets may deviate from equilibrium if the arbitrageurs liquidate their position when the index derivatives expire due to price pressure or the speculators attempt to change the prices in the spot markets in order to increase their profits.

If there is indeed such a price pressure on the expiration day, then stock prices should on average reverse in the opposite direction after the derivative contracts have expired. That is, if the stock index goes up on the expiration day, then it will go down after the expiration day and the price pressure will disappear. If the stock index goes down on the expiration day, then it will go up after the expiration day. The temporary price overreaction should be observed in the underlying spot market. In other words, examining the existence of the price reversal effect on the underlying spot market can be another measure of the existence of the expiration effect of index futures trading. The approach is used by Stoll and Whaley (1991).

Here we use three measures to test whether the reversal effect exists when the index derivatives in the TAIFEX and SGXTW expire. Similar to the approach adopted by Stoll and Whaley (1991) and Hancock (1993), we first define the last 60-minutes index return, R_t, and the return after expiration, R_{t+1}, as:

$$R_t = \frac{P_{close,t} - P_{close\text{-}60,t}}{P_{close\text{-}60,t}} \tag{6}$$

and

$$R_{t+1} = \frac{P_{open,t+1} - P_{close,t}}{P_{close,t}} \tag{7}$$

respectively, where $P_{close\text{-}60,t}$ is the index price 60 minutes before the market close on the expiration day, $P_{close,t}$ is the closing price on the expiration day, and $P_{opne,t+1}$ is the opening price of the next trading day after expiration. The first measure of the price reversal effect is defined as:

$$Rev = \begin{cases} R_{t+1} & \text{if} \quad R_t < 0 \\ -R_{t+1} & \text{if} \quad R_t \geq 0 \end{cases}. \tag{8}$$

A positive Rev indicates a price reversal effect because the price trend for the day after the expiration day moves in the opposite direction to that for the expiration day. If Rev is negative, it indicates that the stock market does not exhibit a price reversal effect because the price trend for the day after the expiration day moves in the same direction as that for the expiration day. The price reversal may be caused by aggregated price pressure of index derivatives or it may happen in the stock market randomly. If it occurs randomly there is a 50% probability that the price trend will either move continuously forward or in reverse. We use the binomial test to determine whether the price reversal in the stock market after the expiration day occurs randomly or is statistically significant.

We may be unable to capture the potential price reversal effect in the Taiwan stock market by using the above measure provided by Stoll and Whaley (1991, 1997). Because the TAIFEX uses the average trading price of the first 15 minutes of the settlement day (the day after the expiration day) as the final settlement price, speculators may attempt to influence this price to their benefit. Major price pressure may occur during this 15-minute period, so we use two more measures to capture the price reversal effect. We use the average return for the stock market during the first 15 minutes of the day after the expiration day (settlement day of TAIFEX) as a comparison base to compare the return during the last 60-minute of the expiration day to the opening price of the day after the expiration day. That is, we define two new measures for the price of returns as:

$$R_t = \frac{P_{open,t+1} - P_{close-60,t}}{P_{close-60,t}} \tag{9}$$

and

$$R_{t+1}' = \frac{P_{open+15,t+1} - P_{open,t+1}}{P_{open,t+1}} \tag{10}$$

where $P_{open,t+1}$ is the opening index on the day after expiration day (settlement day of TAIFEX), $P_{close-60,t}$ is the index price 60 minutes before the market close on the expiration day, and $P_{open+15,t+1}$ is the index for the first 15-minute period after the opening of the day after the expiration day (the settlement day of TAIFEX). We define the second measure of the price reversal measure Rev' as:

$$Rev' = \begin{cases} R_{t+1}' & \text{if} \quad R_t < 0 \\ -R_{t+1}' & \text{if} \quad R_t \geq 0 \end{cases}. \tag{11}$$

Similarly, if Rev' is positive, then it indicates that the stock market exhibits a price reversal effect, because the price trends before and after the opening index price of the day after expiration day (settlement day of TAIFEX) moves in opposite directions. If Rev' is negative, then it indicates that the stock market does not exhibit a price reversal effect because the price trends before and after the opening index price of the day after expiration day move in the same direction. Because price reversal in the stock market after the opening index price of the day after expiration day is decided may take place randomly or may be caused by price pressure as the index derivatives expire, we still use the binomial test to determine whether the price reversal taking place in the stock market after the opening index price of the day after expiration day occurs randomly or whether it is statistically significant.

Moreover, we use the spot index right after the 15-minute period as a comparison base to compare the average return for the stock market during the first 15 minutes of the settlement day of TAIFEX. That is, we define two more measures for the price of returns as:

$$R_{t+1}'' = \frac{P_{open+15,t+1} - P_{open,t+1}}{P_{open,t+1}}, \tag{12}$$

and

$$R_{t+1}''' = \frac{P_{close,t+1} - P_{open+15,t+1}}{P_{open+15,t+1}}, \tag{13}$$

where $P_{open+15,t+1}$ is the index for the first 15-minute period after the opening of the settlement day of TAIFEX (the day after the expiration day), $P_{open+15,t+1}$ is the opening index on the settlement day of TAIFEX (the day after the expiration day), and $P_{close,t+1}$ is the closing index of the settlement day. We then define the third measure of the price reversal measure Rev" as:

$$\text{Rev}''=\begin{cases} R_{t+1}''' & \text{if} \quad R_{t+1}''<0 \\ -R_{t+1}''' & \text{if} \quad R_{t+1}''\geq 0 \end{cases}. \tag{14}$$

Similarly, if Rev" is positive, then it indicates that the stock market shows a price reversal effect, because the price trends before and after the final settlement price of TAIFEX is decided move in opposite directions. If Rev" is negative, then it indicates that the stock market does not exhibit a price reversal effect because the price trends before and after the settlement price of TAIFEX is decided move in the same direction. Because price reversal in the stock market after the settlement price is decided may take place randomly or may be caused by price pressure as the index derivatives expire, we still use the binomial test to determine whether the price reversal taking place in the stock market after the settlement price is decided occurs randomly or whether it is statistically significant.

Abnormal Price Volatility

If either arbitrageurs liquidate their positions or speculators attempt to influence the spot market price when index derivatives expire, which may cause the spot markets out of equilibrium. If such temporal price disequilibrium exists, then the index returns around the expiration should on average be more volatile. In general, trading volume and volatility within the stock market are higher during the opening and closing of the market.

Therefore, we test the volatility effect during the whole expiration day and the last hour before closing on the expiration day. We used the standard deviation of the return for the stock index to estimate the volatility of the stock market. A pooled F-test is used to test whether expiration day returns are associated with a higher volatility from that of the comparison period.

The null hypothesis is $H_0{:}S_e \leq S_c$, where S_e and S_c are the standard deviations of the returns of the stock index on the expiration day and the non-expiration day, respectively, and the F-statistic is:

$$F=\frac{S_e^2}{S_c^2}. \tag{15}$$

However, because the reaction time of expiration effects might move to the settlement day of TAIFEX, we also set the first 15-minute period of the settlement day of TAIFEX and the whole settlement day of TAIFEX as the second comparison base for examining the abnormal volatility effects.

Abnormal Volume Effect

If large stock positions from index arbitrage carry into the derivatives contract expiration day and are unwound with a large imbalance orders, then the volume of trading in index portfolio stocks should be abnormally high on expiration days. We used two measures to test the abnormal volume effect. The first measure compares the average trading volume of the last one hour on expiration days (denoted as V_e) with that of the daily average in the current month (denoted as V_c).

As in the case of testing for index reversals, a simple binomial test is used to test whether expiration days are associated with abnormally heavy trading volume. We denote the probability of $V_e > V_c$ as p, and the probability of $V_e \leq V_c$ as q = 1 -p. Under the null hypothesis that there are no abnormal trading volumes on expiration days, we expect an equal probability of observing either $V_e > V_c$ or $V_e \leq V_c$, i.e., p = q = 0.5.

The second measure to capture the abnormal volume effect, involves using the first 15 minutes of the day after the expiration day (the settlement day of TAIFEX) as the basis for comparison. Since the TAIFEX uses the average trading price of the first 15 minutes of the settlement day (the day after the expiration day) as the final settlement price, the price pressure may cause the trading volume to inverse suddenly. After this 15-minute period when the price pressure disappears, the trading volume will back to the normal level. Therefore, the major abnormal trading volume effect may occur only during this 15-minute period. This measure compares the stock market volume during the first 15-minutes period of the day after the expiration day (the settlement day of TAIFEX) with that on the non-settlement day, so as to capture the possible abnormal volume effect.

EMPIRICAL RESULTS AND ANALYSIS

In this section we present the results of our empirical investigation. The results of abnormal return effects are discussed first, followed by the results for the price reversal, abnormal return volatility, and trading volume tests on the expiration days and settlement days.

Abnormal Return Effect

First, we compare the average return of TAIFEX index of the expiration day with that of the four non-expiration days to observe the abnormal return effect of the stock market that is caused by each kind of index derivative in the TAIFEX on the expiration day from 2002 to 2007. On the days immediately before(T-1) and after (T+1) the expiration day is a natural choice for the purpose of examining the abnormal price or volume effect. We also capture 5 trading days before (T-5) and after (T+5) the expiration day as comparison days to alleviate the possible day-of-the-week effect, which has been extensively documented. The results are reported in Table 2, where Panel A reports the differences in terms of average return of the stock market between the expiration day and the four non-expiration days. We find that in the whole period, compared with the day after the expiration day, the phenomenon of abnormal return effect is signicant at the 5% level. However, there is no abnormal return effect during the sub-periods. Panel B shows the difference in terms of the average return of the stock market during the last hour before closing between the expiration day and the four non-expiration days. Our findings indicate that the stock market does not exhibit any abnormal return during the last hour before closing on the expiration day throughout the sample period or sub-periods. However, the returns are generally smaller during the closing of the market.

Furthermore, we compare the average return of the settlement day with that of the four non-settlement days. The results are presented in Table 3, where Panel A reports the differences in terms of the average return of the stock market between the settlement day and the four non-settlement days. Our findings show 5% significance levels on the fifth day before the settlement day and the day before the settlement day. Nevertheless, the abnormal return effect apears in the sub-period during the period (2003/07~2005/03) after T5F entered the maeket, the difference on terms of the average return of the stock market between the settlement day and the fifth day before the settlement day is 0.009946, which is significant at 1% level. Another two non-settlement days also exhibit 5% significant level. We also compared the first 15-minute average return between the settlement day and four non-settlement days. However, the results in Panel B indicate that the stock market still does not show abnormal return effects during the first 15-minute after opening on the settlement day, but in general returns are larger during the opening of the market.

Table 2: Abnormal Return Effect of TAIFEX Index on Expiration Day

	T-5	T-1	T+1	T+5
Panel A: Whole Expiration Day (T)				
2002/01~2007/12	-0.00236	-0.0004	-0.00449	-0.00152
(whole period)	(-1.29906)	(-0.20147)	(-2.28097)*	(-0.78446)
2002/01~2003/06	-0.00626	-0.00328	-0.00713	-0.00483
	(-1.53043)	(-0.64518)	(-1.52691)	(-0.92464)
2003/07~2005/03	0.002399	0.00289	-0.00568	4.67E-05
	(0.643863)	(0.76974)	(-1.43194)	(0.013502)
2005/04~2006/03	-0.00366	-0.00264	-0.00438	-0.00283
	(-0.98258)	(-0.77874)	(-1.09328)	(-0.83655)
2006/04~2007/12	-0.00299	6.41E-05	-0.0011	0.000602
	(-1.02851)	(0.019007)	(-0.37741)	(0.202287)
Panel B: Last Hour on Expiration Day				
2002/01~2007/12	-0.00082	-0.00051	0.002316	-0.00116
(whole period)	(-1.0435)	(-0.62879)	(-0.41363)	(-1.39457)
2002/01~2003/06	-0.00199	-0.00214	-0.00158	-0.00216
	(-0.89861)	(-0.99113)	(-0.65317)	(-0.89815)
2003/07~2005/03	0.002235	0.001569	4.67E-05	-0.00011
	(1.797434)	(1.309778)	(0.40157)	(-0.10086)
2005/04~2006/03	-0.00099	-0.00012	-0.00034	-0.00072
	(-0.75639)	(-0.105)	(-0.20106)	(-0.63755)
2006/04~2007/12	-0.00277	-0.00141	-0.0003	-0.00171
	(-2.37684)*	(-0.89549)	(-0.24426)	(-1.13219)

*This table shows the difference in the average return of the stock market between the expiration day and the non-expiration day of TAIFEX during the whole sample period and four sub-periods. To reduce the influence on the empirical results due to the weekly effect, we have chosen four non-expiration days as the comparison base. In parentheses are the statistical values for the modified Levene t-test. The sign ** and * indicate a significance levels of 1% and 5%, respectively.*

Table 3: Abnormal Return Effect of TAIFEX Index on Settlement Day

	(T+1)-5	(T+1)-1	(T+1)+1	(T+1)+5
Panel A: Whole Settlement Day (T+1)				
2002/01~2007/12	0.004757	0.004491	0.002316	0.003273
(whole period)	(2.56111)*	(2.28097)*	(1.35902)	(1.76061)
2002/01~2003/06	0.002733	0.00713	-0.0006	0.0057
	(0.58683)	(1.52691)	(-0.14661)	(1.16605)
2003/07~2005/03	0.009946	0.00568	0.007191	0.008541
	(2.83548)**	(1.43194)	(2.24844)*	(2.65413)*
2005/04~2006/03	0.002165	0.00438	0.002069	-0.00147
	(0.65035)	(1.09328)	(0.56458)	(-0.39645)
2006/04~2007/12	0.002874	0.0011	0.000108	-0.00125
	(0.93066)	(0.37741)	(0.04104)	(-0.44662)
Panel B: First 15-minutes on Settlement Day				
2002/01~2007/12	0.000323	0.000813	0.000745	0.000385
(whole period)	(0.52810)	(1.36979)	(1.19899)	(0.61986)
2002/01~2003/06	-0.00094	0.000704	-0.00111	-0.00033
	(-0.55031)	(0.43923)	(-0.64709)	(-0.19068)
2003/07~2005/03	0.000939	0.000572	0.00119	0.001158
	(0.90607)	(0.56218)	(1.19972)	(1.19926)
2005/04~2006/03	0.001331	0.001234	0.002396	0.001029
	(0.87167)	(0.77282)	(1.41011)	(0.67269)
2006/04~2007/12	0.000221	0.000906	0.00095	-9.1E-05
	(0.29095)	(1.34237)	(1.33277)	(-0.10583)

*This table shows the difference in the average return of the stock market between the settlement day and the non-settlement day of TAIFEX during the whole sample period and four sub-periods. To reduce the influence on the empirical results due to the weekly effect, we have chosen four non-settlement days as the comparison base. The statistical values for the modified Levene t-test are reported in parentheses. The sign ** and * indicate a significance levels of 1% and 5%, respectively.*

Finally, we compare the average return of the expiration day of SGXTW with that of the four non-expiration days to observe the abnormal return effect of the stock market. With the same period and sub-period we compare in the TAIFEX from 2002 to 2007. The results are exhibited in Table 4, where Panel A indicates the differences in terms of the average return of the stock market between the expiration day and the four non-expiration days. For the whole period, compared with the day before the expiration day, the phenomenon of abnormal return effect is significant with a 5% level. We also compare the difference in terms of the average return of the stock market during the last hour before closing between the expiration day and non-expiration day. The results in Panel B show that the returns are generally larger during the last hour before the closing of the market although no one is significant.

Table 4: Abnormal Return Effect of SGXTW on Expiration Day

	T-5	T-1	T+1	T+5
Panel A: Whole Expiration Day (T)				
2002/01~2007/12	-0.00069	0.00396	-0.00166	-0.00151
(whole period)	(-0.32916)	(1.93703)*	(-0.80666)	(-0.68887)
2002/01~2003/06	-0.00338	0.008913	0.000874	-0.00333
	(-0.59663)	(1.94664)	(0.17145)	(-0.56399)
2003/07~2005/03	-0.0004	0.000409	-0.00422	-0.00621
	(-0.11422)	(0.10922)	(-1.04076)	(-1.89485)
2005/04~2006/03	0.00514	0.004745	-0.00272	-0.00115
	(1.28785)	(1.26831)	(-0.63264)	(-0.26424)
2006/04~2007/12	-0.00195	0.001238	-0.00067	0.003105
	(-0.62789)	(0.77007)	(-0.23408)	(0.78213)
Panel B: Last Hour on Expiration Day				
2002/01~2007/12	0.001403	0.001466	7.35E-05	0.000334
(whole period)	(1.41279)	(1.52509)	(0.07661)	(0.34797)
2002/01~2003/06	0.001301	0.003997	0.0015	0.00142
	(0.46983)	(1.62558)	(0.57602)	(0.60459)
2003/07~2005/03	0.001382	-0.00027	-0.00166	-0.00048
	(0.73155)	(-0.15171)	(-1.0696)	(-0.26679)
2005/04~2006/03	0.002092	0.001229	0.000428	-2.5E-05
	(1.20107)	(0.67220)	(0.21570)	(-0.01278)
2006/04~2007/12	0.002818	0.001171	0.000382	0.000456
	(1.06654)	(0.78401)	(0.24723)	(0.29141)

*This table shows the difference in the average return of the stock market between the expiration day and the non-expiration day of SGXTW during the whole sample period and four sub-periods. To reduce the influence on the empirical results due to the weekly effect, we have chosen four non-expiration days as the comparison base. The statistical values for the modified Levene t-test are enclosed on parentheses. The sign ** and * indicate a significance levels of 1% and 5%, respectively.*

By further examining the test outcomes in Table 2, Table3, and Table 4 for abnormal returns effects, we discover that abnormal effects exist on the expiration day regardless of TAIFEX or SGXTW. In addition, when we compare the abnormal returns effects on settlement day of TAIFEX, we find that abnormal return effect also appears on the settlement day. This means that under the special settlement system in TAIFEX, the abnormal return effect also appear on the settlement day.

Price Reversal Effect

If there is a negative price pressure on the underlying stock market during expiration days, as found by Klemlosky (1978), Pope and Yadav (1992), and Stoll and Whaley (1991), then the index price will reverse at the opening of the next trading day. We adopt three measures to capture the possible price reversal. The results of TAIFEX are shown in Table 5 and those of SGXTW are shown in Table 6. First, we compared the stock index returns of TAIFEX between the last hour before the closing of the market and the day after the expiration day. There are 72 expiration day samples during the whole sample period, and 39 of them indicate price reversal or approximately 54.17% of the total number of samples. For the whole period, the price reversal effect is not significant (the p-value of the binomial test is 0.2048). However by observing the sub-period, we find that during the period (2002/01~2003/06) after the TXO entered the markets, the phenomenon of price reversal effect is significant with a 5% level. We also compare the stock index returns of SGXTW. The result shows that there are 72 expiration day samples during the whole sample period, with 38 of them showing price reversal. This finding is not significant (the p-value of the binomial test is 0.2780). However during the sub-period, we discover the period (2005/04~2006/03) after T5F entered the markets, 75% of the total number of samples show price reversal effect, which is significant at the 5 percent level. For the whole period neither the TAIFEX nor the SGXTW shows significant price reversal effect, and the significant results could be found during the sub-periods. Thus, we suggest that the phenomenon of price reversal effect in the stock market appear in some periods.

Table 5: Price Reversal Effect of TAIFEX Index

	Measure 1 Number of Reversals	Measure 2 Number of Reversals	Measure 3 Number of Reversals	Number of Expiration Days
2002/01~2007/12	39 (0.2048)	40 (0.1444)	31 (0.8556)	72
2002/01~2003/06	12 (0.0481)*	11 (0.1189)	9 (0.4072)	18
2003/07~2005/03	11 (0.3318)	13 (0.0946)	15 (0.0133)*	21
2005/04~2006/03	5 (0.6128)	6 (0.3872)	6 (0.3872)	12
2006/04~2007/12	11 (0.3318)	10 (0.5)	10 (0.5)	21

*This table explains the price reversal phenomenon of the stock market when the index derivatives expired during the whole sample period and the four sub-periods. We used three measures to test for the reversal effect. Measure 1 compared the return of the stock market during the last hour before closing on the expiration day with that on the day after the expiration day. Measure 2 compared the return of the stock market during the last hour before closing on the expiration day with that before the 15 minute on the day after the expiration day. Measure 3 compared the return during the first 15 minutes of the settlement day with that during the others times after the first 15 minutes of the settlement day. The null hypothesis is that price reversal occurs "randomly." That is, the probability of reversal was 50%. The p-value were calculated using a binomial distribution and are shown in parentheses. The sign ** and * indicate significance level of 1% and 5%, respectively.*

Second, we compare the last one hour return before the closing of the market of the expiration day and the return of the first 15-minute after the opening of the market of the day after expiration day (settlement day of TAIFEX). The results are presented in Table 6. For the whole sample period comprising 72 expiration days of TAIFEX, the price reversal phenomenon occurs on 40 days, or approximately 55.56%. The price reversal phenomenon of TAIFEX (the p-value of the binomial test is 0.1444) form the expiration day to settlement day is not significant. We also compare the price reversal effect of SGXTW from the expiration day to the next day. However, in the whole sample period, including 72 expiration days, 38 days of them show price reversal, or approximately 52.78%. The price reversal effect of SGXTW (the p-value of the binomial test is 0.2780) is not significant. By observing the results of the price reversal effect of the sub-periods of TAIFEX and SGXTW, there is no significant situation. However, by noting two sub-periods of SGXTW, when T5F entered the market, the phenomenon of price reversal effect becomes more significant after 15 minutes from the opening of the market of the day after expiration day (the p-value of the binomial test is 0.0946). The same situation is observed at the sub-period when MSF and MSO entered the market (the p-value of the binomial test is 0.7020). Thus, we suggest that the settlement system of SGXTW affect the price reversal effect more than that of TAIFEX when comparing the last one hour return before the closing of the market of the expiration day and the return of the first 15-minute after the opening of the market of the day after expiration day.

Finally, for the third price reversal measure, we compare the first 15-minute return after the opening of the market and the return after the first 15-minute of the settlement day of TAIFEX (the day after the expiration day of SGXTW) by using the time after the final settlement price is set as the basis for comparison basis. For the whole sample period of TAIFEX including 72 expiration days, 40 of them show price reversal, or approximately 55.56%. The phenomenon of price reversal effect (p-value of the binomial test is 0.1444) is still not significant for the whole period. However, by observing when T5F entered the market, the phenomenon of price reversal is significant at the 5% level. We also adopt the same measure to capture the possible price reversal of SGXTW in the stock market. There are 72 expiration day samples during the whole sample period, and 31 (43.06%) of them exhibit price reversal. With respect to the whole period, the price reversal effect is not significant on the day after expiration day (the p-value of binomial test is 0.8556). The statistic results from the sub-period are not significant either.

Table 6: Price Reversal Effect of SGXTW

	Measure 1 Number of Reversals	**Measure 2 Number of Reversals**	**Measure 3 Number of Reversals**	**Number of Expiration Days**
2002/01~2007/12	38 (0.2780)	38 (0.2780)	31 (0.8556)	72
2002/01~2003/06	7 (0.7597)	9 (0.4072)	6 (0.8811)	18
2003/07~2005/03	12 (0.1917)	13 (0.0946)	10 (0.5)	21
2005/04~2006/03	9 (0.0193)*	8 (0.0720)	6 (0.3872)	12
2006/04~2007/12	10 (0.5)	8 (0.8083)	9 (0.6682)	21

*This table shows the price reversal phenomenon of the stock market when the index derivatives expired during the whole sample period and the four sub-periods. We used three measures to test for the reversal effect. Measure 1 compared the return of the stock market during the last hour before closing on the expiration day with that on the day after the expiration day. Measure 2 compared the return of the stock market during the last hour before closing on the expiration day with that before the 15 minute on the day after the expiration day. Measure 3 compared the return during the first 15 minutes of the day after expiration day with that during the others times after the first 15 minutes of the day after the expiration day. The null hypothesis is that price reversal occurs "randomly." That is, the probability of reversal was 50%. The p-value were calculated using a binomial distribution and are shown in parentheses. The sign ** and * indicate significance level of 1% and 5%, respectively.*

By further comparing the results of these three price reversal measures, we discover that the first measure can capture the phenomenon of price reversal of SGXTW and the earlier period of TAIFEX. However, the second measure can capture the phenomenon of price reversal more on the expiration day of SGXTW. This means that the settlement system of SGXTW reflects the price reversal effect more than TAIFEX. Finally, we find that the third measure, which uses the time after the settlement price of TAIFEX, is able to capture the price reversal effect of TAIFEX only. This means that the mechanism in the TAIFEX, which separates the settlement day from the expiration day, can reduce the probability of price reversal in the stock market on the expiration day and the day after the expiration day. The purpose of using the average trading price during 15 minutes after the opening of the market was to reduce the influence of speculators and to alleviate the disequilibrium caused by market volatility. However, the average trading price during 15 minutes seems not enough as the price reversal effect still exist. Thus, if the TAIFEX were to extend the sampling time in relation to determining the settlement price like in the case of the HSI, the price reversal effect on the stock market would be weakened.

Abnormal Price Volatility

Next, we compare the price volatility on the expiration day with that on the four non-expiration days to observe the abnormal price volatility effect of the stock market that is caused by each kind of index derivative in the TAIFEX on the expiration day. We also focus on the first 15 minute of the settlement day and whole settlement day by comparing the data on the settlement day and the four non-settlement days. The results are shown in Table 7 and Table 8. First, from Table 7, for the whole period we do not observe significant results during the whole trading day and last hour on the expiration day. During the sub-periods, we can only observe a significant result from the whole expiration day. However, from Panel A of Table 8, by observing the whole settlement day, the F-test results exhibit that the abnormal volatility obviously exist at the 5% and 1% significance levels as compared with the fifth day before the settlement day and the day after the settlement day. In the sub-period when the MSF and MSO entered the market the stock market showed signs of abnormal volatility at the 1% significance level on the settlement day. Panel B of Table 8 exhibits that the abnormal volatility effect, with a significance level of 5%, also existed during 15 minute of the market on the settlement day regardless if compared to the fifth day before the settlement or the day before the settlement day. We infer that this phenomenon occurs because of the special settlement mechanism used by the TAIFEX. Speculators attempt to raise or lower the price on the settlement day in order to influence the final settlement price, so that the abnormal volatility was found to be more significant during the settlement day as compared with that on the non-settlement days.

Table 7: Abnormal Volatility Effect of TAIFEX Index on Expiration Day

	Expiration Day (T)	T-5	T-1	T+1	T+5
Panel A: Whole Expiration Day (T)					
2002/01~2007/12	0.01165	0.0095 (1.48958)	0.01226 (0.90323)	0.01198 (0.94659)	0.01144 (1.03699)
2002/01~2003/06	0.01288	0.01118 (1.32638)	0.01734 (0.55163)	0.01507 (0.73007)	0.01806 (0.50828)
2003/07~2005/03	0.01343	0.00959 (1.96249)	0.01032 (1.69485)	0.01226 (1.19986)	0.00786 (2.9184)*
2005/04~2006/03	0.00905	0.00918 (0.97095)	0.00747 (1.46649)	0.01053 (0.73854)	0.00748 (1.46297)
2006/04~2007/12	0.00986	0.00853 (1.33719)	0.01169 (0.71193)	0.00895 (1.21483)	0.00940 (1.09949)
Panel B: Last Hour on Expiration Day					
2002/01~2007/12	0.00477	0.00445 (1.14887)	0.00489 (0.95173)	0.00564 (0.71501)	0.00513 (0.86426)
2002/01~2003/06	0.00734	0.00559 (1.72876)	0.00550 (1.78536)	0.00715 (1.05448)	0.00707 (1.07960)
2003/07~2005/03	0.00326	0.00456 (0.51129)	0.00436 (0.55881)	0.00622 (0.2742)	0.00375 (0.75374)
2005/04~2006/03	0.00283	0.00356 (0.63062)	0.00286 (0.97822)	0.00508 (0.31005)	0.00271 (1.08795)
2006/04~2007/12	0.00404	0.00322 (1.57203)	0.00591 (0.4672)	0.00405 (0.99677)	0.00560 (0.52092)

*This table presents the estimates of price volatility for the stock market (measured by standard deviation of the returns) on the expiration day and the non-expiration day of TAIFEX during the whole sample period and the four sub-periods. To reduce the influence on the empirical results due to the weekly effect, we chose four non-expiration days as the comparison base. We compared the standard deviation of the stock market on the expiration day and the non-expiration day by means of an F-test. The pooled F-statistics are reported in parentheses. The signs ** and * indicate that the price volatility on the expiration day is higher than that on the non-expiration day with significance level of 1% and 5 %, respectively.*

We also investigate the abnormal price volatility on the expiration day of SGXTW as a comparison. The statistic results are shown in Table 9. As Panel A of Table 9 indicates, during the whole trading day and the different sub-periods, as various index derivatives entered the market consecutively, the abnormal volatility effect of the stock market does not appear to be significant. Meanwhile, from Panel B of Table 9, by observing the last hour before the closing of the market, the statistical results from the whole period are not significant. However, by observing the sub-periods, abnormal volatility appears. In particular, during the sub-period (2005/05-2006/03) after MSF and MSO enter the markets, the F-test result exhibited a 1% significance level for the existence of abnormal volatility on the expiration day as compared with that on the fifth day before the expiration day. We infer that the launch of MSF and MSO affect the spot market both on TAIFEX and SGXTW.

<u>Abnormal Volume Effect</u>

Finally, we investigate the abnormal volume effect of the stock market when index derivatives in the TAIFEX and SGXTW expire. The results of the tests are shown in Table 10 and Table 11, respectively. We adopt two measures to capture the possible abnormal volume effect in the stock market. By observing table 10, the first measure which uses the last hour of the expiration day as a basis, shows that the trading volume of the stock market on the expiration days of TAIFEX is not significantly larger than that on non-expiration days. For the whole sample period of 72 expiration days, we find that there were 28 expiration days whose trading volumes are larger than that on non-expiration days, or about 38.89%. Both the whole sample period and sub-period results show that no significant abnormal volume exists during the last hour on the expiration day (the p-value of the binomial test is 0.9618). With the same measure, we set the last hour of the expiration as a basis for the observed abnormal volume effect on the stock market on the expiration days of SGXTW. From Table 11, for the whole sample period comparing 72 days, there are 49 expiration days with larger trading volume than non-expiration days, or about 68.06% of total number of samples. The phenomenon of abnormal volume effect is significant.

Table 8: Abnormal Volatility Effect of TAIFEX Index on Settlement Day

	Settlement Day (T+1)	(T+1)-5	(T+1)-1	(T+1)+1	(T+1)+5
Panel A: Whole Settlement Day (T+1)					
2002/01~2007/12	0.01198	0.00959	0.01165	0.00798	0.01006
		(1.559)*	(1.056421)	(2.2534)**	(1.41637)
2002/01~2003/06	0.01507	0.01160	0.01288	0.00885	0.01377
		(1.68927)	(1.369741)	(2.90211)*	(1.19837)
2003/07~2005/03	0.01226	0.00909	0.01343	0.00802	0.00819
		(1.82027)	(0.83343)	(2.33532)*	(2.24302)*
2005/04~2006/03	0.01053	0.0047	0.00905	0.00710	0.00662
		(4.989)**	(1.354015)	(2.19849)	(2.52661)
2006/04~2007/12	0.00895	0.01078	0.00986	0.00777	0.00916
		(0.68828)	(0.823164)	(1.32429)	(0.95392)
Panel B: First 15-Minutes on Settlement Day					
2002/01~2007/12	0.00399	0.00308	0.00307	0.00341	0.00338
		(1.6776)*	(1.688)*	(1.36677)	(1.39702)
2002/01~2003/06	0.00551	0.00429	0.00399	0.00474	0.00473
		(1.64783)	(1.91141)	(1.34915)	(1.35957)
2003/07~2005/03	0.00318	0.00328	0.00341	0.00325	0.00308
		(0.94057)	(0.86824)	(0.95775)	(1.06679)
2005/04~2006/03	0.00500	0.00173	0.00238	0.00311	0.00169
		(8.313)**	(4.4255)*	(2.5815)	(8.7346)**
2006/04~2007/12	0.00234	0.00252	0.00202	0.00221	0.00316
		(0.86427)	(1.34305)	(1.12085)	(0.55016)

*This table presents the estimates of price volatility for the stock market (measured by standard deviation of the returns) on the settlement day and the non-settlement day of TAIFEX during the whole sample period and the four sub-periods. To reduce the influence on the empirical results due to the weekly effect, we chose four non-expiration days as the comparison base. We compared the standard deviation of the stock market on the expiration day and the non-expiration day by means of an F-test. The pooled F-statistics are reported in parentheses. The signs ** and * indicate that the price volatility on the expiration day is higher than that on the non-expiration day with significance level of 1% and 95 %, respectively.*

Table 9: Abnormal Volatility Effect of SGXTW on Expiration Day

	Expiration Day (T)	T-5	T-1	T+1	T+5
Panel A: Whole Expiration Day (T)					
2002/01~2007/12	0.01125	0.01348	0.01320	0.01338	0.01453
		(0.69661)	(0.72611)	(0.70731)	(0.6)
2002/01~2003/06	0.01465	0.01905	0.01275	0.01590	0.01997
		(0.59159)	(1.32038)	(0.84927)	(0.53827)
2003/07~2005/03	0.00998	0.01215	0.01397	0.01568	0.01099
		(0.67477)	(2.8211)	(0.4051)	(0.82514)
2005/04~2006/03	0.01046	0.00905	0.00766	0.01057	0.01033
		(1.33609)	(1.86465)	(0.97867)	(1.02525)
2006/04~2007/12	0.00943	0.01041	0.01387	0.00926	0.01542
		(0.81956)	(0.4621)	(1.03727)	(0.373)
Panel B: Last Hour on Expiration Day					
2002/01~2007/12	0.00608	0.00569	0.00544	0.00541	0.00521
		(1.14157)	(1.24902)	(1.26116)	(1.36243)
2002/01~2003/06	0.00836	0.00825	0.00624	0.00722	0.00502
		(1.02662)	(1.79417)	(1.34079)	(2.772)*
2003/07~2005/03	0.00568	0.00627	0.00594	0.00428	0.00591
		(0.82154)	(0.91386)	(1.75897)	(0.92382)
2005/04~2006/03	0.00561	0.00220	0.00293	0.00397	0.00323
		(6.495)**	(3.6829)*	(1.99703)	(3.029)*
2006/04~2007/12	0.00420	0.00313	0.00541	0.00570	0.00575
		(1.79256)	(0.60106)	(0.54114)	(0.5324)

*This table presents the estimates of price volatility for the stock market (measured by standard deviation of the returns) on the expiration day and the non-expiration day of SGXTW during the whole sample period and the four sub-periods. To ease the influence on the empirical results due to the weekly effect, we chose four non-expiration days as the comparison base. We compared the standard deviation of the stock market on the expiration day and the non-expiration day by means of an F-test. The pooled F-statistics are reported in parentheses. The signs ** and * indicate that the price volatility on the expiration day is higher than that on the non-expiration day with significance level of 1% and 95 %, respectively.*

at the 1% level. Moreover by observing the sub-periods with various index derivatives entering the market continually, the abnormal volume effect also emerge with 1% and 5% significance levels after the

MSF and MSO entered the market and after the XIF and XIO entered the market respectively.

Table 10: Abnormal Volume Effect of TAIFEX

	Measure 1 Number of Abnormal Volumes	Measure 2 Number of Abnormal Volumes	Number of Expiration Days
2002/01~2007/12	28 (0.9618)	45 (0.0122)*	72
2002/01~2003/06	5 (0.9519)	9 (0.4073)	18
2003/07~2005/03	7 (0.9053)	14 (0.0392)*	21
2005/04~2006/03	8 (0.073)	9 (0.0193)*	12
2006/04~2007/12	8 (0.8083)	13 (0.0946)	21

*This table presents the test results for abnormal volume in the stock market as the index derivatives of TAIFEX expire during the whole sample period and the four sub-periods. Two comparison bases were used to test the abnormal volume effect. Measure 1 compared the stock market volume during the last one hour on the expiration day with that on the non-expiration day. The null hypothesis of the test is that the probability that the stock market on the expiration day is larger than that on the non-expiration day occurs randomly. That is, the probability is 50%. Measure 2 then compared the stock market volume during the first 15 minutes of the settlement day with that of the non-settlement day. The null hypothesis of the test is that the probability that the stock market volume on the 15-minute period of the settlement day is larger than that on the non-settlement day occurs randomly. That is, the probability is also 50%.. The p-value are calculated using a binomial distribution and shown in parentheses. The sign ** and * indicate a significance level of 1% and 5%, respectively.*

In the second measure for abnormal effect, we set the first 15 minutes of the settlement day (the day after expiration day) of TAIFEX and the day after the expiration day of SGXTW as a basis for comparison. In the Table 10, for the whole sample period of 72 days, 45 of them exhibit abnormal volume effects, or approximately 62.50% of total number of samples. The phenomenon of abnormal volume effect is significant at the 5% level during the 15-minute period for the whole period. We also observe the sub-periods with various index derivatives entering the market. We find that the abnormal volume effect of the stock market is also significantly at the 5% level. Moreover, we observe the first 15 minutes of the day after the expiration day of SGXTW to investigate the abnormal volume effect. The results are shown in Table 11. There are 72 expiration day samples during the whole sample period, of which 26 appear abnormal volume, or about 36.11% of total number of samples. This result shows that the trading volume of the stock market during the first 15 minutes after the opening of the market on the day after the expiration day is not significant. By observing the sub-periods, the abnormal volume effect of the stock market is significant at the 1% level during the period after T5F entered the market.

Table 11: Abnormal Volume Effect of SGXTW

	Measure 1 Number of Abnormal Volumes	Measure 2 Number of Abnormal Volumes	Number of Expiration Days
2002/01~2007/12	49 (0.0006)**	26 (0.9878)	72
2002/01~2003/06	11 (0.1189)	15 (0.0007)**	18
2003/07~2005/03	13 (0.0946)	9 (0.6682)	21
2005/04~2006/03	10 (0.0032)**	6 (0.3872)	12
2006/04~2007/12	15 (0.0133)*	8 (0.8083)	21

*This table presents the test results for abnormal volume in the stock market as the index derivatives of TAIFEX expire during the whole sample period and the four sub-periods Two comparison bases were used to test the abnormal volume effect. Measure 1 compared the stock market volume during the last one hour on the expiration day with that on the non-expiration day. The null hypothesis of the test is that the probability that the stock market on the expiration day is larger than that on the non-expiration day occurs randomly. That is, the probability is 50%. Measure 2 then compared the stock market volume during the first 15 minutes of the day after the expiration day with the four comparison days. The null hypothesis of the test is that the probability that the stock market volume on the 15-minute period of the day after the expiration day is larger than that on the four comparison days occurs randomly. That is, the probability is also 50%. The p-value are calculated using a binomial distribution and shown in parentheses. The sign ** and * indicate a significance level of 1% and 5%, respectively.*

By further comparing the results for these two abnormal volume measures, we find that the first measure, setting the last hour period of the expiration day as the basis for comparison, can catch the abnormal volume effect on SGXTW more clearly. However, the second measure, setting the first 15-minute period of the settlement day as the basis for comparison, is able to capture the abnormal volume effect on TAIFEX more clearly. This implies that the two different settlement mechanisms in the SGXTW and TAIFEX determine the abnormal volume effect.

The change of the settlement mechanism can reduce the abnormal volume effect on the expiration day. However, the abnormal volume effect seems to be moved to the 15-minute period of the settlement day. Therefore, if the TAIFEX could extend the sampling time for determining the settlement prices as in the case of the HIS, the abnormal volume on the stock market would be further weakened.

SUMMARY AND CONCLUSIONS

This study uses data on the Taiwan Weighted Stock Price Index and MSCI Taiwan Index to examine the stock market effect of index derivatives on the TAIFEX and SGXTW. The data covers derivatives expiring over the period from January 2002 through December 2007. Seven index derivatives entered the market consecutively from 2002 to 2007. We also divide the whole period to four sub-periods. The differences in terms of returns, volatility, and trading volumes between the expiration day and the non-expiration day are compared and tested in order to identify possible expiration effects.

The empirical results show that the expiration effects of the index derivatives in the TAIFEX appear on the expiration day. It also shows that dividing the settlement day from expiration day cannot reduce the expiration effect effectively. Moreover, under the special settlement procedure on the TAIFEX, using the 15 minutes average price on the settlement day as the settlement price, the expiration effects of index derivatives appear on the settlement day. This result shows that using the opening 15-minute period on the settlement day to determine the final settlement price does not prevent speculators from manipulating the market during this period, but rather extends the period of expiration day effect. The empirical results also show that the expiration effects of the index derivatives in the SGXTW occur on the expiration day. This means that the expiration effect occur in the same stock market no matter what the mechanism or where is market is.

Based on our empirical results that the expiration effect appear both on the expiration day and settlement day of TAIFEX, we argue that the TAIFEX's unique settlement procedure does not successfully reduce expiration effects. Moreover the expiration day of SGXTW also generates the expiration effect. Therefore we close by suggesting that regulators consider changing the settlement procedure on Hang Seng Index derivatives if they are eager to reduce the expiration effects. We also suggest the SGX consider changing their settlement mechanism to prevent affecting the stock market of another country.

APPENDIX

Appendix 1: The Features of the Index Derivative in TAIFEX

Item	Description
Delivery Months	Spot month, the next calendar month, and the next three quarterly months.
Last Trading Day	The third Wednesday of the delivery month of each contract.
Trading Hours	08:45AM-1:45PM Taiwan time. Monday through Friday of the regular business days of the Taiwan Stock Exchange.
Margin	The initial and maintenance margin levels as well as the collecting measures prescribed by the FCM to its customers shall not be less than those required by the TAIFEX. The margin levels will be adjusted and announced by the TAIFEX in accordance with "the Criteria and Collecting Methods regarding the Clearing Margins".
Final Settlement Day	The first business day following the last trading day. All of the open interests after the final settlement day shall be settled on the final settlement price.
Final Settlement Price	The final settlement price for each contract is computed from the first fifteen-minute volume-weighted average of each component stock's prices in the index on the final settlement day. For those component stocks that are not traded during the beginning fifteen- minute interval on the final settlement day, their last closing prices would be applied instead.
Settlement	Cash settlement.

REFERENCES

Aggarwal, R. (1988) "Stock Index Futures and Cash Market Volatility," *Review of Futures Markets,* 7, 290-9.

Bacha, O. and Vila, A. F. (1994) Futures Markets, Regulation and Volatility: the Case of the Nikkei Stock Index Futures Markets, *Pacific-Basin Finance Journal,* 2, 201-25.

Bessembinder, H. and Seguin, P. J. (1992) Futures-Trading Activity and Stock Price Volatility. *Journal of Finance,* 47, 2015-34.

Chamberlain, T. W., Cheung, C. S. and Kwan, C. C. Y. (1989) Expiration-Day Effects of Index Futures and Options: Some Canadian Evidence. *Financial Analysis Journal,* 45, 67-71.

Chow, Y. F., Yung, H. H. M., and Zhang H. (2003) Expiration Day Effects: the Case of Hong Kong. *Journal of Futures Markets*, 23, 67-86.

Gannon, G. L. (1994) Simultaneous Volatility Effects in Index Futures. *Review of Futures Markets,* 13, 1027-65.

Hancock, G. D. (1993) Whatever Happened to the Triple Witching Hour? *Financial Analysts Journal,* 49, 66-72.

Illueca, M. and Lafuente J. A. (2006) New Evidence on Expiration-Day Effects Using Realized Volatility: An Intraday Analysis for the Spanish Stock Exchange. *Journal of Futures Markets*, 26, No.9, 923-938.

Karolyi, A. G. (1996) Stock Market Volatility Around Expiration Days in Japan. *Journal of Derivatives*, 4, 23-43.

Klemkosky, R. C. (1978) The Impact of Option Expirations on Stock Prices. *Journal of Financial and Quantitative Analysis*, 13, 507-518.

Masulis, R. W. (1980) The Effects of Capital Structure Change on Security Prices, *Journal of Financial Economics*, 8, 139-178.

Pope, P. F. and P.K. Yadav (1992) The Impact of Expiration on Underlying Stocks: The UK Evidence. *Journal of Business Finance and Accounting*, 19, 329-344.

Schlag, C. (1996) Expiration Day Effects of Stock Index Derivatives in Germany. *European Financial Management*, 1, 69-95.

Stoll, H. R. and Whaley, R. E. (1987) Program Trading and Expiration-Day Effects. *Financial Analysts Journal,* 43, 16-28.

Stoll, H. R., and Whaley, R. E. (1990) Program Trading and Individual Stock Returns: Ingredients of the Triple-Witching Brew. *Journal of Business*, 63, 165-192.

Stoll, H. R. and Whaley, R. E. (1991) Expiration-Day Effects: What Has Changed? *Financial Analysts Journal,* 47, 58-72.

Stoll, H. R., and Whaley, R. E. (1997) Expiration-Day Effects of the All Ordinaries Share Price Index Futures: Empirical Evidence and Alternative Settlement Procedures. *Australian Journal of Management*, 22, 139-174.

Swidler, S., Schwartz, L., and Kristiansen, R. (1994) Option Expiration Day Effects in Small Markets: Evidence from the Oslo Stock Exchange. *Journal of Financial Engineering*, 3, 177-195.

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ARE AMERICAN AND FRENCH STOCK MARKETS INTEGRATED?

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ABSTRACT

Within a nonlinear framework, this article studies the market integration hypothesis between the French and American stock markets, on a short- and long-term basis. We use two nonlinear Error Correction Models (ECM): the Exponential Switching Transition ECM (ESTECM) and the nonlinear ECM-Rational Polynomial (NECM-RP). Our results provide strong evidence of integration between French and American stock markets. They show that the stock market integration process is non-linear and time-varying and that it has strengthened over time.

JEL: C22, G15

INTRODUCTION

Stock market integration has been studied in several papers (Jeon and Chiang, 1991; Richards, 1995; Bekaert and Harvey, 1995; Heimonen, 2002; Carrieri *et al.,* 2006). Overall, stock market integration can be assimilated to a situation in which the assets in different countries display the same expected risk-adjusted returns. Two stock markets are perfectly integrated if investors can pass from one market to the other without paying any additional costs and if there are possibilities of arbitration which ensure the equivalence of stock prices on both markets. These studies have focused on several questions: Are stock markets integrated? What are the factors of this financial integration? What is the financial integration dynamics? Is the integration process continuous or discontinuous? The answer to these questions is important because it helps understand the degree of interdependence and correlation between stock markets and helps apprehend the international investors' strategies. More interesting, the recent increases in the number of international investors and the recent development of stock markets and financial liberalization have alimented these studies and interrogations.

Exploration of the literature on stock market integration shows that most of the previous studies either examined the integration hypothesis only in a linear framework or used the usual correlation tools, thus limiting the convergence between stock markets to be symmetric and linear. However, the stock market integration dynamics should be nonlinear and asymmetric. The nonlinearity and asymmetry can be justified in different ways through the presence of heterogeneous transaction costs (Dumas, 1992 and Anderson, 1997), the coexistence of different shareholders (De Grauwe and Grimaldi, 2006), information asymmetry and through mimetic behavior (Jawadi, 2008), which can induce discontinuities, persistence and inertia effects in the financial integration dynamic. We use two kinds of Nonlinear Error Correction Models (NECM): the NECM-RP and the ESTECM. The first model considers several potential sources of nonlinearities, due to the abrupt changes in adjustment speeds, can be taken into account and adjustment may be differed according to stock price misalignment, etc. The second model helps to reproduce the asymmetry characterizing the financial integration process and stock price deviations.

The main idea of this paper is thus to check the integration hypothesis between the French and American stock markets in the short and long run, using recent nonlinear econometric modeling. The contribution of this study is twofold. On the one hand, the approach used helps to understand the interdependence, contagion and integration between the French and American stock markets. These questions are of crucial importance with regard to decisions related to international portfolio diversification. On the other hand, using two different NECM would specify a more appropriate convergence process between these stock markets than the one induced by the linear cointegration tools used in previous studies.

The motivation of this paper is twofold. First, it is clear that the fusion between the Euronext group, a merger of Paris', Brussels', Amsterdam's and Lisbon's stock markets, and the NYSE, has affected Paris' and New York's quotations. It has also profited professionals and stimulated the convergence process between the French and American stock markets. Secondly, many stylized facts show significant independencies between these stock markets, notably in periods of crises, scandals and stock crashes. For example, after a prosperous period in 1999, the CAC40 lost 15% in about four months in September 2000 as a result of the first signs of a slowing of American growth. In 2001 and 2002, the CAC40 lost respectively 21.97% and 33.75% because of the American decrease, the fall of new technology markets and the fear of an American recession. The Internet Bubble also showed that both stock markets are strongly correlated and that investors are imitating each others. Indeed, while the NASDAQ recorded a fall of 75% in October 2002, the IT CAC40 fell by 90% between March 20, 2000 and October 8, 2002 (Jawadi, 2008).

More recently, due to a recession risk in the United States and after Georges Bush's declaration on January 18th 2008 concerning his plan to raise the American economy and suppress the "subprime" crisis, the CAC40 lost 6.83% on January 21st 2008, its most important fall since September 2001. The DAX also lost 7.6%, the Eurostoxx50 7.31%, the FTSE 5.48% and the BEL20 5.48%. But two days after the Fed reaction to the decrease of the interest rate by 75 points, a reaction which aimed at damming up the crisis and at restoring the investors' confidence, the CAC40 increased by 2.2%. This indicates that the American conjuncture leads world conjunctures. It also implies that the French and American Stock markets and economies seem to be strongly correlated. This interrelationship was justified differently in the finance literature: the decrease of transaction costs, the convergence of risk premium, the synchronization between markets, the internationalization of companies and investors, the ongoing liberalization process, etc. This article investigates the financial aspect of this interdependence. More precisely, we shall check whether the French and American stock markets are integrated and examine their convergence dynamic. The plan of the article is as follows. Section 2 presents a quick literature review on stock market integration. The NECM's are presented in the third section. Section 4 discusses the empirical results. Section 5 concludes the paper.

LITERATURE REVIEW

The main idea of this section is to present the literature review on stock market integration while explaining and answering two questions: i) How did the previous studies justify financial integration? ii) Which econometric tools have been used to check the financial integration hypothesis?

Stock market integration has been studied by several authors (Gourinchas and Gurgand, 1990; Jeon and Chiang, 1991; Richards, 1995; Bekaert and Harvey, 1995; Heimonen, 2002; and Hardouvelis *et al.,* 2006). According to these studies, there is some evidence of interdependence and convergence between stock markets. Overall, this interdependence seems to be due to fewer cross-country restrictions on stock investment and foreign ownership (Bekaert, 1995), to the contagion effect (Roll, 1988; and Bekaert *et al,.* 2005), to the strong economic ties and policy coordination among countries (Engle and Susmel, 1993), to the global cooperation under technological and financial innovation (Chen *et al.,* 2002), and to the internationalization of companies and institutional investors (Lin *et al.,* 1998).

In practice, stock market integration was checked using different techniques and several results were proposed, either confirming or rejecting the perfect integration hypothesis, for instance, linear cointegration tests. According to the mixed results of previous studies, stock markets come off as very ambiguous: Cho *et al.* (1986) and Wheatley (1988) showed that stock markets are integrated whereas Ko and Lee (1991) suggested that they are segmented. Chan *et al.* (1997) obtained mixed results across different periods. Other studies focusing on return correlations adopted a short-term approach. Their results were also mixed (Campbell and Hamao, 1992; and Ammer and Mei, 1996). Nevertheless, we have noted two remarks. First, these studies retained several definitions of integration and the results vary among these definitions (Heimonen, 2002). Second, most of the previous studies evaluated the integration within a linear framework.

However, linear tests cannot show whether stock markets have become more integrated and whether their integration process is gradual. In other words, the usual linear cointegration cannot detect the partial integration of the stock market because it does not take into account the fact that integration is a gradual and on-going process (Canarella *et al.,* 1990). To remedy these limits, other studies focused on nonlinear modeling. Bekaert and Harvey (1995) showed that the stock market integration degree is time-varying. The authors studied integration using a regime-switching framework. They assumed that stock markets are segmented in a part of the sample and become integrated in another part. Okunev and Wilson (1997) also used nonlinear tests to examine the integration between real estate and stock markets. Recently, Li (2006) has allowed for the risk premium function to be nonlinear as the cause of nonlinear linkages among international stock markets. Overall, these studies justified the nonlinearity and the persistence through the smoothness characterizing the stock market integration, liberalization and barrier removal between countries.

Few studies focus exclusively on the integration between American and French stock markets. Indeed, our exploration of the empirical literature identified only the study of Gourinchas and Gurgand (1990). The authors focused on the CAC40 and Dow Jones over the period September 1987-March 1990. They suggested that the French market is correlated with the American market variations and that there is a transmission effect from Wall Street toward Paris. However, Wall Street can only amplify, reduce, initiate movements and delay the tendencies of the French market; it cannot determine them. Thus, both stock markets are neither completely integrated nor segmented: they are partially integrated. In fact, Gourinchas and Gurgand (1990) established that both indexes have parallel movements and similar tendencies but the French stock market does not systematically nor exactly reflect the American one. They justified this result by the presence of information and transaction costs and currency risk.

However, these results are henceforth questionable because the reduction of transaction costs over the last few years can justify a progressive integration between these markets. Furthermore, the use of new information and communication technologies and the recent fusion of NYSE and Euronext supports the hypothesis of financial integration between American and French markets. To check this hypothesis, we focused on the study of the integration between these markets within a nonlinear framework. The nonlinearity is useful to reproduce the persistence characterizing stock market convergence and the integration dynamic. The next section presents our framework.

NONLINEAR ERROR CORRECTION MODELS

This section presents the nonlinear tools that we used to check the integration hypothesis between the American and French stock markets. First, we present the definition of the NECM as given by the theorem of Escribano and Mira (1998). Second, we describe the nonlinear modeling of these models.

The Nonlinear Error Correction Model Representation: Theorem of Escribano and Mira (1998)

We develop this representation according to Escribano and Mira (1998) and Dufrénot and Mignon (2002). Consider two sets of I(1) variables: an endogenous variable y_t and a vector X_t of K explanatory variables. A NECM is written as follows:

$$\Delta y_t = \alpha_0 + \rho \; z_{t-1} + f(z_{t-1}, \theta) + \sum_{i=1}^{q} \gamma'_i \, \Delta X_{t-i} + \sum_{j=1}^{p} \delta'_j \, \Delta y_{t-j} + \mu_t$$
$$\Delta X_t = \nu_t \tag{1}$$
$$z_t = y_t - \beta' X_t$$

Where: γ'_i and δ'_j are vectors of parameters, ρ is the adjustment term, z_t is the residual term of the linear cointegration relationship. It is assumed that: μ_t and ν_t are mixing processes with finite second-order moments and cross-moments; f is a nonlinear function that is continuously differentiable and that satisfies some regularity conditions: $-2 < \frac{\partial f(zt-1, \gamma)}{\partial z_{t-1}} < 0$, the roots of $\left| 1 - \sum_{i=1}^{p} \delta'_j L^P \right| = 0$, all lie outside the unit circle. μ_t is a martingale difference process with zero mean and constant variance. Under these assumptions, z_t is NED (Near Epoch Dependent) and y_t and X_t are cointegrated with cointegrating vector $(1, -\beta')$. The NED assumption can be tested in the same way as the mixing hypothesis (Dufrénot and Mignon, 2002).

In order to check the integration in a nonlinear framework, we first test the mixing hypothesis. To do so, we use two tests: a nonparametric test given by the KPSS test and a parametric test defined by Lo (1991). Secondly, we reproduce the potential integration process using two particular classes of NECM's: the ESTECM (equation (2)) and the NECM-RP (equation (3)):

$$\Delta y_t = \alpha_0 + \rho_1 z_{t-1} + \sum_{i=0}^{q} \gamma_i \Delta x_{t-i} + \sum_{j=1}^{p} \delta_j \Delta y_{t-j} + (\rho_2 z_{t-1} + \sum_{i=0}^{q} \gamma'_i \Delta x_{t-i} + \sum_{j=1}^{p} \delta'_j \Delta y_{t-j}) \times \left[1 - \exp\left\{-\gamma (z_{t-1} - c)^2\right\}\right] + \mu_t \tag{2}$$

Using NECM's helps study the integration hypothesis in the short- and long-term, while the ESTECM allows gradual reproduction and smooth changes in the integration process, during the successive periods of stock price under- and overvaluation. It also helps detect temporal paths governed by smooth changing regimes and accounts for a slow adjustment mechanism. This modeling was used in several empirical studies to model the asymmetries in the misalignment dynamics of the financial assets (a, b, c and d are the parameters of the rational polynomial function).

$$\Delta y_t = \alpha_0 + \rho_1 z_{t-1} + \sum_{i=0}^{q} \gamma_i \, \Delta x_{t-i} + \sum_{j=1}^{p} \delta_j \, \Delta y_{t-j} + (\rho_2 + \sum_{i=0}^{q} \gamma'_i \, \Delta x_{t-i} + \sum_{j=1}^{p} \delta'_j \, \Delta y_{t-j}) \times \frac{(z_{t-1} + a)^3 + b}{(z_{t-1} + c)^2 + d} + \mu_t \tag{3}$$

The NECM-RP is more general because this modeling is a useful parametric approximation to unknown function forms. Indeed, as suggested by Chaouachi *et al.* (2004), this modeling can take into account many potential sources of nonlinearities (i.e. abrupt changes in adjustment speeds, effects of negative and positive shocks on stock price adjustment, multiple long-term attractors, etc.). In addition, the NECM-RP can highlight asymmetric dynamics between the overvaluation and undervaluation regimes and allows these asymmetric dynamics to be modeled in a more flexible way than with the ESTECM.

Nonlinear modeling for Stock Market Integration

We propose to check the integration between the American and French stock markets using NECM that not only allow us to test stock market integration while taking into account the persistence and the asymmetry characterizing stock price dynamics, but also to specify the integration process per regime. Furthermore, the comovements among various national stock markets may well take nonlinear forms, and a lot more evidence of stock market integration can emerge from nonlinear cointegration analysis than from linear cointegration analysis. This empirical investigation will involve several tests. To begin with, we apply the usual unit root tests (Augmented Dickey-Fuller (ADF) and Phillips-Perron (PP) tests) to check the integration order of stock price series. Second, by applying two tests - the KPSS test and *R/S* test - on the residual term ($\hat{z}_t$), we check the mixing hypothesis in order to verify the nonlinear cointegration hypothesis. Third, accepting the mixing hypothesis implies that stock prices are nonlinearly mean-reverting and allows the NECM to be estimated by the Nonlinear Least Squares (NLS) method. These estimations are useful for reproducing the dynamics of the market integration process.

DATA AND EMPIRICAL RESULTS

Data and Preliminary Tests

Our empirical study concerns the daily American and French stock indexes (CAC40 and Dow Jones) over the period of January 1st 1988 – September 28th 2007. Using daily data provides us with enough observations for nonlinear modeling. Both indexes are expressed in US dollars so as to have homogenous data and to take currency risk into account. This stock price data was obtained from DATASTREAM while the exchange rate series were obtained from the Federal Reserve Bank of St. Louis. The indexes are transformed in logarithm.

First, both ADF and PP tests show that the CAC40 and Dow Jones are I(1). Second, we estimated a static regression modeling the CAC40 on a constant before estimating the Dow Jones and applying usual linear cointegration tests (ADF). Table 1 shows the results.

By comparing the linear cointegration statistic with the critical value of Engle and Yoo (1987), we rejected the hypothesis of linear cointegration. This implies that the American and French stock markets are segmented. However, it is important to be careful when analyzing this result as the rejection of linear cointegration hypothesis can be explained differently: misspecification, low power of linear cointegration tests, etc. In order to remedy these limits, we proposed to check the cointegration hypothesis within a nonlinear framework using "mixing" tests. Accepting the nonlinear cointegration hypothesis implies that the CAC40 is nonlinearly mean-reverting toward the Dow Jones and that both stock markets are nonlinearly integrated.

Table 1: Linear Cointegration and Mixing Tests

		Coefficients (t-statistic)
Constant		1.12 (37.87)
Dow Jones		0.79 (235.3)
$\overline{R}^2$		0.91
Linear Cointegration test: ADF		-2.32
Mixing Tests		
KPSS		**R/S**
l_4	l_{12}	**Andrews**
0.44	**0.26**	2.4^*

This table presents the estimation results of the long-run relationship and those of the linear cointegration tests. It also gives the results of the nonlinear cointegration tests. The values between brackets are the t-ratio. (*) denotes the rejection of the null hypothesis at the 5% significance level.

Third, we applied, two "mixing" tests: the KPSS test and the *R/S* test. The first one is a nonparametric test, which tests the null hypothesis of "mixing" against its "non-mixing" alternative. The second one is a parametric test that tests the null hypothesis of null or short-range dependence (mixing) against its "non-mixing" or long-range dependence alternative. As far as the KPSS is concerned, we empirically retained the values recommended by Schwert (1989) for the truncation parameter: $l_4 = \text{int}\left[4\left(\frac{T}{100}\right)^{\frac{1}{4}}\right]$ and $l_{12} = \text{int}\left[12\left(\frac{T}{100}\right)^{\frac{1}{4}}\right]$ where T is the number of observations (Int [.] denotes the interior part). Concerning the choice of q for the *R/S* test, we used the value of Andrews (1991) corresponding to the following formula: $q_t = [K_T]$, where $K_T = \left(\frac{3T}{2}\right)^{\frac{1}{3}}\left(\frac{2\hat{\rho}}{1-\hat{\rho}^2}\right)^{\frac{2}{3}}$, $[K_T] = \text{int}(K_T)$ and $\hat{\rho}$ is the first-order autocorrelation coefficient.

Our mixing tests did not show the same conclusion insofar as the mixing hypothesis is rejected according to the test of Lo (1991), whereas the KPSS test accepted this hypothesis. In what follows, we retained the mixing hypothesis and preferred the KPSS test because it is a nonparametric one which is more powerful than the *R/S* test (parametric test). To check this mixing hypothesis and assess whether French and American stock prices are nonlinearly mean-reverting, two NECM - the ESTECM and the NECM-RP - will be examined in the next step.

ESTIMATION RESULTS OF NECM

The estimation of the NECM is done through the NLS method and the use of a nonlinear optimization program. We need to define initial values in order to start the nonlinear estimation. We proceed in many steps according to the methodology proposed by Escribano and Mira (1998) and developed by Van Dijk *et al.* (2002) to estimate the NECM. The results are presented in Table 2.

Several conclusions may be drawn from Table 2. First, for both models, the US parameters are statistically significant in the two regimes, showing the statistical dependence of the French stock market toward the American one. The AR parameters for the CAC40 are more significant in the first regime, indicating that the CAC40 depends on its past tendencies while it becomes more correlated to the actual and previous American stock price variations in the second regime.

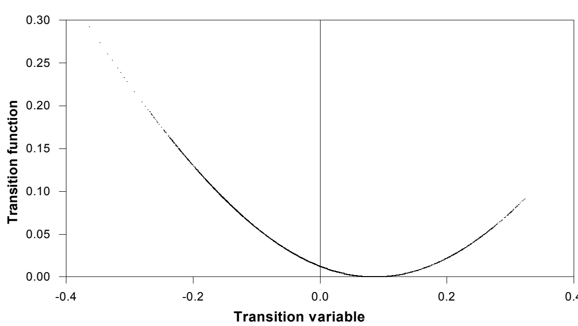
Second, for the ESTECM, γ and c are also statistically significant validating the choice of this nonlinear representation. The estimated value for γ is low, thus confirming the slowness in the transition between regimes. More interestingly, the estimation of the nonlinear adjustment terms (ρ_1 and ρ_2) shows an important result and feature which had also been pointed out by several previous empirical studies (Michael *et al.,* 1997; Peel and Taylor, 2000). Indeed, ρ_1 is negative but non significant and ρ_2 is negative and statistically significant at 5%. In addition, the sum ($\rho_1 + \rho_2$) is negative, implying a significant nonlinear error correction adjustment dynamic for the French index and showing a nonlinear integration process between the Paris and New York stock markets. This also indicates that in the first regime, while the French deviations are small, the CAC40 cannot follow the Dow Jones fluctuations and both markets are rather segmented. By contrast, in the second regime, while French stock price deviations are large, a nonlinear integration process will be active and its convergence speed will increase with the stock price deviation size, as it is shown in Figure 1. This figure also shows that the estimated function did not exceed 30%, reflecting the persistence, asymmetry and smoothness associated with the stock price adjustment dynamic. Applying the misspecification tests, we showed that the residuals of ESTECM are mixing and stationary.

Table 2: NECM Estimation Results

Coefficients	ESTECM (3,1)	NECM-RP
α_0	-0.0048 (-0.01)	-0.0009 (-0.42)
ρ_1	-0.0052 (-0.62)	-0.026 (-1.43)
ρ_2	-0.0024* (-2.27)	0.011 (1.25)
δ_1	0.0768* (2.72)	0.4973* (2.25)
δ_2	-0.0335 (-1.27)	-0.0624 (-0.30)
δ_3	0.0779* (2.88)	-0.3131 (-1.51)
γ_0	0.3604* (8.48)	1.372* (4.67)
γ_1	0.3746* (9.95)	1.507* (4.73)
α'_0	0.0003 (0.34)	0.0021 (0.27)
δ'_1	-0.0513 (-1.36)	0.1961** (1.77)
δ'_2	-0.0041 (-0.12)	0.0140 (0.13)
δ'_3	0.0619** (1.76)	0.139 (1.33)
γ'_0	0.2363* (4.75)	-0.436* (-2.94)
γ'_1	0.0658 (1.27)	-0.5505* (-3.43)
γ	1.707* (1.98)	-
C	0.0856* (5.94)	-
ADF^{GLS}	-50.64	-35.13

This table reproduces the estimation results of the two NECM. The second column presents the estimation results of the ESTECM whereas the estimation results of the NECM-RP are given in the third column. The values in brackets are the t-statistic of nonlinear estimators. (*) and (**) denote significance respectively at 5% and 10%.

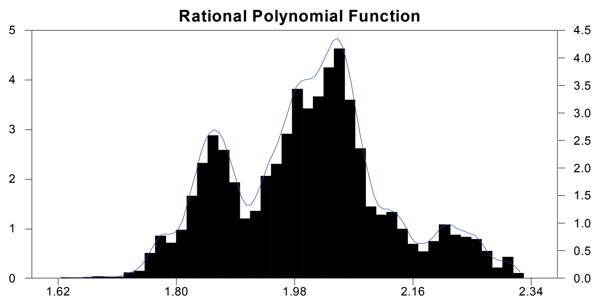
Figure 1: Exponential Transition Function



This figure reproduces the graph of the exponential transition function which describes the transition between the central and upper regimes.

Third, in order to simplify convergence of the algorithm, we estimated the NECM-RP under the following restrictions: a = c = d =1 and b = 0 as in [Chaouachi *et al.* (2004)]. The estimation results show a significant correlation between American and French indexes. The analysis of the histogram of the rational polynomial function (figure 2) which is plotted in function of the estimated misalignment values ($\hat{z}_{t-1}$) shows that the NECM-RP has captured the asymmetry in the integration process between American and French stock markets. Indeed, it shows a bimodal density with two modes of unequal heights. The coexistence of these unequal modes reflects the important and extreme stock price deviations between the regimes of segmentation and integration. This asymmetry in the distribution of the rational polynomial function shows the persistence and the smoothness in the integration process between the Paris and New York stock markets. It also indicates the presence of an integration process which is activated by regime, varying over time and increasing, particularly if integration was observed in the preceding phase.

Figure 2: Histogram of the Rational Polynomial Function



This figure reproduces the graph of the rational polynomial function that characterizes the NECM-RP.

CONCLUSION

The main contribution of this article is to study the stock market integration between France and the United States of America within a nonlinear framework and to specify the integration dynamic process using nonlinear cointegration tools. The estimation results indicated a nonlinear financial integration between Paris and NYSE and established that the CAC40 and Dow Jones are nonlinearly mean-reverting. They also showed that the ESTECM and NECM-RP could respectively reproduce the slow convergence in the integration process and capture the asymmetry and the persistence associated with the nonlinear dynamic integration between France and the United States of America. However, this study was limited to two developed countries: France and the United States, and limited to two kinds of nonlinearity. Therefore, one possible extension would be to check the integration process for other markets belonging to the group of the G8 countries, and also check the integration hypothesis for emerging stock markets. It would also be interesting to test the integration hypothesis between stock markets while looking for other kinds of nonlinearity and modeling. Alternatively, our methodology may be used to check for nonlinear integration using microeconomic data from smaller sectors and firms.

REFERENCES

Ammer J., and Mei, J. (1996), “Measuring International Economic Linkages with Stock Market Data”, *Journal of Finance*, 51, pp. 1743-1763.

Andrews, D. (1991), “Heteroscedasticity and Autocorrelation Consistent Covariance Matrix Estimation”, *Econometrica*, 59, 817-858.

Anderson, H., M. (1997), “Transaction Costs and Nonlinear Adjustment Towards Equilibrium in The US Treasury Bill Markets", *Oxford Bulletin of Economics and Statistics*, Vol.59, pp.465-484.

Bekaert, G. (1995), “Market Integration and Investment Barriers in Emerging Equity Markets”, *The World Economic Review*, Vol. 9, n° 1, pp. 75-107.

Bekaert, G., and Harvey, C. (1995), “Time Varying World Market Integration”, *Journal of Finance*, 50(2), pp. 403-44.

Bekaert, G., Harvey, C., and Ng, A. (2005), "Market Integration and Contagion", *Journal of Business*, vol. 78, pp. 39-69.

Canarella, G., Pollard, S., K., and Lai, K., S. (1990), "Cointegration between Exchange Rates and relative Prices: Another View", *European Economic Review*, 34, pp.1303-22.

Carrieri, F., Errunza, V., and Hogan, K. (2006), "Characterizing World Market Integration Through Time", *Journal of Financial and Quantitative Analysis*, forthcoming.

Chan, K., C., Gup, B., E., and Pan, M., S., (1997), "International stock market efficiency and integration: a study of eighteen nations", *Journal of Business Finance and Accounting*, 24, pp. 803–13.

Chaouachi S., Dufrénot, G., and Mignon, V. (2004), "Modelling the Mesalignements of the Dollar-Sterling Real Exchange Rate: A Nonlinear Cointegration Perspective", *Economics Bulletin*, Vol. 3, 19, pp.1-11.

Chen, G., Firth, M., and Rui, O., M. (2002), "Stock market linkages: evidence from Latin America", *Journal of Banking and Finance*, 26, pp. 1113–41.

Cho D., Eun C., S., and Senbet, L., W. (1986), "International arbitrage pricing theory: an empirical investigation", *Journal of Finance*, 41, 2, pp. 313–29.

De Grauwe, P., and Grimaldi, M., (2006), "Heterogeneity of Agents, Transaction Costs and the Exchange Rate", *Journal of Economic Dynamics and Control*, forthcoming.

Dufrénot, G., and Mignon, V. (2002), *Recent Developments in Nonlinear Cointegration with Applications in Macroeconomics and Finance*, Kluwer Academic Publishers.

Dumas B. (1992), "Dynamic Equilibrium and the Real Exchange Rates in a Spatially Separated World", *Review of Financial Studies*, Vol. 2,pp. 153-180.

Engle, R., F., and Susmel, R. (1993), "Common volatility in international equity markets", *Journal of Business and Economic Statistics*, 11, 2, pp. 167–76.

Escribano, A., and Mira, S. (1998), "Nonlinear Cointegration and Nonlinear Error Correction", Working Paper, Universidad Carlos III, Madrid.

Gourinchas, P., O., and Gurand, M. (1990), « L'autonomie du marché boursier français : les marchés français et américain sont-ils segmentés ? », *Economie et Statistique*, n° 236, octobre, 81-94.

Hardouvelis G., Malliaropoulos, P., and Priestley, D., (2006), "EMU and Stock Market Integration", *Journal of Business*, vol. 79, pp. 365-392.

Heimonen, K., (2002), "Stock market integration: evidence on price integration return convergence", *Applied Financial Economics*, Vol.12, pp. 415-429.

Jawadi, F. (2008), "Threshold Mean Reversion in Stock Prices Adjustment" in *Risk Management and Value Valuation and Asset Pricing*, Jean Luc Prigent *et al. (Eds),* World Scientific Publishing.

Jeon, B., N., and Chiang, T., C. (1991), "A system of stock prices in world stock exchanges: common stochastic trends for 1975-1990?" *Journal of Economics and Business* 43, 329–337.

Ko, K., S., and Lee, S., B. (1991), "A comparative analysis of the daily behaviour of stock returns: Japan, the US and the Asian NICs", *Journal of Business Finance and Accounting*, 18, 2, pp. 219–34.

Li X., M. (2006), "A Revisit of International Stock Market Linkages: New Evidence from Rank Tests for Nonlinear Cointegration", *Scottish Journal of Political Economy*, Vol. 53(2), pp.174-197.

Lin, Y., X., Mccrae, M., and Gulati, C., M. (1998), "Cointegration between exchange rates: a generalised linear cointegration model", *Journal of Multinational Financial Management*, 8, pp. 333–52.

Lo, A., W. (1991), "Long-Term Memory in Stock Market Prices", *Econometrica*, 59(5), 1279-1313.

Michael, P., Nobay, A., R., and Peel, D., A. (1997), "Transaction costs and Nonlinear adjustment in Real Exchange Rates: An Empirical Investigation", *Journal of Political Economy*, N°105, pp.862-879.

Okunev, J., and Wilson, P., J. (1997), "Using nonlinear tests to examine integration between real estate and stock markets", *Real Estate Economics*, 25, pp. 487–503.

Peel, D., A., and Taylor, M., P. (2000), "Nonlinear Adjustment, Long-run Equilibrium and Exchange Rate Fundamentals", *Journal of International Money and Finance*, 19, pp.33-53.

Roll, R., (1988), "The international crash of October 1987", *Financial Analyst Journal*, 44, 5, pp. 19–35.

Schwert, G., W. (1989), "Tests for Unit Roots: A Monte Carlo Investigation", *Journal of Business & Economic Statistics*, 7, 147-159.

Van Dijk, D., Teräsvirta, T., and Franses, P., H. (2002), "Smooth Transition Autoregressive Models- A Survey of Recent Developments", *Econometric Reviews*, 21, 1-47.

Wheatley, S. (1988), "Some tests of international equity integration", *Journal of Financial Economics*, 21, 2, pp. 177–212.

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INVESTIGATING THE INFLUENCE OF COUNTRY CREDIBILITY ON THE CHANCE OF CURRENCY CRISIS

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ABSTRACT

This paper deals with the relationship between country credibility and currency crises by using cross-country pooled data with a multinomial logit model. We use Corruption Perception Index (CPI) published by Transparency International, which is a global non-governmental organization devoted to combating corruption, as a proxy for country credibility. The main findings from the multinomial logit models indicate that country credibility has a significant influence on currency crisis. The applied corruption index is statistically significant. Negative coefficients on this index prove a relationship between country corruption and the likelihood of currency crisis. Moreover, marginal effects of corruption index are high in relation to marginal effects of other explanatory variables. These results suggest that approaching currency crisis should capture country corruption itself or any other proxy of country prestige.

JEL: E51,F31, G18

INTRODUCTION

The Asian Crisis confirms that currency crises occur when macroeconomic as well as microeconomic fundamentals experience vulnerabilities. As Radelet and Sachs (1998a:26) stated; in the aftermath of the Asian crisis, many observers have decried widespread corruption and crony capitalism as an underlying cause of the financial crisis. Mc Kinnon and Phil (1997,1999) view the problem in the Asian Crisis as the "over borrowing syndrome" and emphasized the role of moral hazard driven by unregulated banks and financial institutes. Corruption, cronyism and lack of transparency were thought to be the ultimate causes of the large open positions banks took and over borrowing throughout the economy. Wei and Wu (2001) note that crony capitalism and self fulfilling expectations by international creditors are tow frequently mentioned opposing explanations for currency crises. They propose corruption as a linkage between these two alternative explanations. Their main argument is that corruption may affect a country's composition of capital inflows in a way that makes it more likely to experience a currency crisis that is triggered by a sudden reversal of international capital flows.

Ghosh and Ghosh (2002) examined the role of corporate sector vulnerabilities in currency crises by using a methodology called binary recursive tree. Their findings suggest that, countries with poor public sector governance, as proxied by public sector corruption index, are much more likely to have a crisis. Eichengreen and Rose (1999) consider the impact of corruption. Their argument is that it is bad for both international direct investors and creditors but direct investment suffers more and hence the composition of capital inflows favors financial over direct investment. Countries with this problem are more likely to default on bank loans, or to nationalize the assets of foreign direct investors. The need for international investors to pay bribery and deal with extortion by corrupted bureaucrats tends to increase with the frequency and the extent of their interaction with local bureaucrats. A country with high levels of corruption receives substantially less foreign direct investment.

This study intends to investigate the influence of country credibility on probability of currency crises by using cross-country pooled data with the multinomial logit model. Corruption Perception Index (CPI) published by Transparency International, which is a global non-governmental organization devoted to combating corruption, is used as a proxy for country credibility in this study. High levels of corruption are expected to increase the chance of currency crisis. The impact is not direct though, as it affects the currency stability through its impact on Foreign Direct Investments. Moreover, membership in international financial and trade organizations might have a reasonable impact on country credibility. However, when checking for the sample countries membership, each country is found to be a member of some international organization. Nevertheless, membership in the European Union was included in to this study in order to analyze the impacts of being an international organization member on country credibility. The remainder of the paper is organized as follows: The following section provides a review of the related literature. Section three discusses the methodological issues and section four presents the empirical results. Finally, section five includes some final remarks and future extension proposals.

A BRIEF REVIEW OF RELATED LITERATURE

The beginning of the currency crisis literature is attributed to Krugman's (1979) classic paper. Krugman's (1979) work was later simplified and extended by Flood and Garber (1984) and surveyed in Agenor et al. (1992). The so called *first generation* crisis models are based on the exhaustible resource literature. The first generation crisis occurs because of an unfavorable macroeconomic policy incompatible with a fixed exchange rate. In this framework, currency crises are essentially caused by excessive credit expansion, which leads to a gradual but persistent loss of international reserves. Under a virtually fixed exchange-rate regime, economic agents anticipate that the authorities will eventually abandon the parity with the depletion of international reserves.

Flood and Garber (1984) showed that the abandonment of the currency peg typically will be enforced by a speculative attack of rationally acting market participants. As soon as the shadow flexible exchange rate, i.e. the exchange rate that would prevail in the absence of the currency peg, is above the fixed parity, domestic speculators can make a profit in an attack by purchasing foreign exchange from the central bank at the fixed price and reselling it at the market-determined price. However, competition between speculators will ensure that the exchange-rate peg will collapse as soon as the shadow exchange rate equals the currency peg. This kind of currency crisis is often called a "fundamental crisis" since it is a fundamental reason that makes it impossible for the central bank to further stabilize the exchange rate. Without such a fundamental reason no currency crisis will occur in these models (Berlemann et al. 2002).

The understanding of a currency crisis based on first generation models is questioned after 1992 due to their inability to explain the crisis of the European Monetary System (EMS) that happened in the same year. To capture the features of the crises in the European Monetary System (EMS) and in Mexico in the 1990s, a *second-generation* of currency crises models was developed (Obstfeld 1986, 1994). These models studied what happend when government policy reacts to change in private behavior or when the government faces an explicit trade-off between the fixed exchange rate policy and other objectives. In Ozkan and Sutherland (1998), for instance, one benefit derived from maintenance of the peg is to obtain credibility in the fight against inflation.

However, an increase in foreign interest rates will lead to an increase in domestic interest rates, which causes a lower level of output. If foreign interest rates rise beyond a certain level, the cost of maintaining the peg becomes larger than the benefits and policymakers will abandon the peg. Therefore, it is changes in important economic variables, due to certain shocks either domestic or external, that make policymakers abandon the peg. Self-fulfilling expectations and multiple equilibria play an important role in these models. The economy can move from one state of equilibrium to another without any noticeable change in the fundamentals and crisis may develop without changes in economic fundamentals.

Obstfeld (1996) presented some mechanisms through which currency crises with self-fulfilling features, or self-fulfilling crises, erupt. One such mechanism is when expectations of a currency depreciation drive up domestic interest rates in a country with a high public debt. In this case, out of concern for the higher cost of servicing the public debt, the government will abandon the peg. Another mechanism is based on expectations of depreciation, which lead to higher domestic interest rates. In this case, rather than face a possible costly bailout of banks, the government will abandon the peg.

The Asian crises of 1997-98 motivated the development of *third generation* currency crises models with financial issues from both banks' and firms' side being the key elements. As Radelet and Sachs (1998b:2) have argued; the Asian Crisis in 'one sense can be understood as a "crisis of success." By conventional fiscal measures the governments of the afflicted economies were in quite good shape at the beginning of 1997; while growth had slowed and some signs of excess capacity appeared in 1996, none of them faced the kind of clear tradeoff between employment and exchange stability. Krugman (1999) suggested that there are two major approaches dominated in the post-1997 theoretical literature. The first, so called "moral hazard approach" represented Krugman (1998, 1999) as well as Mc Kinnon and Phil (1997,1999) modeled "over borrowing syndrome" and emphasized the role of moral hazard driven by unregulated banks and financial institutes.

Corruption, cronyism and lack of transparency were thought to be the ultimate causes of the large open positions banks took and over borrowing throughout the economy. Krugman (1998) argued that the collapse of a fixed exchange rate regime might occur as the result of moral hazard due to governmental guarantees to the financial sector without an adequate system of banking regulation and supervision. Krugman considered a case of over-guaranteed and under-regulated financial intermediaries. Since these institutions do not have to put any capital up-front, and have the liberty to walk away at no personal cost in case of bankruptcy, they engage in excessive lending.

In the alternative view, the so called "financial fragility approach", represented by Chang and Velasco (1998) and Radelet and Sachs(1998a, 1998b), a self-fulfilling pessimism of international lenders caused financial fragility of the Asian countries. A typical model of this kind was presented in Chang and Velasco (1998). This model is based the bank runs work of Diamond and Dybvig (1983). Chang and Velasco (1998) added currency demand to the utility function, which makes a distinction among different exchange rate-monetary regimes necessary. They provide a detailed account of interactions between bank fragility and exchange rates under different monetary regimes.

According to Radelet and Sachs (1998a:2); East Asia had exposed itself to financial chaos because its financial systems were riddled by insider dealing, corruption, and weak corporate governance, which in turn had caused inefficient investment spending and had weakened the stability of the banking system. Wei and Wu (2001) noted that crony capitalism and self fulfilling expectations by international creditors are often suggested as two opposing explanations for currency crises. They proposed corruption as a linkage between these two alternative explanations. Their main argument is that corruption may affect a country's composition of capital inflows in a way that makes it more likely to experience a currency crisis that is triggered by a sudden reversal of international capital flows.

METHODOLOGICAL ISSUES

Data Description and Transformation

Data were collected from the International Financial Statistics (IFS: CD-ROM Version) of the International Money Fund (IMF) for International Reserves, M2 Money Supply and Domestic Credit series. Real Effective Exchange Rate series data were obtained from JP Morgan. Data about corruption comes from Transparency International. Because of data availability, the sample contains 33 countries listed in Table1. This limited number is mostly a result of data availability. It includes a panel of emerging economies of the period 1991-2002.

The following standard definition of currency crises defines Exchange Market Pressure (EMP) as:

$$EMP = \frac{\Delta e}{e} - \frac{\sigma e}{\sigma R} . \frac{\Delta R}{R} \qquad (1)$$

This is a weighted average of the rate of change of the exchange rate, $\frac{\Delta e}{e}$, and of reserves, $\frac{\Delta R}{R}$, with weights such that the two components of the index have equal sample volatilities, that σe is the standard deviation of the rate of change of the exchange rate and σR is the standard deviation of the rate of change of reserves. This way of defining currency crises has an advantage of including instances where a currency came under severe pressure but the authorities successfully defend it. Another advantage of using this definition is that, it can be use to analyze speculative attacks under both fixed and flexible exchange rate regimes. It is worth noting that in order to construct an exchange market pressure index, some researchers employ the difference between domestic and foreign interest rates, or percentage changes in domestic interest rates, while many others avoid using it because many developing countries do not have reliable interest-rate data.

Table 1: Countries in the Sample

Country	Country	Country	Country	Country	Country	Country	Country
Argentina	Bangladesh	Bolivia	Brazil	Bulgaria	Chile	China	Colombia
Croatia	Cyprus	Czech Rep.	Dominican	Ecuador	Egypt	Estonia	Ghana
Hungary	India	Indonesia	Israel	Korea	Latvia	Lithuania	Malaysia
Mexico	Peru	Philippines	Poland	Slovenia	Russia	Thailand	Turkey
Tunisia							

Kaminsky et.al (1997) listed 103 variables used in various studies to explain the likelihood of a crisis occurring and they accepted 16 out of 103 variables as most useful indicators of currency crises. Then in World Economic Outlook of May 1998 IMF, this list is reduced further to three: the real exchange rate, credit growth and M2/Reserves. Corruption Perception Index is used to reflect country prestige as an exogenous variable and to be consistent with the related studies three explanatory variables that might affect currency crises; M2 money supply/net international reserves (denoted M2/NIR), domestic credit (denoted Dcred), and real effective exchange rates (denoted REER), are taken into account as the other exogenous variables. For real effective exchange rates, using the percentage change would be improper as not all real appreciations indicate greater depreciation pressure. Therefore, a trend was estimated by using Hodrick-Prescot (H-P) filter and deviations from the trend are used. A Hodrick-Prescott filter applied to a time series $\{y_t\}$ produces the filtered series $\{\hat{y}_t\}$ which solves the following minimization:

$$\min \sum_{t=1}^{T} (y_t - \hat{y}_t)^2 + \lambda \sum_{t=2}^{T-1} ((\hat{y}_{t+1} - \hat{y}_t) - (\hat{y}_t - \hat{y}_{t-1}))^2 \ . \qquad (2)$$

The first term is the standard sum of squared residuals; the second term is a penalty factor for changing the slope of $\{\hat{y}_t\}$ too much. λ controls the importance given to the penalty. If $\lambda = 0$ then $\hat{y}_t = y_t$ for all *t;* as $\lambda \to \infty$ the H-P filter becomes a simple least squares linear regression The other two exogenous variables; domestic credit and M2/Reserves are calculated as 12 months percentage changes. Avoiding any seasonality that could be incorporated into the data and keeping away from the risk of deriving misleading interpretations are the reasons for the use of the 12 months percentage change.

Method of Estimation

In order to examine how the probability of currency crisis is affected by country credibility, a maximum – likelihood multinomial logit model is used in this study. Eichengreen, Rose, Wyplosz (1994, 1996), Frankel, and Rose (1996) have used the logit approach in estimating models of currency crises. Recently, Berg and Pattillo (1998) and Kumar et al. (2003) have also analyzed the emerging market currency crises by using probit/logit models. The decision to use Logit or Probit (normal distribution) is purely arbitrary as the logistic and normal distributions are quite similar except that the tails are thicker in the logistic case. This approach defines a crisis indicator equal to one or zero depending on whether a currency crisis does, or does not occur within the specified time. In addition, it considers all variables together and looks only at the marginal contributions of each indicator; it disregards variables that do not contribute information that is not already captured in the other variables (Abiad, 2002).

According to this approach, one can write the probability of having crisis as;

$$P(y=1) = f(\beta' x) \tag{3}$$

where $f(\bullet)$ is a probability distribution function. If we assume a logistic distribution, then;

$$(Y_t = 1) = \frac{\exp(\beta' X_t)}{1 + \sum \exp(\beta' X\ )}\ , \qquad (Y_t = 0) = \frac{1}{1 + \sum \exp(\beta' X_t)} \tag{4}$$

where,

$$Y_t = \begin{cases} 1,\ if\ EMP_t > \mu_{EMP} + 1.75\sigma_{EMP} \\ 0,\ otherwise \end{cases} \tag{5}$$

the parameter vector β is estimated by maximum likelihood and the regression is a standard logit one. μ_{EMP} and σ_{EMP} are the sample mean and the standard deviation of the exchange market pressure index for each country.

For the logit model, the estimated coefficients do not have a direct economic interpretation. Measures that are familiar to economists are marginal effects and elasticities. The marginal effect of the k^{th} explanatory variable on the response probability is obtained from the following formula;

$$\frac{\partial P(Y_t = 1 | X_t)}{\partial X_{kt}} = scale.\beta_k \qquad where, \qquad scale = \frac{\exp(X_t\beta)}{[1 + \exp(X_t\beta)]} \tag{6}$$

ESTIMATION RESULTS

The first step is to estimate the impact of corruption on the probability of currency crisis. I estimate a logit model with the dependent variable being an indicator whether a currency crisis has occurred. The explanatory variable is expressed by the Transparency International corruption index (CPI) that varies from zero to 10. The higher the value of the CPI index, the less corrupted a country is. The result from this model are reported in the first column of Table 2. The model gives a log likelihood of -64, 37. As we can see from the table below, corruption is significant at the 5% level. The negative sign of the coefficient on CPI is in line with the common intuition that the more corrupted country, the less stable the country is and the more likely a currency crisis.

However, we have to remember that the estimation above does not capture the influence of all significant explanatory variables. Since there are omitted variables, the estimation is inconsistent. Therefore, I estimate the baseline model by taking into account all basic explanatory variables. The basic explanatory variables that are included in the logit model are; deviations from real effective exchange rates (denoted REER), M2 money supply/net international reserves (denoted M2/NIR) and domestic credit (denoted Dcred). The result is reported in column 2 of Table 2. The baseline model has a better fit to the data. The log likelihood from the baseline model is -60,33, higher than that from the previous one. The *CPI* remains statistically significant and negative. The coefficients on the other explanatory variables except deviations from real effective exchange rate are statistically insignificant at the 5% significance level, as p-values on them are greater than 0.05. Although there are no omitted variables, there are still inconsistent estimates due to the correlation between explanatory variables. Because of this, we should be careful in interpretation of statistical significance.

I also analyze the marginal effects as opposed to simple coefficients because I want to see the effect of corruption on the probability of currency crisis. For instance, the marginal effect would show us by how much the probability of currency crisis increases when there is a one-unit rise in an independent variable x, i.e. Corruption Perception Index. The third column of the Table 2 shows the marginal effects of each variable on the probability of currency crisis. As can seen from the table, the marginal effect of the *CPI* is relatively high among all others. We can expect that the probability of currency crises in an average country will decrease by 0.046, if the *CPI* rises by 1 point. We should remember that the change by 1 in *CPI* stands a substantial shift, as CPI varies from 0 to 10. Such strong changes do occur often in the unstable countries with low *CPI* values, and are less likely in those with high *CPI* scores.

Table 2: Results for Multinomial Logit Models [a]

Variables	(1) *Introducing Model*	(2) *Baseline Model*	(3) *Marginal Effects*	(4) *CPI Mean*	(5) *EU Membership*	(6) *GCI Model*	(7) *Mar. Effects GCI*
CPI	-0.2653*	-0.2941*	-0.0467*	-0.2182*	-0.0951*	-3.2746**	-0.0584*
	(0.1052)	(0.1221)	(0.1952)	(0.1129)	(0.5433)	(1.1444)	(0.0446)
REER	-	0.0478*	0.0235*	0.0056*	0.0037	0.2014*	0.0617*
		(0.0787)	(0.0765)	(0.0035)	(0.0038)	(0.1398)	(0.0084)
M2/NIR	-	-0.0347	-0.01873	-0.0083	-0.2141	0.0098*	-0.0217*
		(0.0007)	(0.0023)	(0.0072)	(0.0092)	(0.2971)	(0.0061)
DCred	-	-.0.0065	-0.0081	-0.0769	-0.0064	0.0081	0.0094
		(0.0006)	(0.0002)	(0.0702)	(0.0087)	(0.0027)	(0.8010)
Cons	-0.3321	-0.2602	-	-0.6192	-1.2761	-2.3947	-
	0.3750	(0.0089)		(0.4189)	(0.1972)	(0.7944)	
Num. of Obs.	194	194	194	193	306	113	113
Prob>chi2	0.0024	0.0198	-	0.0934	0.0221	0.0004	-
Pseudo R2	0.0714	0.0989	-	0.601	0.0572	0.3971	-
LogLikelihood	-64.3725	-60.3321	-	-67.2392	-99.2612	-23.5235	-

** 5% and ** 1% indicates level of significance. [a]This table shows the results of the Logit regressions on the probability of currency crises. The dependent variable is Exchange Market Pressure Index that is defined in Eqn (1). The independent variables are Corruption Perception Index (CPI),Real Effective Exchange Rate (REER), M2 Money Supply/Net International Reserves (M2/NIR), Domestic Credit (DCred), Global Competitiveness Index (GCI) and European Union (EU) Membership. Columns 1 through 7 report the results of the regression using seven different models of currency crises. The first figure reported in each cell is the regression coefficient and the second is the standard error.*

Country prestige tends to be smoothened along time and even if there is an improvement in reliability, a country can suffer from the previous low place in the Transparency International rank. In order to capture this mechanism, I use an average CPI_t instead of pure CPI_t. The fourth column of Table 2 shows the results of the model that is analyzed with the average *CPI*. The results are disappointing. In the *CPImean* estimation only deviations from the real effective exchange rate is statistically significant. The corruption index coefficient is once again negative. The p-value on *CPImean* is the lowest; although the other explanatory variables (except REER) are statistically insignificant. This can be the result of the autocorrelation. We define *CPImean* as an average of *CPI* from all previous periods and we know that the corruption is correlated with other explanatory variables. Thus, the random term from period *t* is correlated with the previous random terms.

Moreover, membership in international financial and trade organizations might have a reasonable impact on country credibility. However, when checking for the sample countries membership I found that each country is a member of some international organization that could not been treated in the same way. Nonetheless, I decide to include membership in the European Union. The sample consists of developing countries and some countries from the new Member States are included. According to this, it is reasonable to check the influence of EU membership on the probability of currency crises. I show the results of this estimation even though the data is based on the years 1990-2002 and in these years, none of the countries from the sample was a member of the EU.

However, since joining the EU means that a country must satisfy particular criteria and changes in politics for a number of years, I assume that the political climate of those countries must have been decent long before the actual accession. In the fourth model, we use *EU* dummy, which is a variable equal to one for all EU Member States and zero for the others, instead of *CPI*. The results obtained from the forth model, which is reported in the fifth column of Table 2, confirm that membership in the EU has a statistically significant impact on currency crisis probability. The log likelihood from this model is -99,26, lower than that from the all other models. *EU* dummy is statistically significant at the %5 level and there is a negative relationship between EU membership and currency crisis, as the sign of coefficient on the *EU* dummy is negative.

The Transparency International corruption index is one of the most popular measures of corruption. However, there are some other measures of corruption. Global Competitiveness Index (GCI) published by the World Bank Group is included in this study as another corruption index (of different methodology) and the results from two corruption definitions are compared. Here, it is worth noting that the CPI index is based on experts' opinions whereas the GCI index is based on firms' records

The sixth column of the Table 2 shows the results from the model conducted with the *GCI* index, while the seventh column of the Table 2 presents the marginal effect of *GCI* index on the probability of currency crisis. The relationship between corruption and currency crisis is positive, as the sign of coefficient on *GCI* is negative. As can be seen from the sixth column of the Table 2, corruption measured in another way is still statistically significant at the 5% and 1% levels. Besides, the marginal effect of *GCI* is again relatively high and negative. The results of this model show that a unit change in *GCI* increases the probability of a currency crisis by 5.8% in the sample countries.

When considering marginal effects it is important to note that currency crises are rare events and even marginal effects of 5% would be notable. This confirms our thesis that corruption does matter regardless of the measure we use to reflect it. *GCI* can be a better measure of corruption because in such a specified model the pseudo R^2 is much higher. Although, the pseudo R^2 estimated for a multinomial logit model cannot be interpreted as it is in a regular linear model, it still gives an indication of the fit since the crisis definition is not dichotomous. The limitation of using pseudo R^2 as a goodness-of-fit measure for usual logit models is due to binary characteristics of the dependent variable rather than its boundaries. In addition, the fit of the model is relatively high, although our results were obtained from a 113 observation sample.

CONCLUSIONS

Empirical research on corruption is quite a new field of interests to economists. Cross-country analyses aim at finding causes and consequences of corruption. There are number of problems arising from the modeling point of view, mostly connected with the methodology of corruption indicator creation. Such indices are mostly base on subjective opinions of the real level of corruption that is country specific. Understanding of corruption varies across cultures. Nevertheless, there are still forms of power abusing behaviors that fall under the 'corruption' definition regardless the culture. Due to this, researchers commonly agree to rely on some commonly approved corruption indices and conduct many projects in the field of corruption impact on different economic and social characteristics.

This research examines unobserved country credibility. Such a defined explanatory variable is difficult to measure, as there is no standard index of country reliability. Corruption Perception Index is used as a proxy for the unobserved credibility in this study. The estimation results confirm the main intuition. The applied corruption index is statistically significant and negative coefficients on this index prove that the higher corruption, the more likely a currency crisis will occur. Moreover, marginal effects of corruption index are high in relation to marginal effects of other explanatory variables. These results suggest that approaching currency crisis should capture country corruption itself or any other proxy of country prestige. In order to make the proposal of using membership in international trade and financial organizations as a proxy for country credibility useful, I encourage new studies to more detailed analysis of different organizations structure to capture only significant elements and enlarge the sample of countries used in the estimation.

REFERENCES

Abiad, Abdul (2002) "Early Warning Systems for Currency Crises: A Markov-Switching Approach with Application to Southeast Asia," *Ph.D. Dissertation; Philadelphia: University of Pennsylvania*

Agénor, Pierre-Richard, Jagdeep S. Bhandari and Robert P. Flood (1992) "Speculative Attacks and Models of Balance of Payments Crises," *NBER* WP No. 3919, 1992

Berg, Andrew, Pattillo Catherine (1998) "Predicting Currency Crises: The Indicators Approach and an Alternative", *Journal of International Money and Finance,* Vol. 18, p.561-86

Berlemann, Michael, Hristov Kalin and Nikolay Nevosky (2002) "Lending of Last Resort, Moral Hazard and Twin Crises: Lessons From the Bulgarian Financial crises 1996/1997," *William Davidson Working Papers*, No: 464

Chang, Roberto and Andres Velasco (1998) "The Asian Liquidity Crisis," *NBER Working Paper Series*, No. 6796

Eichengreen, Barry, Andrew Rose, and Charles Wyplosz (1996) "Contagious Currency Crises: First Tests," *Scandinavian Journal of Economics*, 98(4), p. 463-84

——— (1999) "The Empirics of Currency and Banking Crises," *NBER Reporter*, Winter.

Flood, Robert and Peter Garber (1984) "Collapsing Exchange Rate Regimes: Some Linear Example," *Journal of International Economics,* 17, p. 1-13

Flood, Robert, and Nancy Marion (1999) "Perspectives on the Recent Currency Crisis Literature," *International Journal of Finance and Economics*, Vol. 4, No. 1, p. 1-26

Frankel, Jeffrey, and Andrew Rose (1996) "Currency Crashes in Emerging Markets: An Empirical Treatment," *Journal of International Economics*, Vol. 41, p. 351–66 (November)

IMF (1998) "World Economic Outlook" (May) Financial Crises: Causes and Indicators

Ghosh, Swati, and Atish Ghosh, 2002 "Structural Vulnerabilities and Currency Crises," *International Monetary Fund Working Paper*: WP/02/9, January

Kaminsky, Graciela, L., Saul Lizondo, and Carmen Reinhart (1997) "Leading Indicators of Currency Crises," *IMF Staff Papers*, 45(1), March, p. 1–48

Kaminsky, Graciela, L., and Carmen Reinhart (1996) "The Twin Crises: Causes of Banking and Balance-of-Payments Crises," *American Economic Review* 89(3), June, p. 473–500

Krugman, Paul (1979) "A model of balance of payments crises," *Journal of Money, Credit, and Banking* 11, p. 311-325

——— (1998) "What happened to Asia?," mimeo Cambridge, Massachusetts: MIT, 1998 http://www.pkarchive.org/crises/DISINTER.html, (date of acces: 16/11/2005)

——— (1999) "Balance Sheets, The Transfer Problem, and Financial Crises," in *International Finance and Financial Crises: Essays in Honor of Robert P. Flood, Jr.,* ed. by Peter Isard, Assaf Razin, and Andrew K. Rose Boston: Kluwer Academic; Washington: International Monetary Fund

Kumar, Manmohan, Uma Moorthy, and William Perraudin (2003) "Predicting Emerging Market Currency Crashes," *Journal of Empirical Finance* 10, p. 427– 454

McKinnon, R. and Pill, H. (1997) "Credible Economic Liberalization and Overborrowing," *American Economic Review*, Papers and Proceedings, 87(2), p.189- 93

——— (1999) "Exchange Rate Regimes for Emerging Markets: Moral Hazard and International Overborrowing", *Oxford Review of Economic Policy*, 15(3), p. 19-38

Obstfeld, Maurice (1986) "Rational and Self-Fulfilling Balance of Payments Crises," *American Economic Review* 76 (Mar), p.72-81

——— (1994) "The Logic of Currency Crises," Cahiers *Économiques et Monétaires*, No. 43, p. 189-204

——— (1996) "Models of Currency Crises with Self-Fulfilling Features," *European Economic Review*, 40, p. 1037-1047

Ozkan, F. Gulcin ve Alan Sutherland (1998) "A Currency Crisis Model with an Optimising Policymaker," *Journal of International Economics*, 44, p. 339-364.

Radelet, Steven and Jeffrey Sachs (1998a) "The East Asian Financial Crisis: Diagnosis, Remedies, Prospects" *Brookings Papers on Economic Activity*, 1998(1), p. 1-74

——— (1998b) "The Onset of the East Asian Financial Crisis," *Harvard Institute for International Development.*

Szczurek, Mateusz (2005) "Currency Crises and Optimal International Liquidity," *Ph.D. Dissertation; University of* Sussex

Wei Shang-Jin and Yi Wu (2001) "Negative Alchemy? Corruption, Composition of Capital Flow, and Currency Crisis," *NBER Working Paper,* No 8187

http://www.transparency.org/ - data for CPI index

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