# ESSAYS ON ASSET PRICES

A Dissertation

by

SANG BONG KIM

Submitted to the Office of Graduate Studies of
Texas A&M University
in partial fulfillment of the requirements for the degree of

DOCTOR OF PHILOSOPHY

August 2008

Major Subject: Economics

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# ABSTRACT

Essays on Asset Prices. (August 2008)

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In this dissertation I explain the relationship among inflation volatility, rational bubbles, and asset prices. In addition, I investigate the transmission of asset prices and volatility among countries.

In the second chapter, which deals with the relationship between inflation volatility and asset prices, my empirical analysis shows that real stock returns tend to co-vary negatively with expected inflation during periods of stable inflation, but co-vary positively with expected inflation during periods of volatile inflation for 16 countries.

To investigate the relationship between rational bubbles and asset prices in the third chapter, I formulate an information error model which allows one to derive the measure of non-fundamentals in stock prices in a straightforward manner. This study provides a new method by specifying rational bubble measures that follow the Weibull distribution. As a result, my empirical analysis is the first step in applying survival analysis to bubbles, and it reveals preliminary evidence that there is the increasing bursting rate at a decreasing rate for extraneous or instrinsic bubbles in the U.S. stock market.

In the fourth chapter, which deals with the transmission of asset prices and volatility, I investigate how the 1997 crisis has changed the Korean market by focusing on price and volatility spillovers from the U.S., Chinese, and Japanese markets. I have used daily stock prices from January 3, 1995 to July 31, 2007 and employed an EGARCH model. New information on stock prices originated in the U.S. market was more transmitted to the Korean market for all periods. The price spillover effect from the Japanese market to the Korean market became stronger from the crisis period. The influence of U.S. and Japanese innovations on market volatility increased after the crisis period. However, the magnitude of spillover effects from the Chinese market to the Korean market remained small and stable between the prior- and post-crisis periods and the volatility spillover effect remained stable for all periods. Asymmetry in the spillover effects on market volatility was pronounced in the Korean market after the financial crisis.

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# CHAPTER I

# INTRODUCTION

There have been many theories and empirial results for explaning asset prices. Inflation volatility and rational bubbles have been important factors. Recently, the transmission of asset prices and volatility among countries also has been one of important topics in asset pricing.

First of all, I consider the relationship between inflation volatility and asset pricing. Conventional wisdom suggests that if investors hold stocks as a hedge against inflation, stock returns and expected inflation should be positively related. Such a positive relation appears to be consistent with the investors' optimal portfolio choice. While this assumed positive relation is intuitively appealing, the empirical evidence on the issue is, at best, not persuasive and, at worst, contradictory. There has been a wealth of evidence that a negative relation between stock returns and inflation has prevailed since the 1950s in the United States and other countries as well.

The relations between real stock returns and expected inflation are contingent on the degree of the volatility of inflation. More specifically, my results suggests that real stock returns tend to co-vary negatively with expected inflation in periods of stable inflation and to co-vary positively with expected inflation in periods of volatile inflation. In this study, I argue that stocks are effective hedges against inflation during periods of volatile inflation.

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This dissertation follows the style of *Econometrica*.

Second, I investigate rational bubbles in asset prices. Broadly speaking, bubbles include both rational and irrational bubbles. Irrational bubbles mean that market participants do not recognize the existence of bubbles. However, rational bubbles describe the fact that market participants recognize the existence of bubbles and expect increasing prices regardless of the internal value of firms. The traditional studies of bubbles focus mainly on the rational aspect.

Although the stock market seems to be overheated, people do not recognize if bubbles exist or not. That is to say, stock prices can be adjusted into appropriate levels so far as the prices are not developed into explosive bubbles. Therefore, Bubbles will be detected only after the market is collapsed. Moreover, economists have understood that it is difficult to test if bubbles exist.

In this study I formulate a model which allows one to derive the measure of non-fundamentals in stock prices in a straightforward manner. This study provides a new method by specifying bubble measures that follow the Weibull distribution, and it is the first attempt to apply the Weibull distribution to the test for rational bubbles. There is not only a parallel between the burst of speculative bubbles and a material's burning out, but there is also a good reason to believe that non-fundamentals in stock prices can be appropriately modeled using the Weibull specification.

Third, I investigate the transmission of stock prices and volatility from the U.S., Chinese, and Japanese markets on the Korean market. Recently, the world economy has become increasingly interdependent through trade, common creditors, and similar macroeconomic trends. The growing integration of world markets raises several fundamental questions. Does the globalization of financial markets precipitate the transmission of information from advanced markets into the relatively small markets? If the small markets become more globalized and liberalized, then information on stock prices produced in a leading market such as the U.S., Chinese, and Japanese markets will be more rapidly disseminated into the small market, thus prompting price spillovers. In fact, since the 1997 financial crisis changed the financial landscape in Asia, the influence of advanced stock markets on the small market has gained steadily.

What is of particular interest from both practical and theoretical perspectives is whether such co-movements in stock prices amplify the volatility of Korean market. The co-movements of stock prices across markets may change stock prices above or below the levels dictated by market fundamentals, potentially creating market volatility. However, if investors become more informed as a result of globalization of markets, the increased interdependence and linkage of financial markets could reduce the transmission of volatility from one market to another. Thus, one interesting hypothesis to be tested concerns whether or not an increased integration of financial markets leads to a reduction in market volatility in the Korean market.

The main purpose of this study is to investigate how the Korean stock market has been affected by the 1997 financial crisis. I am particularly interested in whether the influences of shocks originated in the U.S., Chinese, and Chinese markets on prices and market volatility in the Korean market increased or decreased after the 1997 crisis. It is well documented that the volatility of the Korean market was historically high. Interestingly enough, the volatility of this market has been much dampened since 2000 when the aftermath of the financial crisis has been substantially subdued. This study aims to explore whether much of the slowdown in Korean markets' volatility after the financial crisis in 1997 is a fundamental shift or a temporary fad.

This dissertation is organized as follows: In chaper II, I investigate a relationship between inflation and asset prices. Chapter III discusses rational bubbles in asset prices. Chapter IV describes the transmission of stock prices and volatility from the U.S., Chinese, and Japanese markets on the Korean market. Chapter V contains conclusion.

# CHAPTER II

# INFLATION AND ASSET PRICES

## 2.1. INTRODUCTION

Conventional wisdom suggests that if investors hold stocks as a hedge against inflation, stock returns and expected inflation should be positively related. Such a positive relation appears to be consistent with the investors' optimal portfolio choice. While this assumed positive relation is intuitively appealing, empirical evidence on the issue is, at best, not persuasive and, at worst, contradictory. There has been a wealth of evidence that a negative relation between stock returns and inflation has prevailed since the 1950s in the United States and other countries as well.

The theoretical basis for the assumed positive relation between returns on common stocks and inflation is provided by the extended version of the Fisher effect, which states that the expected rate of return on common stocks consists of a real return and the expected inflation rate and that the real return is not affected by the rate of inflation. The earlier work of Nelson (1976), Jaffe and Mandelker (1976), Bodie (1976), Fama and Schwert (1977), Solnik (1983), and Gultekin (1983) was mainly concerned with whether the Fisher hypothesis holds true in the U.S. stock market. Most of the studies found that nominal stock returns were negatively related to inflation rates over the post-war period, thus rejecting the view that common stocks are effective hedges against inflation.[1]

[1] Fama and Schwert (1977) show that U.S. government bonds were a complete hedge against expected inflation, and private residential real estate was a complete hedge against both expected and unexpected inflation.

Subsequent studies including Fama (1981), Stulz (1986), Kaul (1987), Lee (1992) have been extended to analyze the relationship between real (ex post and ex ante) stock returns and expected inflation, changes in expected inflation, and unexpected inflation, and the majority of studies have confirmed negative relations between real stock returns and both expected and unexpected inflation, although the strength of the result of a negative relationship is varying.[2] Similarly, Schwert (1981) has also observed that the stock market reacted negatively to the announcement of unexpected inflation, but concluded that the most puzzling result of why aggregate stock returns were negatively related to the level of expected inflation remained a mystery.

The literature in this area has evolved with different layers of assumptions and specifications over the past three decades or so. Some studies such as LeRoy (1984), Stulz (1984), Danthine and Donaldson (1986), and Bakshi and Chen (1996) have sought to explain the anomalous results in the context of general equilibrium models, while others have analyzed the relation in the partial equilibrium framework. Some models such as Fama (1981), Fama and Gibbons (1982), Geske and Roll (1983), and Hasbrouck (1984) have attempted to establish a link between inflation and real activity, but many studies including LeRoy (1984), Stulz (1986), Danthine and Donaldson (1986), Boyle

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[2] Fama (1981) examined relations between real stock returns and both expected and unexpected inflation, but the evidence on the relations between real stock returns and unexpected inflation is less consistent. Stulz (1986) investigated the relationship between real stock returns and expected inflation, changes in expected inflation, and unexpected inflation and found that the strength of the relationship was weaker when the increase in expected inflation was caused by an increase in money growth rather than by a worsening of the investment opportunity set. Kaul (1987) found negative relations between real stock returns and expected, unexpected, and changes in expected inflation under counter-cyclical monetary policy regimes and positive relations under pro-cyclical monetary policy regimes. Lee (1992) observed that nominal stock returns and changes in expected inflation are weakly negatively correlated, and real stock returns and ex post inflation are mildly negatively correlated.

(1990), Marshall (1992), Bakshi and Chen (1996) have explained the stock return-inflation relation in the monetary asset pricing setting.

The return-inflation relation has recently provoked further controversies among economists as some studies have found even positive relations between real stock returns and expected inflation. These recent studies focus on the theoretical reconciliation of the seemingly contradictory results. The recent literature can be categorized as (1) the measure of expected inflation constructed from time series models versus survey data (Hasbrouck, 1984); (2) increases in expected inflation caused by an increase in money growth versus a worsening of the production opportunity set (Stulz, 1986); (3) counter-cyclical versus pro-cyclical monetary policy responses (Kaul, 1987, Park and Ratti, 2000); (4) inflation generated by monetary fluctuations versus inflation resulting from real economic fluctuations (Danthine and Donaldson, 1986, Marshall, 1992); (5) short-horizon versus long-horizon returns (Boudoukh and Richardson, 1993); and (6) cyclical versus non-cyclical movements in industry output (Boudoukh, Richardson, and Whitelaw, 1994). Although these recent studies have attempted to reconcile the conflicting results, the reconciliation is far from being reached.

The purpose of this study is to explore the return-inflation paradox. One important channel connecting real stock returns and expected inflation in this study is the behavior of velocity. The importance of velocity in explaining the behavior of asset prices has been noted by several authors such as Fama (1981), Friedman (1988), Boyle (1990), Marshall (1992), and Bakshi and Chen (1996). This empirical analysis explains a relation between asset returns and consumption velocity and establishes a link between the high volatility of asset returns and the high volatility of velocity. [3]

This approach is particularly important in that the traditional representative-agent model, given its reliance on consumption growth, has failed to reconcile the high variability of real asset returns with the low variability of consumption growth. Cochrane (1991, 1992, 1996), Kim (2003), and others have noted that production variables such as output growth, investment, and technology shocks are more volatile than consumption and proposed production-based (or investment-based) asset pricing models in which expected stock returns are significantly correlated to production. This study offers a new perspective on the relation between stock returns and expected inflation by linking explicitly the behavior of stock returns to the velocity of consumption which contains information about inflation, production, and monetary growth. This study aims to present a theoretical framework and empirical evidence for the apparent dependence of the relation between real returns and expected inflation on inflation volatility.

The analysis of the velocity-based asset prices is that relations between real stock returns and expected inflation are contingent on the degree of the volatility of inflation. More specifically, my analysis suggests that real stock returns tend to co-vary negatively with expected inflation in periods of stable inflation and to co-vary positively with

[3] Fama (1981) implicitly recognizes the importance of the role of velocity in explaining the relation between real stock returns and expected inflation. He notes that the spurious negative relations between inflation and expected real returns are induced by a somewhat unexpected characteristic of the money supply process during the post-1953 period, in particular, the fact that most of the variation in real money demanded in response to variation in real activity has been accommodated through offsetting variation in inflation rather than through nominal money growth. The money supply process, real money demanded, real activity, and inflation are all essential elements of velocity. Bakshi and Chen (1996) also note that nominal stock prices are negatively related to the contemporaneous velocity of money, while real stock prices are positively correlated with the velocity of money three quarters ahead (Friedman, 1988).

expected inflation in periods of volatile inflation. One possible explanation for the positive relation associated with volatile inflation is that increased inflation uncertainty may lead to an increased required risk premium for stocks. Thus, I argue that stocks are effective hedges against inflation during periods of volatile inflation and over a long period of time.

To a surprising extent, the results derived appear to be similar to those obtained from rational expectations macroeconomic models which have shown that the degree of inflation volatility matters in economic relations. For example, Lucas (1973) has demonstrated that the trade-off between inflation and unemployment holds more strongly in a stable price country than in a volatile price country. Put it in a different way, the Phillips curve is steeper in a volatile price regime than in a stable price regime. In a recent study, Akerlof, Dickens, and Perry (2000) have found that the relation between inflation and the natural rate of unemployment depends on the volatility of inflation, indicating that unemployment can be reduced below its natural level without inducing a rise in inflation when inflation is low and stable. This study is in parallel with these studies in that real stock returns are contingent on the volatility of inflation volatility.

The plan of this study is as follows: Section 2.2 reviews the recent theoretical and empirical literature. Section 2.3 shows an empirical analysis. Section 2.4 contains conclusions.

## 2.2. A REVIEW OF THE LITERATURE

Recent studies have attempted to reconcile with seeming contradictions found in the literature concerning the relation between stock returns and expected inflation. Day

(1984) has found that the actual or *ex post* real rate of return has an inverse relation to the *ex post* rate of inflation and maintained that the correlation between expected real returns and expected inflation is affected by investors' preferences and the form of the production process. When the production function exhibits stochastic constant returns to scale, the model explains the negative relation between expected real returns and expected inflation.

Hasbrouck (1984) has argued that with expected inflation measure constructed from time series models, the negative relationship is invariably significant and observed that the Livingston forecasts of economic activity are shown to be somewhat negatively related to expected inflation when monetary growth is held constant. When the proxy for real uncertainty is included in the return specification along with economic activity and expected inflation, the coefficient of expected inflation using the Livingston expected inflation proxy becomes positive (but not significant). While Hasbrouck has asserted that the negative relation of stock returns with expected inflation disappears when an *ex ante* measure of variability of real activity is taken into account, Stulz (1986) has shown that variability of real activity and expected inflation are negatively related. He has argued that the fall in real wealth associated with an increase in expected inflation decreases the expected real rate of return of the market portfolio. The expected real rate of return of the market portfolio falls less, for a given increase in expected inflation, when the increase in expected inflation is caused by an increase in money growth rather than by a worsening of the production investment opportunity set.

Danthine and Donaldson (1986) have argued that asset and commodity prices, rates of return, and rates of inflation are variables which are simultaneously determined and consequently are not independent of one another. Their general equilibrium model supports the evidence that real rates of return are negatively correlated with the rate of inflation. Their study also suggests that common stocks are not a good hedge against inflation of a non-monetary origin, but stocks will offer protection over the long run against purely monetary inflation. Bakshi and Chen (1996) have also offered a monetary asset pricing model in a general equilibrium setting in which the price level, inflation, asset prices, and the real and nominal interest rates are determined simultaneously and in relation to each other. They have shown that for many types of monetary economies, real stock returns are negatively correlated with expected or unexpected inflation, and are positively correlated with money growth.

Kaul (1987) has examined two types of monetary responses: counter-cyclical and pro-cyclical. He has proposed that counter-cyclical monetary responses lead to negative relations between real stock returns and expected inflation (or unexpected inflation or changes in expected inflation). Kaul has further hypothesized that money demand effects combined with pro-cyclical monetary responses would result in insignificant or even positive relations between stock returns and inflation, as was confirmed in the United States during the 1930s. Marshall (1992) has investigated the relation between real stock returns and expected inflation in the dynamic context of a monetary inter-temporal asset pricing model. He has found that the relation between real stock returns and expected inflation is strongly negative when inflation is caused by real economic fluctuations and is ambiguous in sign and small in magnitude when inflation is caused by monetary fluctuations.

Boudoukh and Richardson (1993) have employed a long-horizon representation of the Fisher equation to test the hypothesis that long-horizon nominal returns are positively related to long-term inflation, and short-horizon returns are negatively related to short-term inflation. For U.S. stock returns during the period 1802 –1990, they have found that the coefficient of five-year stock returns on the contemporaneous five-year inflation rate was significantly positive and concluded that long-term nominal stock returns and inflation tend to move together.[4] Boudoukh, Richardson, and Whitelaw (1994) have explored the stock return-inflation relation in a Fisherian context, and found that the sign and magnitude of the covariance between nominal stock returns and expected inflation may be affected by cyclical movements in industry output. They have found that returns on stocks of cyclical industries tend to co-vary negatively with expected inflation while the reverse holds for non-cyclical industries.

Gultekin (1983) has ascribed the lack of the positive relation between stock returns and inflation to the errors-in-variables problem. If observed stock returns and observed inflation are related to their ex ante counterparts with error terms, then the regression of the actual rate of return against the actual rate of inflation could cause the estimate of the coefficient to be biased. If the covariance between the two errors is negative, that is, the market reacts negatively to the unexpected inflation, the bias could result in a negative relation.

[4] Boudoukh and Richardson have obtained similar results using ex ante inflation.

Some recent evidence on international stock markets is further puzzling. Rapach (2002) has shown that estimates of the long-run real stock price response to a permanent inflation shock are zero or positive for 16 industrialized countries. Chatrath, Ramchander and Song (1997) for India, Najand and Noronha (1998) for Japan, Zhao (1999) for China and Crosby (2001) for Australia have confirmed the negative relationship between inflation and stock returns. Adrangi, Chatrath and Raffiee (1999) have studied the effects of macroeconomic variables on stock returns for Korea and Mexico and shown that a negative relationship between real stock returns and unexpected inflation exists. Choudhry (2001) has found a positive relationship between current stock returns and current inflation in four high-inflation countries (Argentina, Chile, Mexico and Venezuela). Spyrou (2004) has reported a positive relationship between the two variables for 10 emerging stock markets (ESM), namely Chile, Mexico, Brazil, Argentina, Thailand, South Korea, Malaysia, Hong Kong, Philippines and Turkey during the period of the 1990s. Omran and Pointon (2001) have documented a negative relationship for Egypt. Apergis and Eleftheriou (2002) have obtained a negative relationship for Greece, whereas Spyrou (2001) has observed a negative relationship between inflation and stock returns only for the period until 1995 while for the remaining period until 2000 he has found no statistically significant relationship. Finally Hondroyiannis and Papapetrou (2006) show that real stock returns are not related to expected and unexpected inflation.

Although recent studies, especially those by Kaul (1987), Marshall (1992), Boudoukh and Richardson (1993), Boudoukh, Richardson, and Whitelaw (1994) have substantially contributed to unraveling the mystery concerning stock return-inflation relations, it cannot be denied that these models still remain unsatisfactory in providing consistent explanations for the puzzle. Most of all, the recent attempts to reconcile the empirical conflicts are based on strong assumptions about the behavior of economic agents. For example, the Boudoukh and Richardson, and Boudoukh, Richardson, and Whitelaw models are constructed based on the Fisher equation, not on the optimizing behavior of market participants. As Marshall (1992) notes, the Fisher relation does not generally address the implications of dynamic economic equilibria when the role of money is explicitly taken into consideration. Dokko and Edelstein (1987) warn that the Fisher relation should be viewed as a reduced-form equation derived from a set of unknown behavioral equations.

The Marshall model is an important improvement in that it investigates correlations between real asset returns and inflation in the dynamic monetary equilibrium context. However, the model itself is not clear about the distinction between inflation induced by real economic shocks and inflation induced by monetary shocks. Furthermore, although the Marshall model recognizes some association between the variability of consumption velocity and the variability of real asset returns, it does not explicitly incorporate this point into the model.

## 2.3. METHODOLOGY

### 2.3.1. *The Estimation Equation*

In order to examine the relationship between the rate of inflation and the return on common stocks empirically, I regress real stock returns on the rate of expected inflation, the growth rate of expected real GDP, and the growth rate of the expected money supply. In estimating the relationship, one delicate estimation issue is how one can obtain an estimate of expected inflation. In the literature, several different methods have been proposed and used. First, one can use the contemporary inflation rate as a proxy for expected inflation (Gultekin (1983)). However, when observed stock returns and observed inflation are related to their ex ante counterparts with error terms such as

(2.1) $$r_t = E(r_t \mid \Omega_{t-1}) + u_t$$

(2.2) $$\pi_t = E(\pi_t \mid \Omega_{t-1}) + v_t \,,$$

the slope coefficient in the regression of $r_t$ on the expected inflation rate is biased (Nelson (1976), Gultekin (1983)).

Many researchers use short-term interest rates such as the Treasury Bill rate as proxies for expected inflation assuming a constant real interest rate. (Jaffe and Mandelker (1976), Fama and Schwert (1977), Schwert (1981), Gultekin (1983)). Some authors decompose inflation into expected and unexpected components by ARIMA models, and use inflation forecasts from ARIMA as estimates of expected inflation (Gultekin (1983)). However, Fama (1977) and Schwert (1981), Hillion and Solnik (1982), Solnik (1983) and others have shown that the ARIMA representation of expected inflation did not significantly outperform short-term nominal interest rates as a predictor of expected inflation. Finally one can use lagged inflation rates as an estimate of expected inflation. For instance, Jaffe and Mandelker (1976) related the return on stocks to past three lagged inflation rates. Nelson (1976) included the four lags of inflation rates to examine the relationship between real stock returns and inflation. Gultekin

(1983) has used the four lagged inflation rates to represent expected inflation. Ball (2000) also has used the four lagged inflation rates and the output to represent expected inflation and the output. I include four lagged inflation rates and the growth rate of real GDP to represent expected inflation and the expected growth rate of real GDP. People can adjust their expectations of inflation and the growth rate of real GDP on the basis of past inflation rates. Thus, it is appropriate to include lagged inflation rates and the growth rate of real GDP. As Gultekin indicates, regressing stock returns on the past inflation rates should eliminate the errors-in-variables bias due to the negative covariance, since past inflation rates contain no new information for the market.[5] I further note that $r_t$ (the real rate of return) is less than – 1.0 in some sample observations. Since $\ln R_t$ is not defined for the negative value of $r_t$ that is smaller than – 1.0, I approximate $\ln R_t = \ln(1+r_t)$ by $r_t$ (the real rate of return).

I employ a GARCH (1,1)-M model. The mean and variance equations of the GARCH (1,1)-M model are specified as follows:

(2.3) $$r_t = \beta_0 + \sum_{j=1}^{4} \beta_j INF_{t-j} + \sum_{j=1}^{4} \beta_{j+4} RGDP_{t-j} + \sum_{j=1}^{4} \beta_{j+8} M_{t-j} + \gamma \log(\sigma_t) + u_t$$

(2.4) $$\sigma_t = \sqrt{Var(u_t | \Omega_{t-1})}$$

(2.5) $$u_t | \Omega_{t-1} \sim N(0, \sigma_t^2)$$

(2.6) $$\sigma_t^2 = \alpha_0 + \alpha_1 \varepsilon_{t-1}^2 + \alpha_2 \sigma_{t-1}^2 + \alpha_3 \sigma(INF_{t-1})$$

[5] See Nelson (1976) and Gultekin (1983).

where INF is the rate of inflation, RGDP is the growth rate of real GDP, and M is the growth rate of money supply. In Equation (2.4), $\sigma_t$ denotes the conditional standard deviation of the error term, that is, $\sigma_t$ is the one-period ahead forecast standard deviation based on past information, $\Omega_{t-1}$. The logarithm of the standard deviation represents the effect of volatility on stock returns that is less than proportional in the mean.[6] Equation (2.6) is the conditional-variance equation. The conditional-variance equation consists of three parts: $\alpha_1$ measures the effect of squared innovations on volatility from the previous period (the ARCH term), and the coefficient $\alpha_2$ denotes the effect of the forecast variance of the last period (the GARCH term). The sum of $\alpha_1$ and $\alpha_2$ measures the extent of the persistence of volatility. The last term is the conditional volatility of inflation $\sigma(INF_t)$. Incorporating the effect of inflation volatility on the stock return distribution is important because inflation volatility influences the volatility of real stock returns in my model. Therefore, $\alpha_3$ captures the effect of unexpected conditional inflation volatility on the conditional volatility of real stock returns.[7]

### 2.3.2. *The Description of the Data*

The data used in this study have been taken from the International Monetary Fund's online *International Financial Statistics* (IFS). I use quarterly data to estimate the GARCH(1,1)-M model. The countries that are included in my sample are those that have the following data series: the consumer price index (CPI), nominal share price index,

[6] Empirically, the logarithm of the conditional variance is better than the standard deviation but this method cannot affect the significance of coefficients.
[7] See Elyasiani and Mansur (1998), Ryan and Worthington (2004) for using more other volatility variables in conditional variance equation.

GDP, GDP deflator, and M2. There are 33 countries that meet this selection criterion. I have eliminated countries that have fewer than 70 observations.[8] Thus, I have finally obtained 16 countries. The starting dates and ending dates differ for countries in my data set.

Real stock returns are obtained from the nominal share price index series (IFS series 62..ZF..) deflated by the consumer price index. The inflation rate is calculated by taking the first difference of the natural logarithm of the consumer price index series (IFS series 64..ZF..). The growth of real GDP is calculated from the GDP series (IFS series 99B..ZF.. and 99B..CZF..) deflated by the GDP deflator (IFS series 99..BIR..ZF..). The M2 series (IFS series 35L..ZF.. and 59MB..ZF..) is the sum of currency outside banks, demand deposits other than those of central government, time and savings deposits, and foreign currency deposits of residents other than the central government.[9] The annualized values of the series are used.

### 2.3.3. *Empirical Results*

Table 2.1 reports basic statistics for the variables used in this study. The table shows that the average annual rate of real stock returns ranges from -1.23% to 24.04%, and the average annual inflation rate ranges from 2.6% to 51.2%. The standard deviation of real

[8] I eliminate some countries since quarters of data are not enough to estimate the GARCH specification. Germany, Portugal, Sweden, Switzerland in OECD countries and Argentina, Brazil, Chile, and Colombia in South America and most of developing countries have below 70 quarters.

[9] For monthly data, there are 37 countries that can be used. The starting dates and ending dates are differ for these countries in my data set. Real stock returns are obtained from the MCSI index series (datastream) deflated by the consumer price index. The inflation rate is calculated by taking the first difference of the natural logarithm of the consumer price index series (IFS series 64..ZF..). The growth of real GDP is calculated from Industrial Production series (IFS series 66..ZF.. , 66..BZF.. , 66..CZF..)[9]. The M2 series from IFS and datastream (IFS series 35L..ZF.. , 59MB..ZF.. , datastream). The annualized values of the series are used.

stock returns is relatively large in the Philippines (83.63%), Mexico (73.86%), Peru (76.46%), Israel (50.29%), and South Korea (50.04%) and the smallest in the United States (24.52%). On the other hand, the standard deviation of the inflation rate is pronouncedly large in Peru (129.76%), Israel (40.99%), and Mexico (25.74%), whereas other countries have a standard deviation of the inflation rate that is less than 10%. (The standard deviation of the Philippines (10.25%) is close to 10%.) Thus, Peru, Israel, and Mexico can be classified as volatile-price countries, whereas the rest of the countries in my sample are characterized as stable-price countries. Because my data set includes countries that exhibited highly volatile price movements, my sample provides a good laboratory for testing relationships.

If I use the Euler equation for the optimal consumption path and consumption velocity of money, then stock return can be expressed as a function of the expected values of inflation rate, the growth rate of real GDP growth, and the growth of money supply. Table 2.2 reports the results of simple OLS estimation when realized inflation rates are used for proxies of expected inflation rates. The coefficients are significantly negative except France, Italy, Mexico, and U.K. However, these coefficients are negatively biased since the observed inflation rates are related with ex-ante counterparts with error terms. Therefore, there exists the lack of positive relation between stock returns and inflation. In addition, this simple OLS result do not contain the lagged terms for expected values and inflation volatility.

TABLE 2.1

BASIC STATISTICS FOR VARIABLES

| COUNTRY | Real Stock Returns | | Inflation Rates | | Real GDP Growth Rates | | M2 Growth Rates | | Sample Period |
|---|---|---|---|---|---|---|---|---|---|
| | Mean | Std. Dev. | Mean | Std. Dev. | Mean | Std. Dev. | Mean | Std. Dev. | |
| AUSTRALIA | 2.02 | 33.50 | 5.17 | 4.44 | 3.54 | 4.63 | 10.17 | 8.93 | 59.Q4-07.Q1 |
| CANADA | 2.26 | 28.54 | 4.04 | 3.39 | 3.57 | 4.05 | 10.54 | 23.10 | 57.Q2-07.Q1 |
| FRANCE | 8.28 | 35.38 | 4.96 | 4.18 | 2.04 | 1.97 | 6.01 | 13.60 | 77.Q3-98.Q4 |
| ISRAEL | 4.58 | 50.29 | 30.97 | 40.99 | 4.69 | 113.24 | 37.65 | 71.67 | 71.Q2-06.Q2 |
| ITALY | 7.81 | 48.51 | 7.07 | 4.90 | 1.92 | 3.05 | 7.22 | 21.67 | 80.Q2-98.Q4 |
| JAPAN | 3.49 | 32.03 | 3.47 | 5.06 | 4.91 | 39.17 | 10.95 | 17.86 | 57.Q2-07.Q1 |
| KOREA | 1.95 | 50.04 | 6.13 | 6.75 | 7.19 | 56.78 | 15.77 | 13.03 | 78.Q2-06.Q4 |
| MEXICO | 16.04 | 73.86 | 22.86 | 25.74 | 2.54 | 18.37 | 26.24 | 44.03 | 84.Q2-07.Q1 |
| NETHERLAND | 8.04 | 31.58 | 2.92 | 2.46 | 2.29 | 3.47 | 8.63 | 24.99 | 77.Q2-97.Q4 |
| NEW ZEALAND | 3.67 | 49.08 | 2.65 | 2.38 | 2.69 | 4.29 | 10.84 | 32.78 | 87.Q2-06.Q4 |
| NORWAY | 3.40 | 43.07 | 5.78 | 4.07 | 3.67 | 20.80 | 9.66 | 13.53 | 66.Q3-01.Q2 |
| PERU | 24.04 | 76.46 | 51.25 | 129.76 | 4.11 | 28.04 | 58.37 | 121.88 | 89.Q2-07.Q1 |
| PHILIPPINES | 4.39 | 83.63 | 9.00 | 10.24 | 2.71 | 39.25 | 15.66 | 38.31 | 81.Q2-07.Q1 |
| SPAIN | -1.23 | 42.95 | 9.27 | 6.04 | 2.90 | 3.63 | 11.99 | 16.35 | 70.Q2-98.Q4 |
| U.K. | 3.43 | 33.59 | 6.34 | 6.02 | 2.44 | 4.30 | 10.91 | 23.38 | 58.Q2-99.Q1 |
| U.S. | 2.54[10] | 24.52 | 3.98 | 3.04 | 3.21 | 3.60 | 7.39 | 6.25 | 57.Q2-07.Q1 |

[10] For nominal stock return, it is almost same as Gultekin (1983)'s stock return as 0.58% in monthly data. Annualized nominal value using quarterly data is 6.5%. In the appendix A, I report the sources of stock returns from *IFS*.

TABLE 2.2

OLS ESTIMATION WITH PROXIES FOR VARIABLES

| | $\beta_0$ | $\beta_1$ | $\beta_2$ | $\beta_3$ |
|---|---|---|---|---|
| AUSTRALIA | 11.310** | -2.078** | 0.068 | 0.122 |
| CANADA | 7.973** | -1.663** | 0.397 | -0.038 |
| FRANCE | 12.002 | -0.060 | -0.778 | -0.305 |
| ISRAEL | 10.970** | -0.327** | -0.035 | 0.104** |
| ITALY | 0.082 | 1.263 | 1.299 | -0.513** |
| JAPAN | 7.421** | -1.998** | -0.214** | 0.370** |
| KOREA | 24.113** | -1.687** | 0.089 | -0.789** |
| MEXICO | 15.526 | -0.218 | -0.200 | 0.229 |
| NETHERLAND | 8.025 | -1.364 | 1.054 | 0.185** |
| NEW ZEALAND | 22.792* | -5.846* | -1.450 | 0.029 |
| NORWAY | 23.482** | -1.903** | -0.013 | -0.933** |
| PERU | 10.357 | -0.933** | 0.223 | 1.038** |
| PHILIPPINES | 26.558* | -1.894** | -0.303 | -0.274 |
| SPAIN | 30.503** | -2.790** | 0.268 | -0.554** |
| UK | 3.646 | -0.606 | 0.715 | 0.172** |
| US | 11.234** | -2.746** | 0.854* | -0.065 |

Note: (1) $r_t = \beta_0 + \beta_1 INF_t + \beta_2 RGDP_t + \beta_3 M_t + \varepsilon_t$
(2) Standard Error is corrected for heteroskedasticity
(3) ** and * are significant at 5% and 10% level, respectively.

Table 2.3 presents the results of the GARCH(1,1)-M estimation. My primary concern lies in the relationship between real stock returns and expected inflation. I am particularly concerned with whether the rate of real stock returns is negatively related to the inflation rate in a stable-price country and is positively related to the inflation rate in a volatile-price country. My sample contains three volatile-price countries (Mexico, Peru, and Israel) and 13 stable-price countries.[11]

Interestingly enough, the relationship between inflation (measured by the sum of the coefficients of the lagged inflation terms) and real stock returns has turned out to be negative in all stable-price countries. Furthermore, my Wald test shows that the coefficients taken together are significant at the conventional level of significance. On the other hand, Mexico and Peru have a positive relationship between inflation and real stock returns. Israel is only the exception in my empirical analysis: The relationship has been negative in Israel for the full sample, but the coefficients taken together are insignificant. Thus, I can conclude that my empirical analysis unravels some puzzling observations concerning the relationship between real stock returns and expected inflation in a consistent and unambiguous manner.

My findings can be compared with those of Gultekin (1983) who has investigated the relationship between the rate of return on common stocks and the inflation rate (four lags) for 14 advanced countries from 1947 to 1979 using the same data source as ours (*International Financial Statistics*). He has found that with the exception of the United Kingdom, the sum of the coefficients of the four lagged terms of the inflation rate was negative. His findings were in the same context as ours, but his model was not able to distinguish between stable-price regimes and volatile-price regimes in identifying the relationship between stock returns and expected inflation.

[11] In this monthly data, the standard deviation of the inflation rate in Argentina, Brazil, Hungary, Poland, Taiwan, and Turkey is greater than 10%. We also get the positive relationship except Taiwan.

# TABLE 2.3

## GARCH(1,1)-M ESTIMATION WITH PROXIES FOR EXPECTED INFLATION

| | $\beta_0$ | SUM of INF | SUM of RGDP | SUM of M | $\gamma$ | $\alpha_0$ | $\alpha_1$ | $\alpha_2$ | $\alpha_3$ | Q(12) | $Q^2(12)$ |
|---|---|---|---|---|---|---|---|---|---|---|---|
| AUSTRALIA | 2.21 | -2.77** | -0.69** | 1.97** | -0.18** | 23.65 | 0.63** | 0.42** | 13.17 | 42.06** | 14.54 |
| CANADA | 185.66 | -0.11 | -2.78** | -0.16** | -26.38 | 262.75 | 0.01 | 0.22 | 79.88 | 17.18 | 7.30 |
| FRANCE | -48.51 | -2.95** | 2.02 | 0.50** | 9.31 | 1184.05** | 0.31** | -0.95** | 40454 | 6.53 | 7.38 |
| ISRAEL | -504.43 | -0.65 | -0.13 | 0.26** | 66.95 | 2834.82 | 0.01 | -0.39 | 14.45 | 13.75 | 2.38 |
| ITALY | -59.14 | -3.61** | -1.66* | 4.94** | 8.18 | 777.33** | 0.04 | 0.64** | -84.36** | 20.06* | 12.17 |
| JAPAN | -4.26 | -1.26** | 0.15* | 0.09 | 1.65 | 826.97** | 0.01 | 0.29** | -44.96** | 29.96** | 6.96 |
| KOREA | -315.93* | -0.15 | 1.37 | -2.22** | 45.28* | 717.75** | 0.26** | 0.39** | 4.70 | 19.39* | 8.56 |
| MEXICO | 1300.36* | 2.14* | -1.29** | -0.09 | -162.56* | 1214.79** | 0.13 | 0.45** | 18.46 | 11.30 | 16.25 |
| NETHERLAND | -11.59 | -0.53** | 0.86 | 0.62** | 2.20 | 583.99* | -0.09** | 0.42 | -33.84 | 19.52* | 6.80 |
| NEW ZEALAND | 1225.57** | -14.42** | 18.83 | -3.50** | -142.46** | 1327.20** | 0.44** | 0.18** | 216.75** | 65.32** | 7.75 |
| NORWAY | 123.12** | -1.85 | 0.29** | -0.74** | -13.88** | 1259.71 | -0.09** | 0.58** | -99.06 | 21.03* | 15.26 |
| PERU | -626.13 | 1.56** | 0.50 | 0.01** | 78.70 | 3000.57** | 0.17 | -0.27 | 26.63 | 9.67 | 9.95 |
| PHILIPPINES | 510.63** | -1.79** | -0.91** | -1.50** | -53.68** | 9582.63** | 0.17** | -1.04** | -59.90** | 16.82 | 11.93 |
| SPAIN | 9.74 | -2.14** | 3.34* | -0.96** | 1.27 | 719.86** | 0.27** | 0.48** | -51.36** | 18.56 | 6.04 |
| UK | -25.97 | -0.97 | -1.25 | -0.03** | 5.99 | 653.89** | 0.39** | 0.38* | 53.89 | 13.01 | 6.89 |
| US | 64.13* | -2.26** | -2.35** | 0.55** | -7.95 | 187.03** | 0.24 | 0.04 | 59.00* | 15.02 | 15.18 |

Note: (1) $r_t = \beta_0 + \sum_{j=1}^{4} \beta_j INF_{t-j} + \sum_{j=1}^{4} \beta_{j+4} RGDP_{t-j} + \sum_{j=1}^{4} \beta_{j+8} M_{t-j} + \gamma \log(\sigma_t) + u_t$ and $\sigma_t^2 = \alpha_0 + \alpha_1 \varepsilon_{t-1}^2 + \alpha_2 \sigma_{t-1}^2 + \alpha_3 \sigma(INF_{t-1})$

(2) Bollerslev and Wooldridge's robust variance estimator is employed.

(3) Q-test and is the test for serial correlation and $Q^2$-test is the test for dependency in squared residuals.

(4) ** and * are significant at 5% and 10% level, respectively.

I should expect the contingency of real stock returns on inflation volatility to hold in a time series over time. Fama (1981) found a negative relation between real U.S. stock returns and inflation from 1954 to 1976, whereas Kaul (1987) found a positive relation between U.S. real stock returns and expected inflation from 1926 to 1940. These findings can be reconciled within my empirical results. The standard deviation of the U.S. annual inflation rate during the 1954 -1976 period was 2.43 percent while the standard deviation of the U.S. annual inflation rate during the 1926 -1940 period was much higher at 4.98 percent.[12]

More recently, McCown and Fitzgerald (2006) have examined the simple correlation between real stock returns and expected inflation for industrialized countries during the pre-World War II period. Table 2.4 shows McCown-Fitzgerald's results. McCown-Fitzgerald's evidence showed a stylized pattern in the relation between real stock returns and expected inflation: when the volatility of inflation is relatively low (that is, the standard deviation of the inflation rate is roughly less than 10%), the return-inflation relation is negative; when the volatility of inflation is relatively high (that is, the standard deviation of the inflation rate is greater than 10%), the return-inflation relation is positive. The only exception occurred in France during the 1857 -1913 period, but the correlation was insignificant at the 5 percent level. There is a striking resemblance between their stylized pattern and my finding.

[12] The standard deviation of the inflation rate for each period was calculated on the basis of the GDP deflator.

TABLE 2.4

CORRELATION BETWEEN REAL STOCK RETURNS AND EXPECTED INFLATION

| Country | Time Period | Standard Deviation | Correlation |
|---|---|---|---|
| Denmark | 1923 – 1939 | 6.13% | -0.060 |
| France | 1857 – 1913<br>1919 – 1937 | 4.27%<br>16.10% | 0.237<br>0.004 |
| Germany | 1871 – 1913 | 8.57% | -0.389* |
| UK. | 1868 – 1913<br>1919 – 1939 | 4.68%<br>14.96% | -0.539*<br>0.438* |
| US | 1802 – 1939<br>1802 – 1913<br>1919 – 1939 | 8.83%<br>9.13%<br>5.61% | - 0.076<br>- 0.089<br>- 0.091 |

Note: *indicates significance at the 5 percent level.

## 2.4. CONCLUDING REMARKS

Since negative relations between real (or nominal) stock returns and expected inflation were observed in the post-war U.S. data, there has been a proliferation of studies which have attempted to resolve the return-inflation puzzle. Although the existing studies have explained some important aspects of the anomalous relations, the stock return-inflation puzzle still remains unresolved. The purpose of this study is to provide theoretical foundations and empirical evidence for the seemingly paradoxical results within an inter-temporal portfolio-choice framework. One novel feature of this study is that it converts the consumption-based capital asset pricing model into a relationship between real stock returns and consumption velocity and links the volatility of asset returns to the volatility of velocity. The traditional consumption-based asset pricing model fares poorly in explaining asset price movements because it is unable to reconcile the high variability of real asset returns with the low variability of consumption growth.

In my empirical analysis, expected real stock returns are determined by market fundamentals such as the expected values of inflation, money growth, real output growth, and monetary and real shocks. Given monetary and real shocks, the relation between real stock returns and expected inflation can be of either sign depending on the degree of inflation volatility. My empirical analysis suggests that real stock returns are negatively related to expected inflation in periods of low volatility of inflation and positively related to expected inflation in periods of high volatility of inflation. My empirical analysis provides unambiguous implications for the relation between real stock returns and expected inflation.

In order to test for relationships, I have employed an GARCH(1,1)-M model and conducted an empirical investigation using quarterly data for 16 countries. The data set includes 13 stable-price countries and three volatile-price countries. My empirical results have confirmed that the relationship between real stock returns and expected inflation was negative in all stable-price countries and positive in two volatile-price countries. The only exception was Israel where the volatility of inflation measured by the standard deviation of the inflation rate was relatively high during the sample period, but a negative relation between real stock returns and expected inflation was found, albeit the coefficient was not significant.

This result is in line with McCown and Fitzgerald (2006)'s paper. My conclusion can be compared with that of Akerlof, Dickens, and Perry who have found that the unemployment can be reduced below the natural level without stimulating higher inflation when the inflation rate is between zero and 4 percent, but the unemployment rate eventually approaches the conventional natural level when the inflation rate is higher than 4 percent. The most important conclusion drawn from this study is that the relation between real stock returns and expected inflation is significantly affected by the degree of inflation volatility.

# CHAPTER III

# TESTING FOR BUBBLES IN ASSET PRICES

## 3.1. INTRODUCTION

Broadly speaking, bubbles include both rational and irrational bubbles. Irrational bubbles mean that market participants do not recognize the existence of bubbles. However, rational bubbles describe the fact that that market participants recognize the existence of bubbles and expect increasing prices regardless of the internal value of firms. The expected value of rational bubbles of stock prices either would increases or decreases. If a positive rational bubble exists, market participants might expect it eventually to dominate stock prices, which would then bear little relation to market fundamentals. On the other hand, given free disposal, a negative rational bubble can not exist because market participants can not rationally expect stock prices to decrease without bound, and hence, to become negative.

Although the stock market seems to be overheated, people do not recognize if bubbles exist or not. That is to say, stock prices can be adjusted into appropriate levels so far as the prices are not developed into explosive bubbles. Therefore, Bubbles will be detected only after the market is collapsed. The traditional studies of bubbles focus mainly on the rational aspect of bubbles.

In general, economists have understood that it is difficult to test what her bubbles exist or not. For example, Evans (1991) has maintained that it seems to be difficult to test the case where the stock prices are high enough but there is no possibility of the burst. Because of this problem, economists traditionally have used indirect methods to test bubbles. Shiller (1981) has used a variance bounds test and interprets excessive volatility as bubbles. The deviation of stock prices, which the present value (PV) model forecasts, may be the result of waves of pessimistic or optimistic market psychology.

There has also been the resurgence of interest in bubbles in stock prices primarily through the works of Hamilton and Whiteman (1985), Diba and Grossman (1988a, b, c), Evans (1991), Froot and Obstfeld (1991), and others. The bubble models elaborated by these authors have represented a significant departure from the conventional paradigm in that they reinterpret rational bubbles in terms of market fundamentals. These bubbles may be termed intrinsic bubbles, as opposed to extraneous bubbles in the traditional view. It has been widely believed that bubbles do not easily lend themselves to direct testing. The attractive feature of the intrinsic bubble specification may be found in its ability to derive testable implications for bubbles by investigating the stationarity properties of stock prices and dividends or by parameterizing a specific bubble relationship as a function of market fundamentals. However, the existing approaches to intrinsic bubbles still remain unsatisfactory. As Evans has pointed out, Diba and Grossman's tests for stationarity have been unable to detect an important class of rational bubbles. The ability of Froot and Obstfeld (1991)'s parametric test to discover bubbles is also doubtful. The essence of their test is that if the price-dividend ratio is significantly (and nonlinearly) related to current dividends, the hypothesis of no bubbles is rejected. However, all the recent tests for bubbles have not been direct measures of bubbles.

In this study, I formulate an information error model which allows one to derive the measure of non-fundamentals in stock prices in a straightforward manner. This study provides a new method by specifying bubble measures as the Weibull distribution, and it is the first attempt to apply the Weibull distribution to the test of rational bubbles. There is not only a parallel between the burst of speculative bubbles and a material's burning out, but there is also a good reason to believe that measured bubbles can be appropriately modeled as the Weibull specification. The plan of this study is as follows: Section 3.2 reviews the literature, section 3.3 develops a theory of bubbles, section 3.4 gives empirical results, and section 3.5 contains conclusions.

## 3.2. A REVIEW OF THE LITERATURE

Shiller (1981a) has been the driving force in challenging the interpretation that stock price movements reflect the efficient discounting of new information on market fundamentals. This study, using the present value (PV) model of stock prices, has demonstrated that stock prices are too volatile to be consistent with the present value of rationally expected future dividends discounted by a constant real interest rate. In particular, Shiller's variance bounds test has established that the variance of the market price of a stock should not be greater than that of the present discounted value of future cash flows. Shiller has shown that the violation of this inequality is overwhelming for U.S. data. The violation of the variance bounds are interpreted as a rejection of the efficient markets hypothesis, although its advocates disagree. Therefore, the deviation of stock prices, which the present value (PV) model forecasts, begins with waves of pessimistic or optimistic market psychology. As mentioned above, this is called extraneous bubbles.

Hamilton and Whiteman (1985) have dealt with bubbles which can be dependent on market fundamental factors for the first time. They have discussed bubbles from the view of a stationarity property between stock prices and market fundamental factors since it is impossible to test the view that the bubbles are caused by self-fulfilling expectations from purely extraneous factors. Therefore, the existence of bubbles is the output of rational actions on market fundamentals by market participants. It has been assumed that bubbles exist if the *d*-th differenced stock prices are stationary and the *d*-th differenced fundamental factors are nonstationary. That is, there are no bubbles if fundamental factors such as dividends are more explosive than stock prices.

Diba and Grossman (1988a) have postulated that stock prices are cointegrated with market fundamentals in a nonlinear fashion if there is no bubble premium in stock prices (nonlinear cointegration). However, they also have suggested that the opposite is not true since non-cointegrated stock prices and dividends can occur from the non-stationarity of a variable that market participants either observe or conduct but that the researcher does not observe.

Campbell and Shiller (1987) have tested cointegration between stock prices and dividends using the Standard and Poor’s 500 Composite Price Index from 1871 to 1976. Campbell and Shiller (1988) have suggested that the dividend-price ratio can be explained by some fundamentals even though stock prices are not explained by macroeconomic variables. They have also found that the dividend-price ratio (D/P) and the long-run return for stock prices have a significant relationship with the growth of dividends. Moreover, they have also found that short run interest rates do not cause the dividend-price ratio (D/P) and dividends, but the growth of dividends causes the dividend-price ratio (D/P).

Evans (1991) has suggested that the sum of stock prices and dividends (P+D) and dividends are not cointegrated, and it cannot be conclude that there are no bubbles using the Standard and Poor's 500 Composite Price Index from 1871 to 1980. Therefore, Diba and Grossman's tests for stationarity are unable to detect an important class of rational bubbles when using test for unit roots. That is, time series can be stationary even if there are bursting bubbles.

Froot and Obstfeld (1991) also have made an assertion with Hamilton and Whiteman that the existence of bubbles is the output of rational actions about market fundamentals of market participants, so bubbles are not extraneous. On the contrary the self-fulfilling expectation of intrinsic bubbles is based on market fundamental factors and they induce the functional relationship between bubbles and market fundamental factors. That is, if the price-dividend ratio (P/D) has a significant relationship with current dividends, bubbles exist. They have found that unexplained parts from the present value model of stock prices have a high positive relationship with dividends using the Standard and Poor's 500 Composite Price Index from 1900 to 1988.

More recently, Koustas and Serletis (2005) have employed fractional integration techniques, which are combined with volatility modeling, to investigate the persistence of the logarithm of the dividend yield for Standard and Poor's 500 Price Index. They have found that a fractionally integrated dividend yield is not consistent with rational bubbles. Cunado et al. (2007) have also used fractional integration techniques using the Standard and Poor's 500 Price Index over the period 1871m1-2004m6. They have found that rational bubbles exist for the whole period.

## 3.3. THE THEORETICAL MODEL

The present value (PV) model illustrates that stock prices are equal to the present value of future cash flows plus fundamental factors such as dividends discounted by a constant real interest rate. The traditional view interprets bubbles as the variation of stock prices deviating from the orbit which the present value model forecasts. Generally, the present value model has a form as follows:

(3.1) $$P_t = \delta E_t(P_{t+1} + D_{t+1})$$

where $P_t$ is stock prices at time $t$ and $D_{t+1}$ is dividends between $t$ and $t$+1, and $\delta$ is a time discount factor, which is given as 1 / (1+$r$) with constant real interest rate $r$.

The solution is

(3.2) $$P_t = \sum_{k=1}^{\infty} \delta^k E_t(D_{t+k}) + \lim_{T\to\infty} E_t \delta^T P_{t+T}$$

The general solution to equation (3.2) is the sum of the market fundamentals component in equation (3.3) and a rational-bubbles component in equation (3.4).

(3.3) $$\tilde{P}_t = \sum_{k=1}^{\infty} \delta^k E_t(D_{t+k})$$

(3.4) $$B_t = \lim_{T\to\infty} E_t \delta^T P_{t+T}$$

Therefore, current stock prices are divided into two parts, (3.5) $$P_t = \sum_{k=1}^{\infty} \delta^k E_t(D_{t+k}) + B_t$$

where a random variable $B_t$ has the following characteristic:

(3.6) $$B_t = \delta E_t B_{t+1}$$

Diba and Grossman (1988c) have shown that the expected value of a rational-bubble component of stock prices either would increases or decreases. If a positive rational bubble exists, stock holders might expect it eventually to dominate stock prices, which would then bear little relation to market fundamentals. On the other hand, given free disposal, a negative rational-bubble component can not exist because stock holders can not rationally expect stock prices to decrease without bound, and hence, to become negative.

Hamilton (1986) have made the model as follows:

(3.7) $$P_t = \delta(D_{t+1} + E_t P_{t+1} + \pi_t)$$

where $\pi_t$ is the random variable that market participants observe but that researchers do not observe. Moreover, $\pi_t$ represents a regression disturbance term, corresponding to omitted variables such as time-varying real interest rates, risk-premia, and changes in tax laws. They have assumed that $\pi_t$ is stationary, that is, I(0).

Diba and Grossman (1988a) have examined stock price volatility in the framework of self-fulfilling expectations of market fundamentals elaborated by Hamilton and Whiteman (1985) and Hamilton (1986). To formulate self-fulfilling expectations, Diba and Grossmann introduce a variable that is observed by economic agents but not the econometrician. Thus, in the Diba and Grossman model, stock prices depend on unobservable variables as well as dividends in the following manner:

(3.8) $$P_t = \delta E_t(P_{t+1} + \alpha D_{t+1} + \pi_{t+1})$$

where $\pi_t$ is the random variable that market participants observe but that the researcher does not observe. $\alpha$ is a positive constant and the ratio of expected dividends relative to expected capital gains. Let dividends and stock prices have unit roots and 1-st differencing dividends, 1-st differencing stock prices. That is to say, stock prices and dividends are I(1). There are no bubbles if stock prices and dividends are I(1) and stock prices and dividends are CI(1,1). The opposite is not true since unobserved and nonstationary $\pi_t$ can make stock prices and dividends non-cointegrated.

Based on equation (3.8), Diba and Grossman (1988a) obtain the estimation equation as follows:

(3.9) $$P_{t+1} + \gamma D_{t+1} - (1/\delta)P_t = e_{t+1} - \pi_{t+1}$$

where $\gamma$ is a positive constant and the ratio of expected dividends relative to expected capital gains, and $e_{t+1} = P_{t+1} + \gamma D_{t+1} + \pi_{t+1} - E_t(P_{t+1} + \gamma D_{t+1} + \pi_{t+1})$. From the view of rational expectations, *e* is serially uncorrelated. Then the stationarity of left-hand side of equation (3.9) becomes equivalent to the stationarity of $\pi_{t+1}$. Their test depends on the existence of a cointegration relationship between $P_{t+1} + \gamma D_{t+1}$ and $P_t$, the cointegration relationship indicates the stationarity of unobservable variable. Moreover, they have shown that there is a cointegration relationship if $\gamma$ is in the range (0.5, 2). If $\gamma$ is one, they significantly reject the null hypothesis of no cointegration. As a result, they have shown that there are no bubbles in the U.S stock market. However, Evans (1991) has raised doubts about this conclusion by rational bubbles that appear to be stationary when unit root tests are applied, even though they are explosive.

Froot and Obstfeld (1991) have introduced intrinsic bubbles more directly. Bubbles can be made from market fundamental factors such as dividends. As Froot and Obstfeld (1991), the compounding present value model is follows:

(3.10) $$P_t = e^{-rt} E_t (D_{t+1} + P_{t+1})$$

where $r$ is the constant, instaneous real rate of interest rate.

Solution becomes

(3.11) $$P_t^{PV} = \sum_{s=t}^{\infty} e^{-r(s-t)} E_t (D_t)$$

If a transversality condition is not imposed,

(3.12) $$P_t = P_t^{PV} + B_t$$

They have assumed that $\ln(D_t)$ can be generated by the geometric martingale.

(3.13) $$d_{t+1} = \mu + d_t + \xi_{t+1}$$

where $\mu$ is the trend growth in dividends and $d_t$ is the logarithm of dividends at time $t$, and $\xi \sim N(0, \sigma^2)$.

Under the geometric martingale, the solution becomes

(3.14) $$P_t^{PV} = (e^r - e^{\mu + \frac{\sigma^2}{2}})^{-1} D_t$$

On the other hand, they have assumed that bubbles have a non-linear relationship with dividends.

(3.15) $$B(D_t) = \gamma D_t^{\lambda}$$

where λ are the roots of

(3.16) $$\frac{\lambda^2\sigma^2}{2} + \lambda\mu - r = 0$$

Plug into equation (3.15) and equation (3.16) into equation (3.12), then they get the equation for intrinsic bubbles as follow:

(3.17) $$P_t(D_t) = P_t^{PV} + B(D_t) = \chi D_t + \gamma D_t^{\lambda}$$

where $\chi = (e^r - e^{\mu + \frac{\sigma^2}{2}})^{-1}$

Bubbles depend on market fundamentals such as dividends *D*. Equation (3.17) shows that bubbles are not originated from extraneous variables, but from dividends. Divide $D_t$ and include the error term to reflect errors which occurs from stock price movements, then the estimation equation for equation (3.17) become

(3.18) $$\frac{P_t}{D_t} = \gamma_0 + \gamma D_t^{\lambda - 1} + \pi_t$$

In equation (3.18), no existence of bubbles means that there are parametric restrictions as following: $\gamma_0 = \chi$ and $\gamma = 0$ for the null hypothesis, and $\gamma_0 = \chi$ and $\gamma > 0$ for the alternative hypothesis. Froot and Obstfeld (1991) employ the F-statistics to test the joint hypothesis and reject the null hypothesis in the U.S. In addition, $\gamma$ is significant and positive. Therfore, bubbles overreact to changes in fundamentals. To introduce extraneous bubbles, they add time trend or time-dependent bubbles ( $e^{r - \mu - \frac{\sigma^2}{2}}$ ) respectively.

(3.19a) $$\frac{P_t}{D_t} = \gamma_0 + \beta_1 t + \eta_t$$

(3.19b) $$\frac{P_t}{D_t} = \gamma_0 + \beta_2 e^{(r-\mu-\frac{\sigma^2}{2})t} + \eta_t$$

Equation (3.19a) denotes the model which has only extraneous time trend and Equation (3.19b) encompass extraneous time-dependent bubbles. In the U.S. market, $(\beta_1, \beta_2)$ are significant. In the non-linear case, they find that the coefficient of bubbles $\gamma$ is only significant when they add non-linear bubbles.

(3.20a) $$\frac{P_t}{D_t} = \gamma_0 + \gamma D_t^{\lambda-1} + \beta_2 t + \eta_t$$

(3.20b) $$\frac{P_t}{D_t} = \gamma_0 + \gamma D_t^{\lambda-1} + \beta_1 e^{(r-\mu-\frac{\sigma^2}{2})t} + \eta_t$$

They conclude that there are bubbles if stock prices have non-linearly correlated with bubbles and the relation is significant. However, it is not plausible that there exist bubbles if the dividend-price ratio (D/P) and dividends (D) have a non-linearly positive correlation.

### 3.3.1. *The Information Error Model*

Here is my information error model. The present value model is

(3.21) $$P_t = \sum_{k=1}^{\infty} \delta^k E_t(D_{t+k})$$

Bring one period backward,

(3.22) $$P_{t-1} = \sum_{k=1}^{\infty} \delta^k E_{t-1}(D_{t+k-1})$$

Multiplying $\delta$ by (3.21) and subtracting it from equation (3.22),

$$P_{t-1} - \delta P_t = E_{t-1}\delta D_t - \sum_{k=1}^{\infty} \delta^{k+1}[E_t - E_{t-1}]D_{t+k} \tag{3.23}$$

Rearranging equation (3.23) to obtain

$$\delta(P_t + D_t) - P_{t-1} = \sum_{k=0}^{\infty} \delta^{k+1}[E_t - E_{t-1}]D_{t+k} \tag{3.24a}$$

$$P_t + D_t = \frac{1}{\delta}P_{t-1} + \sum_{k=0}^{\infty} \delta^{k}[E_t - E_{t-1}]D_{t+k} \tag{3.24b}$$

where $\varepsilon_t$ denotes the present value of the sum of the forecasting error of participants in the market. That is, it is the response of market participants about market fundamental factors since $\varepsilon_t$ is the difference between the forecasting error for dividends at a previous period and the forecasting error for dividends based on new information at the current period. Let's call (3.24b) an information error model. In this model, $\varepsilon_t$ is the measure of non-fundamental in stock prices since rational bubbles are an over-reaction or under-reaction of market participants about new information for market fundamentals.[13] Moreover, $\varepsilon_t$ is a serially uncorrelated.[14]

As a matter of fact, $\varepsilon_t$ is the same as the unobserved variable in the Hamilton-Whiteman and Diba-Grossman models. My information error model can be written as follows:

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[13] If there is no flow of new information about future dividends at time $t$, $v_t = 0$. Then, I obtain a perfect equilibrium situation. When this happens, the current price depend only the previous period's price and current dividends.

[14] See Shiller (1981a) for proving this by using the law of iterated expectations.

(3.25) $$P_{t+1} = \frac{1}{\delta} P_t - E_t D_{t+1} + v_{t+1}$$

where $v_{t+1} = \sum_{k=1}^{\infty} \delta^k [E_t - E_{t-1}] D_{t+k+1}$

Since $E_{t+1} D_{t+1} - E_t D_{t+1} = e_{t+1}$,

(3.26) $$P_{t+1} + D_{t+1} - \frac{1}{\delta} P_t = e_{t+1} + v_{t+1}$$

This is the same that $\varepsilon_{t+1} = e_{t+1} - \pi_{t+1}$ or $\varepsilon_{t+1} = e_{t+1} + v_{t+1}$, and $\gamma = 1$ in the Diba-Grossman model. Therefore, I am able to measure rational bubbles from the deviation of stock prices ($P_t + D_t - \frac{1}{\delta} P_{t-1}$) without an unobserved and arbitrary random variable. One more implication for my model is that if the market is efficient, forecast errors will be completely absorbed in the current price, and it is unlikely that the equilibrium error will be corrected in the next period, because forecast errors at time $t$ do not constitute new information in the next period. Since the efficient markets hypothesis expects the instantaneous adjustment of prices to new information, a pattern of a systematic slow adjustment to new information would imply the existence of profitable arbitrage opportunities.

### 3.3.2. *A Review of the Survival Analysis*

Non-fundamentals in stock prices from the information error model make up two parts. One is rational bubbles which are positive and the other is negative part. Rational bubbles are a rare event. Like other rare events, rational bubbles can be approached in terms of the instantaneous rate at which an event occurs after duration $t$ since some prior event has taken place. It is important to distinguish a random variable from a particular outcome of a random variable. Let $T$ denote a duration which has some distribution in the population and let $t$ denote a particular value. In the survival analysis, $T$ is the length of time the subject lives. $T$ would be the time until the market becomes explosive. If $T$ has a probability density function $f(t)$ and a cumulative density function is defined as $F(t)=P(T\leq t)$, $t\geq 0$, then the existing probability, that is, the survival function $S(t)$ for optimistic forecasting or expectation until a specific survival past time $t$ are as follows:

(3.27) $$S(t) = \Pr(T > t) = \int_t^{\infty} f(t)dt = 1 - F(t)$$

where $F(t)$ is the cumulative density function, also called the survival function.

The probability that bubbles will explode from $T=t$ to $T=t+\Delta t$ is that

Pr(Explosion between $t$ and $t+\Delta t$) / Pr(No explosion until $t$), then

(3.28) $$\theta(t) = \lim_{\Delta t \to 0} \frac{\Pr\left(t \leq T < t + \Delta t \mid T \geq t\right)}{\Delta t}$$

Let $\theta(t)$ be the bursting rate at time $t$. The bursting rate, which is also called the hazard rate in statistics, has the interpretation $\theta(t)$=Prob{the burst of bubbles in the next small unit of time $\Delta t$, given bubbles have survived to time $t$}. Therefore, the bursting rate is a conditional probability between $t$ and $t+\Delta t$.

Similarly, I can express the numerator of the bursting rate in terms of *cdf* as follows:

(3.29) $$\Pr\left(t \leq T < t + \Delta t \mid T \geq t\right) = \Pr\left(t \leq T < t + \Delta t\right) / \Pr\left(T \geq t\right) = \frac{[F(t+\Delta t) - F(t)]}{1 - F(t)} = \frac{[F(t+\Delta t) - F(t)]}{S(t)}$$

When the *cdf* is differentiable, I can take the limit of right-hand side, divided by $\Delta t$, as $\Delta t$ approaches zero, then

(3.30) $$\theta(t) = \lim_{\Delta t \to 0} \frac{[F(t+\Delta t) - F(t)]}{\Delta t} \cdot \frac{1}{S(t)} = \frac{F'(t)}{S(t)} = \frac{f(t)}{S(t)} = \frac{f(t)}{1-F(t)}$$

Therefore, the bursting rate is

(3.31) $$\theta(t) = -\frac{S'(t)}{S(t)} = -\frac{d \ln S(t)}{dt}$$

From (3.31), I can derive the probability density function as follows:

(3.32) $$\ln S(t) = -\int \theta(t) dt + \ln c$$

then, $S(t) = ce^{-\int \theta(t) dt}$ and $F(t) = 1 - ce^{-\int \theta(t) dt}$

If $f(t)$ follows the Weibull distribution,

(3.33) $$f(t) = \alpha \lambda t^{\alpha - 1} \exp(-\lambda t^{\alpha})$$

then the bursting rate, the so called hazard function, is given as

(3.34) $$\theta(t) = \alpha \lambda t^{\alpha - 1}$$

When λ is equal to a constant, the bursting rate of extraneous bubbles is given as $\theta(t) = \alpha t^{\alpha-1} \exp(\beta_0)$. Let $\lambda = \exp(X'\beta)$, the bursting rate of intrinsic bubble model is given as $\theta(t) = \alpha t^{\alpha-1} \exp(X'\beta)$. To sum up these arguments,

(3.35a) $$\theta(t) = \alpha t^{\alpha-1} \exp(\beta_0)$$

(3.35b) $$\theta(t) = \alpha t^{\alpha-1} \exp(X'\beta)$$

where $X$ includes a constant and the set of market fundamentals. $\alpha$ denotes the shape parameter, also known as the Weibull slope. Different values of the shape parameter can have marked effects on the behavior of the distribution. As a matter of fact, some values of the shape parameter will cause the distribution equations to reduce to those of other distributions. For example, when $\alpha = 1$ and $\lambda = \exp(X'\beta)$, the *pdf* of the two-parameter Weibull reduces to that of the one-parameter exponential distribution. The bursting rate (hazard function) $\theta(t)$ depends on $\alpha$. There are four cases.

(a) If $\alpha$ is smaller than one, the bursting rate decreases.

(b) If $\alpha$ is equal to one, the bursting rate is constant.

(c) If $\alpha$ is between one and two, the bursting rate increases at a decreasing rate.

(d) If $\alpha$ is greater than two, the bursting rate increases at an increasing rate. Figure 3.1 shows several cases of Weibull distributions.

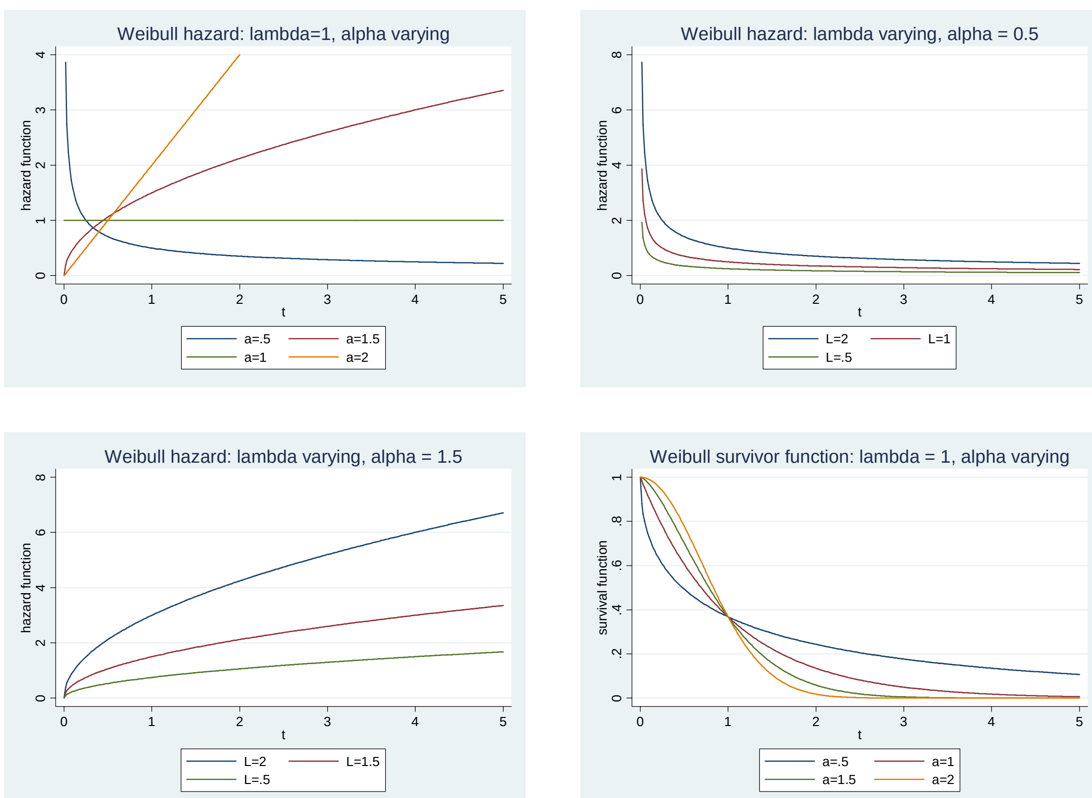


FIGURE 3.1. − Weibull distributions

## 3.4. EMPIRICAL ANALYSIS

### 3.4.1. *The Description of the Data*

I use variables for measuring stock prices and the state of the economy. First, variables that measure stock prices consist of the S&P 500 Composite Price Index, the Price Earnings Ratio (PER), and the Dividend Yield (DY). I get dividends (DIV) which have an index type using the S&P 500 Composite Price Index and the Dividend Yield.

Second, variables for the state of the economy consist of discount factors and market fundamental factors. The former are short-term interest rates (3-Month T-Bills), long-term interest rates (10-Year T-Bonds), and a term premium, exchange rates (Dollar/Yen). The latter is unemployment rates and default rates. The data period is from 1980m1 to 2007m8. The data are from *http://www.datastream.com.*

The term premium shows the forecast of the future economy. In the well-developed market, short-term interest rates reflect the policy rate, but long-term interest rates include the expectation of the economy of market participants. By the expectation theory, which explains the term structure of interest rates, long-term interest rates are equal to a geometric average. If the state of the economy is expected to be better in the future, the market participants' expected inflation goes up, expected short-term interest rates also go up, and current long-term interest rates will increase. The opposite during the bad economy also applies. Therefore, the term premium is the index of the market participants' forecast for the economy. The figure 3.2 shows the trend of the S&P 500 Composite Price Index.

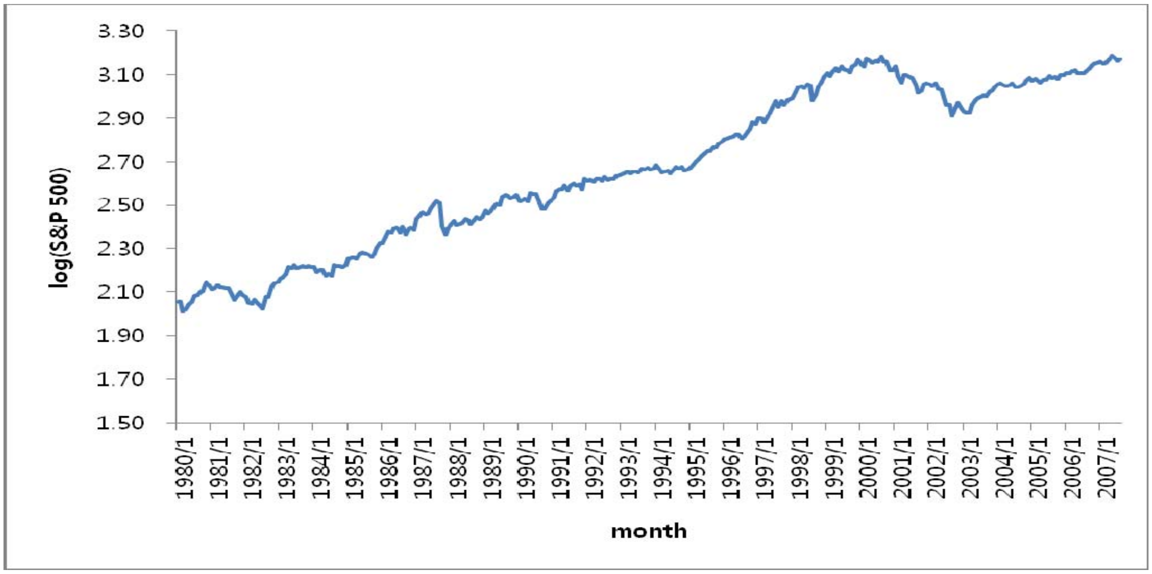


FIGURE 3.2. – Movement in the S&P 500 Index

3.4.2. *Measuring Non-fundamentals in Stock Prices*

To measure non-fundamentals in stock prices, I use the information error model. There needs a criterion which appropriately describes current stock prices. Whether the stock prices are too high or low, the information error model provides this criterion. I divide current stock prices into two parts, fundamental stock prices explained by $P_t = -D_t + \frac{1}{\delta} P_{t-1}$ and rational bubbles which deviate from the expected orbit of the information error model. The coefficient $\frac{1}{\hat{\delta}}$ is 1.022 and significant. Figure 3.3 shows estimated non-fundamentals in stock prices. The series of non-fundamentals is serially uncorrelated since we do not reject the Ljung-Box Q statistic, which follows a chi-square distribution under the null hypothesis that the series exhibit white noise processes, at lag 8 and 16 at the 5% significance level.

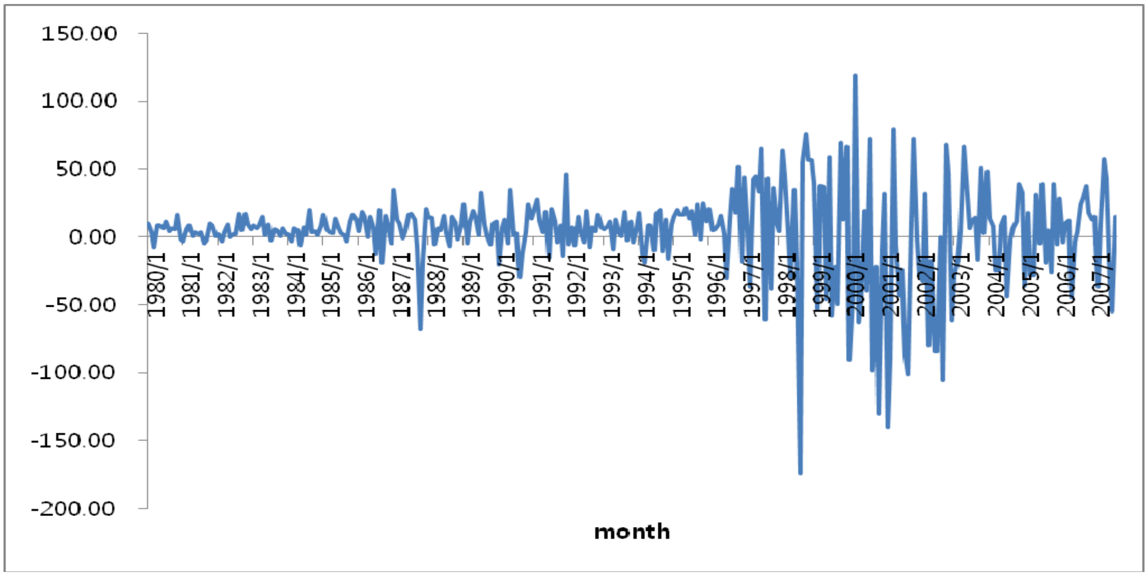


FIGURE 3.3. – Non-fundamentals in stock prices

### 3.4.3. *Tests for Unit Roots*

It is important for time series data to have a stationary property. If some variables are not stationary, regressions using those variables can be spurious. If those variables are stationary, they do not have unit roots. There are several types of tests for unit roots. The Dickey-Fuller test (DF) was widely used before the Phillips-Perron test (PP) was introduced. The Dickey-Fuller test has a problem, which is affected by autocorrelation and the heteroskedasticity of error terms. Therefore, I use the Phillips-Perron test to admit serial dependency and heterskedasticity.[15] The Phillips-Perron test has three types of tests for unit roots.

(a) $$\Delta y_t = \alpha \hat{y}_{t-1} + \hat{u}_t$$

(b) $$\Delta y_t = \mu + \alpha y^*_{t-1} + u^*_t$$

(c) $$\Delta y_t = \mu + \beta(t - T/2) + \alpha y_{t-1} + u_t$$

The corresponding Phillips-Perron statistics are as follows:

(a) $Z(t_\alpha)$ for $H_0 : \hat{\alpha} = 1$ in (a)

(b) $Z(t_\alpha)$ for $H_0 : \alpha^* = 1$ in (b)

(c) $Z(t_\alpha)$ for $H_0 : \alpha = 1$ in (c)

Cases (b) and (c) are for generating general unit roots. The case (b) is appropriate when the alternative hypothesis is that the series is stationary around a fixed mean and the case (c) is appropriate when the alternative hypothesis is that series is stationary around a trend. Table 3.1 shows the results of the Phillips-Perron test. Ratioanl bubbles are stationary. These results are similar with the Diba-Grossman results. The term premium is stationary. However, all other variables have a unit root.

[15] I do not report the results of the ADF test but the results are almost similar to those of the PP test.

TABLE 3.1

TESTS FOR UNIT ROOTS

| Variables | Phillips-Perron test | |
|---|---|---|
| | $\Delta y_t = \mu + \alpha y_{t-1} + u_t$ | $\Delta y_t = \mu + \beta t + \alpha y_{t-1} + u_t$ |
| | All Samples | All Samples |
| Stock Prices | 0.22 | -1.83 |
| Bubbles | -18.22* | -18.29* |
| PER | -2.12 | -2.21 |
| Dividend | 3.73 | 2.10 |
| Short-term Interest Rates | -2.18 | -2.81 |
| Long-term Interest Rates | -1.35 | -3.30 |
| Term Premium | -4.37* | -4.55* |
| Exchange Rates | -1.80 | -1.61 |
| Unemployment Rates | -1.49 | -2.83 |
| Default Rates | -2.16 | -1.92 |

Note: (1) Values denote t-statistics for α.
(2) * Indicates significance at the 5% level.
(3) Test allows for a constant, and a constant and a linear trend respectively; one-sided (lower-tail) test of the null hypothesis that the variable has a unit root.

### 3.4.4. *Tests for Rational Bubbles*

I make a strong assumption in orter to use the survival analysis. The survival analysis requires only two outcomes, which is a burst or not, based on duration. Therefore, I assume that there is a burst when, after a sum of positive values, a non-fundamental in stock prices is negative. This assumption is quite strong and no doubt overstates the number of bubbles, but it allows me to take this first look at the survival analysis and bubbles. Table 3.2 describes the summary of the data on duration.

TABLE 3.2

THE SUMMARY OF THE DATA ON DURATION

| Number of Episode | Average Duration | Max. of Duration |
|---|---|---|
| 102 | 2.34 | 15 |

Purely extraneous bubbles are originated from not the current or future results of firms but from the participants' mood or stimulus. From the hazard model (3.35a) and the assumption for a burst which has a positive non-fundamental, rational bubbles have the increasing burst rate by extraneous factors which are not related to market fundamentals if $\alpha$ is much greater than two and significant. Table 3.3 reports that there increasing bursting rate at a decreasing rate for extraneous bubbles in the U.S. stock market.

TABLE 3.3

TESTS FOR EXTRANEOUS BUBBLES

| Model | All Samples |
|---|---|
| $\theta(t)=\alpha t^{\alpha-1}\exp\left(\beta_0\right)$ | $1.32^{*}$ |

Note: (1) Upper values are α.
(2) A constant term is not reported.
(3) * Indicates significance at the 5% level.

The next step is for checking intrinsic bubbles. I include the available fundamental set of variables which affects stock prices in the survival analysis to investigate whether or not the bursting rate are affected by market fundamentals. Among these market fundamentals factors, I take the first difference if variables include a unit root. Table 3.4 reports the results of tests about intrinsic bubbles in the U.S. stock market. From the hazard model (3.35b), rational bubbles have the increasing burst rate by market fundamental factors if $\alpha$ is much greater than two and significant. Table 3.4 reports that there is the increasing bursting rate at a decreasing rate for intrinsic bubbles in the U.S. stock market. This says that there is the increasing bursting rate at a decreasing rate for bubbles becoming explosive by market fundamental factors.

TABLE 3.4

TESTS FOR INTRINSIC BUBBLES

| Model | All Samples |
| --- | --- |
| $\theta(t) = \alpha t^{\alpha-1} \exp(\beta_0 + \beta_1 PER_t)$ | $1.42^*$<br>$0.04^*$ |
| $\theta(t) = \alpha t^{\alpha-1} \exp(\beta_0 + \beta_1 SR_t)$ | $1.36^*$<br>$-0.08^*$ |
| $\theta(t) = \alpha t^{\alpha-1} \exp(\beta_0 + \beta_1 LR_t)$ | $1.40^*$<br>$-0.11^*$ |
| $\theta(t) = \alpha t^{\alpha-1} \exp(\beta_0 + \beta_1 TERM_t)$ | $1.35^*$<br>$-0.13^*$ |
| $\theta(t) = \alpha t^{\alpha-1} \exp(\beta_0 + \beta_1 EX_t)$ | $1.44^*$<br>$-0.01^*$ |
| $\theta(t) = \alpha t^{\alpha-1} \exp(\beta_0 + \beta_1 UNEMP_t)$ | $1.43^*$<br>$-0.25^*$ |
| $\theta(t) = \alpha t^{\alpha-1} \exp(\beta_0 + \beta_1 DEF_t)$ | $1.35^*$<br>$0.19^*$ |

Note: (1) Upper values are α and lower values are $\beta_1$'s in each column.
(2) * Indicates significance at the 5% level.

### 3.4.5. *Tests for the Market Efficiecy*: *Cointegration Tests*

I test if stock prices and other variables have a conintegration relationship. Some variables are nonstationary, but the linkage of those variables can be stationary. If variables have the same integrated order of *d* and they have a common trend, then those variables can be stationary. This means that variables are in the long-run equilibrium.

There are two different hypotheses for the efficiency of the asset market and cointegration. One is the Granger hypothesis. If the asset market is efficient, the price of two different assets cannot be cointegrated. If the asset price for A and B have the relationship of linear conintegration, the two assets are long-run stationary. In this case, participants can forecast the price of one asset with the price of the other asset. This is a contradiction for the hypothesis of efficiency. The other hypothesis is examined by Campbell and Shiller (1988). The dividend-price ratio (D/P), which is a long-run return for stock prices, has a significant relationship with the growth of dividends although stock prices are not explained by market fundamentals. I can expect that dividend-price ratio (D/P) has a relationship with macroeconomic variables.[16] Therefore, the stock market is efficient in indirect methods. I use an augmented Engle-Granger test, which is a residual based test, since the result of the Johansen test can be different due to the decision of lag terms. Moreover, Johansen test exhibits serious nominal size distortions leading to spurious conintegration in finite samples.

First of all, I test linear cointegration, which is based on the Granger's Hypothesis (1986). The Granger hypothesis has stated that if the asset market is efficient, the price of two different assets cannot be cointegrated (linear cointegration). If I cannot reject the null hypothesis that has no cointegration, there is no common trend among stock prices and other variables. That is to say, market participants cannot forecast stock prices with other variables so the market of stock prices is efficient.

[16] Diba and Grossman (1998a) and Yuhn (1996) also provide that stock prices and fundamental factors have a nonlinear cointegraion relationship.

TABLE 3.5

TESTS FOR THE GRANGER'S HYPOTHESIS

| Variables | Augmented Engle-Granger Test | |
|---|---|---|
| | $Stock_t = \mu + \alpha \cdot Variables_t + u_t$ | $Stock_t = \mu + \beta t + \alpha \cdot Variables_t + u_t$ |
| | All Samples | All Samples |
| PER | -1.95 | -1.77 |
| Dividend | -0.64 | -1.81 |
| Short-term Interest Rates | -2.02 | -2.60 |
| Long-term Interest Rates | -2.07 | -2.34 |
| Term Premium | $-3.52^{+}$ | $-2.74^{+}$ |
| Exchange Rates | -1.46 | -2.08 |
| Unemployment Rates | -2.45 | -1.75 |
| Default Rates | -1.73 | -1.84 |

Note: (1) Values denote AEG statistics with a constant and a linear trend.
(2) * Indicates significance at the 5% level.
(3) $^{+}$ Indicates that the variable is stationary.

Table 3.5 reports that stock prices and all variables except the term premium which is stationary have no cointegration relationship in the whole period. Therefore, these variables can not explain the movement of stock prices. That is to say, these variables and stock prices show that market is efficient indirectly.

In addition, I test nonlinear cointegration, which is based on the Campbell and Shiller's hypothesis (1988). Campbell and Shiller (1988) have found that the dividend-price ratio (D/P), which is a long-run return for stock prices, has a significant relationship with the growth of dividends although stock prices are not explained by market fundamentals. Therefore, the dividend-price ratio (D/P) has a relationship with macroeconomic variables. Table 3.6 shows that there are no cointegration relationships between the dividend-price ratio and fundamentals except long-term interest. Only long-term interest rate can explain the dividend-price ratio so the dividend-price ratio (D/P) has the significant relationship with long-term interest rate even if stock prices are not explained by other market fundamentals.

TABLE 3.6

TESTS FOR THE CAMPBELL AND SHILLER'S HYPOTHESIS

| Variables | Augmented Engle-Granger Test | |
|---|---|---|
| | $Div_t \,/\, Stock_t = \mu + \alpha \cdot Variables_t + u_t$ | $Div_t \,/\, Stock_t = \mu + \beta t + \alpha \cdot Variables_t + u_t$ |
| | All Samples | All Samples |
| PER | -3.41 | -3.32 |
| Short-term Interest Rates | -3.08 | -2.49 |
| Long-term Interest Rates | -3.97$^{*}$ | -3.20 |
| Term Premium | -2.24$^{+}$ | -2.21$^{+}$ |
| Exchange Rates | -2.67 | -2.71 |
| Unemployment Rates | -3.74 | -3.11 |
| Default Rates | -2.61 | -2.48 |

Note: (1) Values denote AEG statistics with a constant and a linear trend.
(2) * Indicates significance at the 5% level.
(3) $^{+}$ Indicates that the variable is stationary.

## 3.5. CONCLUDING REMARKS

The purpose of this study is to test the bursting rate of rational bubbles in the U.S. stock market. There has also been a resurgence of interest in bubbles in stock prices primarily through the work of Hamilton and Whiteman (1985), Diba and Grossman (1988a, b, c), Evans (1991), Froot and Obstfeld (1991), and others.

The bubble models devised by these authors represent a significant departure from the conventional paradigm in that they reinterpret rational bubbles in terms of market fundamentals. These bubbles may be termed intrinsic bubbles, as opposed to extraneous bubbles in the traditional view. It has been widely believed that bubbles do not easily lend themselves to direct testing. The attractive feature of the intrinsic bubble specification may be found in its ability to derive testable implications for bubbles by investigating the stationarity properties of stock prices and dividends or by parameterizing a specific bubble relationship as a function of market fundamentals.

However, the existing approach to intrinsic bubbles still remains unsatisfactory. As Evans explains, Diba and Grossman's tests for stationarity are unable to detect an important class of rational bubbles. The ability of Froot and Obstfeld's parametric tests to discover bubbles is also doubtful. The essence of their tests is that if the price-dividend ratio is significantly (and nonlinearly) related to current dividends, the hypothesis of no bubbles is rejected. More fundamentally, the recent tests for intrinsic bubbles may be characterized as being indirect in that explicit measures of bubbles are not directly related to market fundamentals.

In this study, I formulate an information error model which allows one to derive the measure non-fundamental in stock prices in a straightforward manner. My information error model also enables one to test for market efficiency by relating the flow of information. This study also provides a new method by specifying rational bubble measures as the Weibull distribution. This study is the first attempt to apply the Weibull distribution to the test of rational bubbles. There is not only a parallel between the burst of a speculative bubble and a material's burning out, but also there is a good reason to believe that measured rational bubbles can be appropriately modeled as the Weibull specification. As a result, my empirical analysis is the first step in applying survival analysis to bubbles, and it reveals preliminary evidence that there is the increasing bursting rate at a decreasing rate for extraneous or instrinsic bubbles in the U.S. stock market.

I have also conducted two different cointegration tests for market efficiency: One is concerned with the Granger (1986) hypothesis, and the other is concerned with the Campbell-Shiller (1988) hypothesis. The Granger hypothesis states that if the stock market is efficient, the prices of two assets cannot be cointegrated (linear cointegration). On the other hand, Campbell-Shiller's hypothesis (1988) postulates that the dividend-price ratio, which is a long-run return for stock prices, has a significant cointegration relationship with the growth of dividends even if stock prices are not explained by fundamentals. Therefore, the dividend-price ratio (D/P) has a non-linear relationship with macroeconomic variables. In the Granger hypothesis I have accepted the null hypothesis of no linear cointegration between stock prices and economic variables.

Therefore, the market is efficient indirectly. In the Campbell-Shiller's hypothesis, there are no cointegration relationships between the dividend-price ratio and fundamentals except long-term interest.

# CHAPTER IV

# TRANSMISSION OF STOCK PRICES AND VOLATILITY FROM THE U.S., CHINESE, AND JAPANESE MARKETS ON THE KOREAN MARKET

## 4.1. INTRODUCTION

Recently, the world economy is increasingly interdependent through trade, common creditors, and similar macroeconomic trends. In many countries, the Chinese economy as well as the traditional U.S. economy has affected the Korean economy. The Japanese economy has overcome 'lost 10 years' in the economic growth, its economy has influenced on the Korean economy.[17] Moreover, the Asian financial crisis that occurred in the late 1990s has affected Asian financial markets including the Korean market in a significant way. In particular, Asian emerging markets including the Korean market underwent seismic changes in the wake of the financial turmoil. Many regulations and restrictions on trading activities in this market have been eased or eliminated since the financial crisis, and Asian emerging markets have become much more globalized and liberalized. Foreign portfolio investment in these markets has grown at a galloping pace as these markets have been increasingly integrated into the world financial market.[18]

[17] In August 2008, the market value of the Chinese stock market is 4,720 billion dollar and that of the Japanese stock market is 4,700 billion dollar so the Chinese stock market becomes the biggest market in the Asian stock market.

[18] The most dramatic developments occurred in the Korean market. Since the Korean government began to open its stock market to foreign investors in 1992, it raised the limit of foreigners' investment in a stock traded on the Korean Stock Exchange six times to 26 percent of shares of the stock until the 1997 financial crisis hit the Korean economy, but an individual's acquisition of a Korean stock was limited to 7 percent. As the financial crisis was looming ahead, the limit of foreigners' investment was raised to 50% on December 11, 1997, then to 55 percent on December 30, 1997, and finally the restrictions on foreign investment in Korean stocks were completely eliminated on May 25, 1998. The market value of Korean stocks held by foreign investors was 10,692.2 billion Won at the end of 1998, but the holdings of Korean

Asian emerging markets have emerged as an important segment of the world financial system. In these Asian emerging markets, the Korean stock market is one of important and unusual market to measure the change of the financial market.[19]

The growing integration of world markets raises several fundamental questions. Does the globalization of financial markets precipitate the transmission of information from advanced markets into the relatively small markets? If the small markets become more globalized and liberalized, then information on stock prices produced in a leading market such as the U.S., Chinese, and Japanese markets will be more rapidly disseminated into the small market, thus prompting price spillovers. In fact, since the 1997 financial crisis changed the financial landscape in Asia, the influence of advanced stock markets on the small market has gained steadily. Notably, the stock prices of the Korean market have tended to synchronize with the sharp moves of U.S., Chinese, and Japanese stock prices.

What is of particular interest from both practical and theoretical perspectives is whether such co-movements in stock prices amplify the volatility of Korean market. The co-movements of stock prices across markets may change stock prices above or below the levels dictated by market fundamentals, potentially creating market volatility. However, if investors become more informed as a result of globalization of markets, the increased interdependence and linkage of financial markets could reduce the

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stocks by foreign investors increased to 95,115.4 billion Won as of April 16, 2004, recording a 789.6 percent increase during this period. The market value of stocks of the 10 largest business groups held by foreign investors accounts for 50.3 percent of the total market value of Korean stocks traded on the Korea Stock Exchange and the KOSDAQ.

[19] Nam, Yuhn, and Kim (2008) report that the price spillover effect picked up significantly in the Korean market after the crisis, and the volatility spillover effect increased dramatically in the Korean market.

transmission of volatility from one market to another. Thus, one interesting hypothesis to be tested concerns whether or not an increased integration of financial markets leads to a reduction in market volatility in the Korean market.

The main purpose of this study is to investigate how the Korean stock market has been affected by the 1997 financial crisis. I am particularly interested in whether the influences of shocks originated in the U.S., Chinese, and Chinese markets on prices and market volatility in the Korean market increased or decreased after the 1997 crisis. It is well documented that the volatility of the Korean market was historically high. Interestingly enough, the volatility of this market has been much dampened since 2000 when the aftermath of the financial crisis has been substantially subdued. This study aims to explore whether much of the slowdown in Korean markets' volatility after the financial crisis in 1997 is a fundamental shift or a temporary fad.

The Korean market presents some features. First, Korea enforced strict controls on capital transactions until the 1997 financial crisis. Second, Korea severely suffered from speculative attacks on their currencies that plunged the countries into a full-scale financial crisis in 1997. Third, Korea changed their exchange rate regime in the midst of the financial crisis. Korea shifted gears toward a floating exchange rate system from a market average exchange rate system in December, 1997. There has been a spate of studies on the interaction and interdependence of stock prices and volatility among advanced markets. There are also a number of studies that examine price (or return) and volatility spillovers from advanced markets to other small markets. For example, see Ng (2000), In, Kim, Yoon, and Viney (2001), Edwards and Susmel (2001), Darrat and

Zhong (2002), Worthington and Higgs (2004), Nam, Yoon, and Kim (2008). However, there is little literature that investigates what happened to the Korean market after the 1997 financial crisis and that investigate Chinese and Japanese stock markets' effect on the Korean stock market. This study offers some new evidence on how the Korean market responded to shocks produced in the United States, China, and Japan after the crisis. I compare spillover effects between the prior- and post-crisis periods using daily data from January 3, 1995 to July 31, 2007. To this end, I utilize an EGARCH model which is known to be suitable for modeling the asymmetric transmission of volatility.

This study is organized as follows: In section 4.2 I briefly review the literature on price and volatility spillovers. Section 4.3 discusses the methodology employed in this study. Section 4.4 describes the data used in this study and analyzes basic characteristics of the data. Section 4.5 presents empirical results and their implications. Concluding remarks are provided in section 4.6.

## 4.2. A REVIEW OF THE LITERATURE

Earlier studies focused on the interdependence of stock prices among advanced markets. Jaffe and Westerfield (1985) studied spillover effects among the U.S., U.K., Australian, Canadian, and Japanese markets using daily closing prices and an OLS and found interdependence among these markets. Eun and Shim (1989) investigated price spillovers in nine advanced markets using a VAR model and found that innovations in the U.S. market were rapidly transmitted to other markets, whereas no single foreign market significantly explained U.S. market movements. Barclay, Litzenberger, and Warner (1990) found positive correlations between the New York and Tokyo markets using daily stock prices. King and Wadhwani (1990) provided some evidence in support of the "contagion effect" in the New York, London, and Tokyo markets, showing that negative shocks in one market were immediately transmitted to other markets. Koch and Koch (1991) analyzed the lead-lag relationship among eight stock markets. Their study revealed that there was a tendency for the regional interdependence of stock markets to increase and that the spillover effect of the New York market on the Tokyo market was pronounced.

More recent studies concentrate on the international interactions of stock returns and volatility in terms of the first and second moments of returns utilizing recent advances in time series analysis such as GARCH-type models. For example, Hamao, Masulis, and Ng (1990), using a single-variable GARCH-M model, have examined price spillovers (interdependences of the first moments) and volatility spillovers (interdependences of the second moments) among the New York, London, and Tokyo markets. Their study has confirmed the presence of a spillover effect from the New York and London markets to the Tokyo market, but found no spillover effect from the Tokyo market to either the New York or London market. Subsequent studies such as Ng, Chang, and Chou (1991) and Theodossiou and Lee (1993) using a GARCH model, Susmel and Engle (1994) using an ARCH model, and Gilmorea and McManus (2002) and Hsiao, Hsiao, and Yamashita (2003) using a VAR have also presented evidence for spillover effects mainly in advanced markets.

However, most of previous studies failed to incorporate asymmetry in price and volatility spillovers. Nelson (1991) has developed an EGARCH model to study the asymmetrical effects of shocks on stock return volatility in the U.S. market. He has discovered that in the U.S. market negative shocks had larger impacts on volatility than positive shocks. Koutmos and Booth (1995), noticing that a market's volatility responds asymmetrically to its own past shocks, have shown that negative shocks originated in one market exert greater spillover effects on other markets than do positive shocks. More specifically, they have used a multivariate EGARCH model to analyze spillovers of daily stock prices and volatility among the New York, Tokyo, and London markets and confirmed an asymmetrical spillover effect from the New York market to the Tokyo and London markets, from the Tokyo market to the London and New York markets, and from the London market to the New York and Tokyo markets. Thus, they have concluded that price and volatility spillovers are generally reciprocal in the sense that two markets influence each other.

Several studies have examined price and volatility spillovers among emerging markets including the Korean market. Ng (2000) has studied the magnitude and changing nature of volatility spillovers from the United States and Japan to six Pacific-Basin countries (Hong Kong, South Korea, Malaysia, Singapore, Taiwan, and Thailand) using weekly equity indexes denominated in U.S. dollars and found evidence for the impact of the world factors and significant spillovers from the region to many of the Pacific-Basin markets. In, Kim, Yoon, and Viney (2001) have examined volatility spillovers among three emerging markets, South Korea, Hong Kong, and Thailand during the 1997-1998 period when the Asian financial crisis spread uncontrollably. They have used a multivariable VAR-EGARCH model and observed that there were reciprocal spillovers of volatility between the Hong Kong and Korean markets, and one-directional spillovers from the Korean market to the Thai market.

Edwards and Susmel (2001) have analyzed the behavior of volatility in Argentina, Brazil, Chile, Hong Kong, and Mexico using weekly equity indexes denominated in U.S. dollars from the last week of August 1989 to the third week of October 1999. They have found strong evidence of volatility co-movements across countries, especially among the Mercosur countries. Darrat and Zhong (2002) have investigated whether the U.S. or Japanese market (or both) is the main driving force behind major movements in 11 Asia-pacific emerging markets using weekly data from November 1987 through May 1999. They have confirmed a robust cointegrating relation linking each of the emerging markets with the two matured markets of the United States and Japan. Chow and Lawler (2003) carried out a comparative analysis using Shanghai and the New York stock exchange during 1992-2002. They have not discovered that any integration between the Chinese market and U.S. market in the regression of rate of returns and volatility in the use of autoregressions and Granger causality tests. Worthington and Higgs (2004) have examined the transmission of stock returns and volatility among Asian markets: three developed markets (Hong Kong, Japan, and Singapore) and six emerging markets (Indonesia, South Korea, Malaysia, the Philippines, Taiwan, and Thailand) using daily data from January 15, 1988 to October 6, 2000 and found that all Asian equity markets are highly integrated. They have further discovered that mean spillovers from the developed to the emerging markets are not homogeneous across the emerging markets. Li (2007) have not found any evidence of a direct linkage between the Chinese market and U.S. market. Nam, Yuhn, and Kim (2008) have investigated what happened to the emerging markets (Hong Kong, South Korea, Malaysia, the Singapore, Taiwan, and Thailand) during and after the 1997 Asian financial crisis from the U.S. market. They showed that the Korean market was the only market whose prices as well as volatility were immune from shocks produced in the U.S. market before the 1997 financial crisis. However, price and volatility spillover effects showed up most strongly in the Korean market after the crisis.

Although there is a proliferation of the literature on the transmission of prices (or returns) and volatility among countries, few studies have investigated what happened to the Korean market during and after the 1997 Asian financial crisis from the Chinese, Japanese, and U.S. markets. This study aims to compare price and volatility spillovers from the Chinese, Japanese, and U.S. markets to the Korean market between the prior- and post-crisis periods. This study is also concerned with the comparison of asymmetric spillovers between good and bad news from the most influential markets to the Korean market.

## 4.3. METHODOLOGY

The transmission of information from one market to another market can be explored in two different ways. One can look at the price spillover effect or the volatility spillover effect. For example, the fact that information on U.S. stock prices is transmitted to stock prices in other markets implies that information on U.S. stock prices can be helpful in predicting stock price movements in other markets. On the other hand, an increase in volatility indicates excessive responses of stock prices to new information. The spillover of volatility from the U.S. market to the Korean market implies that excessive responses of stock prices in the Korean market are linked to excessive responses of U.S. stock prices.

It is frequently observed in asset markets that periods of large volatility are followed by periods of low volatility and vice versa (volatility clustering). The ARCH-type model recognizes the presence of successive periods of volatility and stability. Engle, Ito, and Lin (1990) ascribe volatility clustering to two factors. The first explanation for volatility clustering is that information itself comes in a cluster. In such a case, even if market participants react to market conditions rationally, and stock prices reflect all the available information, successive inflows of information result in a volatility clustering. Another explanation for volatility clustering is provided by non-synchronous trading among market participants who possess different volumes of information. Even if the same information is disseminated in the market, volatility associated with the information persists because market participants may have different perceptions toward the information and behave differently.

Spillover effects can be asymmetrical. Suppose that the stock prices of country A increase by $\beta$ percent when the U.S. stock prices increase by $\alpha$ percent. When I observe that the stock prices of country A fall by more than $\beta$ percent when the U.S. stock prices fall by $\alpha$ percent, I have an asymmetric price spillover effect. Such an asymmetric spillover effect can also occur in the transmission of volatility. For example, if negative shocks generated in the U.S. market have greater impacts on the volatility of market A than do positive shocks, an asymmetric volatility spillover exists.

This study adopts a two-variable EGARCH model to investigate asymmetric price and volatility spillovers. I first define $R_{it}$ as

(4.1) $$R_{it} = 100 \ln\left(\frac{P_t}{P_{t-1}}\right)$$

where $P_{t-1}$ is the closing price on the previous trading day, and $P_t$ is the closing price on the current trading day. Thus, $R_{it}$ represents the daily close-to-close return in market *i* at time *t*. Let $\mu_{it}$ be the average rate of return in the market and $\sigma_{it}^2$ be the variance of $R_{it}$ at time *t*, conditional on market information available at *t-1* ($I_{t-1}$). Then, $\varepsilon_{it}$ is a shock or innovation which is given by the difference between $R_{it}$ and $\mu_{it}$.

(4.2) $$\varepsilon_{it} = R_{it} - \mu_{it}$$

I can standardize the innovation as

(4.3) $$z_{it} = \frac{\varepsilon_{it}}{\sigma_{it}}$$

A two-variable EGARCH model can be represented by the following set of equations:[20]

(4.4) $$R_{it} = \beta_{i,0} + \sum \beta_{i,j}\varepsilon_{j,t-1} + \varepsilon_{j,t} \text{ for } i, j = 1, 2$$

(4.5) $$\sigma_{it}^2 = \exp\left[\alpha_{i,0} + \sum \alpha_{i,j} f_j\left(z_{j,t-1}\right) + \gamma_i \ln\left(\sigma_{i,t-1}^2\right)\right] \text{for } i, j = 1, 2$$

(4.6) $$f_j\left(z_{j,t-1}\right) = \left(\left|z_{j,t-1}\right| - E\left(\left|z_{j,t-1}\right|\right)\right) + \delta_j z_{j,t-1} \text{ for } j = 1, 2$$

[20] The EGARCH specification employed in this study follows that of Koutmos and Booth (1995).

Equation (4.4) expresses the daily return in market $i$ as a vector moving average. That is, the conditional mean of the rate of return in market $i$ is expressed as a function of its past innovations as well as the past innovations of other markets. Thus, coefficients $\beta_{i,j}$ for $i \neq j$ measure the magnitude of a price (or return) spillover across markets. Hamao et al. (1990) use an MA(1)-GARCH(1,1)-M formulation. There is the moving average term to consider the existence non-synchronous tradings.[21] This moving average term makes it possible for investors to interpret the same news innovation differently in the same markets or the same news innovation at the different views each other.

Equation (4.5) represents the conditional-variance equation. The effect of a shock in market $j$ (say, the U.S. market) on the volatility of market $i$ (say, the Korean market) is determined by the coefficient of $f_j$, $\alpha_{ij}$. Note that $f_j$ consists of $\left|z_{j,t-1}\right| - E\left(\left|z_{j,t-1}\right|\right)$ and $\delta_j z_{j,t}$. The term $\left|z_{j,t-1}\right| - E\left(\left|z_{j,t-1}\right|\right)$ which is given by the deviation of standardized errors in market $j$ (in absolute value) from their mean measures the size effect of a volatility spillover, and the term $\delta_j z_{j,t}$ measures the sign effect of a volatility spillover. Asymmetry in the spillover effect is present if $\delta_j$ is negative and statistically significant (assuming that a negative shock exerts a larger impact on volatility than a positive shock). If $\delta_j$ is negative and $z_j$ is negative, then the positive value of $\delta_j z_{j,t}$ reinforces the size effect. However, if $\delta_j$ is negative, and $z_j$ is positive, then the negative value of $\delta_j z_{j,t}$ offsets partially the size effect.

[21] More recently, Worthington et al. (2005) and Nam et al. (2008) use this type of a MA term.

The asymmetric spillover effect of a shock in market *j* on the volatility of market *i* (expressed in logs) is measured by

(4.7) $$\frac{\partial \ln \sigma_{i,t}^{2}}{\partial z_{j,t}} = \alpha_{ij} \frac{\partial f_j}{\partial z_{j,t}}$$

It follows from equation (4.6) that

(4.8a) $$\frac{\partial f_j}{\partial z_{j,t}} = 1 + \delta_j \quad \text{for } z_j > 0 \text{ and}$$

(4.8b) $$\frac{\partial f_j}{\partial z_{j,t}} = -1 + \delta_j \text{ for } z_j < 0$$

Thus, the asymmetric effect of a positive shock on market *i*'s volatility ($\ln \sigma_{it}^{2}$) is given by $\alpha_{ij}\left(\left|1+\delta_j\right|\right)$, and the asymmetric effect of a negative shock is given by $\alpha_{ij}\left(\left|-1+\delta_j\right|\right)$. The importance of the asymmetric effect of a negative shock relative to a positive shock or leverage effect is then given by

$$\frac{\left|-1+\delta_j\right|}{\left|1+\delta_j\right|}$$

I can consider three possibilities:

(a) If $\delta_j = 0$, a negative shock has the same effect on volatility as a positive shock.

(b) If $\delta_j < 0$, a negative shock has a larger effect on volatility than a positive shock.

(c) If $\delta_j > 0$, a positive shock has a greater effect on volatility than a negative shock.

Finally, *γ* in equation (4.5) measures the persistence of volatility. If the conditional variance depends on the previous conditional variance, then a GARCH effect is confirmed. If $\gamma$ is less than one, the unconditional variance is finite; if $\gamma$ is equal to one, the unconditional variance does not exist, and the conditional variance follows I(1).

Researchers are typically concerned with a situation in which price and volatility spillovers occur from big markets to relatively small markets, ruling out the possibility of the reverse direction. Although such an assumption may be plausible in light of the fact that the size of the Korean market is small relative to that of the U.S., Chinese, and Japanese markets, that assumption is not necessarily warranted. However, since the spillovers that run from the U.S., Chinese, and Japanese markets to the Korean market is a predominant pattern, the main focus of this study is on the following form of spillovers:[22] (I report the results on price and volatility spillovers from the Korean market to the Chinese and Japanese market, particularly during the crisis period in section 4.5) First, I formulate the model from U.S. market to the Korean market:[23]

(4.9) $$R_{1t} = \beta_{1,0} + \beta_{1,1}\varepsilon_{1,t-1} + \varepsilon_{1,t}$$

(4.10) $$R_{2t} = \beta_{2,0} + \beta_{2,1}\varepsilon_{1,t-1} + \beta_{2,2}\varepsilon_{2,t-1} + \varepsilon_{2,t}$$

where 1 represents the U.S. market, and 2 the Korean market. By the same token, equation (4.5) can be rewritten as

(4.11) $$\sigma_{1t}^2 = \exp\left[\alpha_{1,0} + \alpha_{1,1}f_1\left(z_{1,t-1}\right) + \gamma_1 \ln\left(\sigma_{1,t-1}^2\right)\right]$$

(4.12) $$\sigma_{2t}^2 = \exp\left[\alpha_{2,0} + \alpha_{2,1}f_1\left(z_{1,t-1}\right) + \alpha_{2,2}f_2\left(z_{2,t-1}\right) + \gamma_2 \ln\left(\sigma_{2,t-1}^2\right)\right]$$

Equation (4.6) is also reformulated as

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[23] I report results of the Granger causality which determine endogenous problems in explanatory variables. There is only one direction from the U.S. market to the Korean market.
[23] This model is based on different trading time between the U.S. market and the Korean market.

(4.13) $$f_1\left(z_{1,t-1}\right)=\left(\left|z_{1,t-1}\right|-E\left(\left|z_{1,t-1}\right|\right)\right)+\delta_1 z_{1,t-1}$$

(4.14) $$f_2\left(z_{2,t-1}\right)=\left(\left|z_{2,t-1}\right|-E\left(\left|z_{2,t-1}\right|\right)\right)+\delta_2 z_{2,t-1}$$

I estimate the EGARCH model in two steps. In the first stage, I estimate equation (4.9) (U.S. return equation) and obtain OLS residuals for the U.S. market. I then estimate equation (4.11) (U.S. conditional-variance equation) after I calculate standardized errors and substitute equation (4.13) into equation (4.11). In the second stage, I estimate equation (4.10) (return equation for the Korean market i) and calculate standardized errors for the Korean market *i*. I then estimate equation (4.12) (conditional-variance equation for the Korean market *i*) using standardized errors of the U.S. and the Korean market *i* after I substitute equation (4.13) and (4.14) into equation (4.12). I have computed the values of $\delta_j$ ($j$ = 1, 2) using GAUSS, and standardized errors using the Delta Method.[24]

## 4.4. ANALYSIS OF DATA

The trading time in the United States is from 9:30 a.m. to 4:00 p.m. This trading time corresponds to the time from 11:30 p.m. to 6:00 a.m. (during the day light saving

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[24] I have used E-VIEWS and GAUSS to estimate the set of equations and the Delta Method to estimate the standard errors of $\delta_j$ ($j$ = 1, 2). Letting $\gamma = f\left(\theta\right)$ and $\hat{\gamma} = f\left(\hat{\theta}\right)$, then

$$\operatorname{cov}\left(\hat{\gamma}\right) = F\left(\hat{\theta}\right)\operatorname{cov}\left(\hat{\theta}\right)F\left(\hat{\theta}\right)'$$

where $\hat{\theta} = \begin{bmatrix} a_{1,1} \\ a_{1,1}\delta_1 \end{bmatrix}$ and $cov(\hat{\theta}) = \begin{bmatrix} \operatorname{var}(a_{1,1}) & \operatorname{cov}(a_{1,1}, a_{1,1}\delta_1) \\ \operatorname{cov}(a_{1,1}, a_1\delta_1) & \operatorname{var}(a_{1,1}\delta_1) \end{bmatrix}$

since $f\left(\hat{\theta}\right) = \begin{bmatrix} a_{1,1} \\ \left(a_{1,1}\delta_1\right)/a_{1,1} \end{bmatrix}$, and $F\left(\hat{\theta}\right) = \begin{bmatrix} 1 & 0 \\ -\left(a_{1,1}\delta_1\right)/a_{1,1}^2 & 1/a_{1,1} \end{bmatrix}$

Similarly, I have calculated the standard errors of $a_{2,1}$ using the var-cov matrix of $F\left(\hat{\theta}\right)$.

time period) in Korea. Thus, at the time when the Korean stock market open at 9:00 a.m., information on changes in the prices of stocks traded in the U.S. market on the previous trading day is available to traders in the Korean market, and the arrival of new information can exert some effects on the behavior of traders in the Korean stock market. The Chinese market opens one hour later than the Japanese and Korean markets. Thus, information produced in other markets becomes a part of the information set of traders in the Korean market.

This study uses daily closing prices to calculate the daily returns of the Chinese, Japanese, U.S. markets and the Korean market. The data I have used include the S&P 500 Index (SPX, U.S.), the Shanghai Composite Index (SHCOMP, China), the Nikkei Stock Average Index (NIKKEI, Japan), and the Korea Composite Stock Price Index (KOSPI, Korea). The data are from *http://www.datastream.com*.

The New York Stock Exchange is the largest market in the world by all criteria. The Origin of The New York Stock Exchange started trading on May 17, 1792. It has a total market capitalization over USD 25 trillion (December, 2007). The S&P 500 Index is one of main index in the U.S. All of the stocks in S&P 500 Index are those of large publicly held companies and trade on the two largest U.S. stock markets, the New York Stock Exchange and Nasdaq.

The Shanghai Stock Exchange is the fifth largest in the view of a market capitalization (USD 2.38 trillion, 2008) in the world. It began operation on December 19, 1990. The Shanghai Composite Index was launched July 15, 1991. There are two types of stock being issued. A shares in the Shanghai Composite Index are priced in the domestic currency, while B shares are quoted in U.S. dollars. Initially, trading in A shares are restricted to domestic investors only while B shares are available to both domestic (since 2001) and foreign investors. However, after reforms were implemented in December 2002, foreign investors are now allowed (with limitations) to trade in A shares under the Qualified Foreign Institutional Investor (QFII) system. There is a plan to eventually merge the two types of shares. The data was available from January 3, 1995.

The Tokyo Stock Exchange is the second largest market in the world by market value. It started trading on June 1, 1878. It has a total market capitalization over USD 5 trillion (July, 2007). The Nikkei Stock Average Index is the main index in the Tokyo Stock Exchange.

The Korean Stock Exchange opened in 1953. Since early 1980s, the Korean stock market was gradually open to foreign investors. As the first step, international investment trusts and country funds such as the Korea Fund were launched, thus allowing foreigners to make indirect portfolio investment. In 1992, the Korean stock market was opened to foreign investors with certain restrictions, and the foreign share ownership restrictions were gradually lifted and were fully eliminated in 1998. Additionally, the membership for the Korean Market was opened to foreign securities companies. The Korea Composite Stock Price Index is the main index in the Korean Stock Exchange.

I have broken down the sample into three sub-samples:[25] (1) sample 1: January 3, 1995 -September 30, 1997; (2) sample 2: October 1, 1997 - September 30, 1998, and (3) sample 3: October 1, 1998 - July 31, 2007. The first sample period roughly corresponds to the period prior to the Asian financial crisis. The second sample period falls under the height of the financial crisis. The third sample period deals with a period during which some financial reforms were under way after the financial turmoil.

Table 4.1 presents some basic statistics for the variables used in this study. For sample periods, only the U.S. market shows a right-skewed pattern. In addition, all markets show a leptokurtic distribution. In order to test whether the returns in each market are normally distributed, I have conducted the Jarque-Bera test.[26] The null hypothesis that the returns are normally distributed is rejected for all markets. This finding is broadly consistent with most of previous studies that have tested for the distribution of stock returns.

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[25] For sub-sample periods, I conduct the Chow test (1960) and gain the significance of breaks.

[26] The Jarque-Bera statistic is given by $JB = (n-k)\left[\frac{S^2}{6} + \frac{(k-3)^2}{24}\right]$ where n indicates the number of observations, k the number of parameters estimated, *K* the kurtosis of the distribution, and *S* the skewness of the distribution. The Jarque-Bera statistic follows a chi-square distribution with 2 degrees of freedom under the null hypothesis that the variable is normally distributed.

TABLE 4.1

BASIC STATISTICS FOR STOCK RETURNS

| | S&P 500 | SHCOPM | NIKKEI | KOSPI |
|---|---|---|---|---|
| Mean | 0.000593 | -0.00004 | 0.000352 | 0.000197 |
| Median | 0.000000 | 0.000000 | 0.000322 | 0.000000 |
| SD | 0.017732 | 0.013703 | 0.010486 | 0.01958 |
| Skewness | 0.689536 | -0.036186 | -0.12757 | -0.11902 |
| Kurtosis | 26.96772 | 5.265015 | 6.858876 | 7.055572 |
| Jarque-Bera | 78826.62 | 701.85 | 2043.99 | 2255.59 |
| (Probability) | (0.000000) | (0.000000) | (0.000000) | (0.000000) |
| Sample size | 3280 | 3280 | 3280 | 3280 |

I have also tested whether the return series in each market are white noise using the Ljung-Box test. To this end, I have investigated the autocorrelation of $R_{it}$ and the square of $R_{it}$ for 8, 16, and 24 lags, respectively. The Ljung-Box Q statistic follows a chi-square distribution under the null hypothesis that the series exhibit white noise processes.[27] Table 4.2 and Table 4.3 report the Ljung-Box Q statistics for each market. No autocorrelation is present in the Chinese returns at lag 8, and Japanese returns at lag 16 and 24 but the rest of the return series is serially correlated regardless of the length of lags. I also reject the null hypothesis for the squared returns in all markets including.[28]

[27] The Ljung-Box Q statistic is given by $Q_{LB} = n\left(n+2\right)\sum \rho_k^2 / n - j$, where $\rho_k$ is autocorrelation between $\gamma_t$ and $\gamma_{t-k}$. The Ljung-Box Q statistic is distributed as $\chi^2$ with n degrees of freedom under the null hypothesis that $\rho_1 = \rho_2 = \cdots \rho_k = 0$.

[28] The squared return series can be viewed as a proxy for the variance of the series.

These test results suggest that once volatility gets larger, such large volatility persists for a certain period of time.

TABLE 4.2

LJUNG-BOX STATISTICS FOR RETURNS

| | S&P 500 | SHCOPM | NIKKEI | KOSPI |
|---|---|---|---|---|
| Q(8) | 14.10<br>(0.079) | 7.77<br>(0.456) | 14.05<br>(0.080) | 30.30<br>(0.000) |
| Q(16) | 26.85<br>(0.043) | 27.72<br>(0.034) | 19.08<br>(0.264) | 40.83<br>(0.001) |
| Q(24) | 42.29<br>(0.012) | 46.54<br>(0.004) | 31.41<br>(0.142) | 51.48<br>(0.001) |

TABLE 4.3

LJUNG-BOX STATISTICS FOR THE SQUARES OF RETURNS

| | S&P 500 | SHCOPM | NIKKEI | KOSPI |
|---|---|---|---|---|
| Q(8) | 685.5<br>(0.000) | 489.58<br>(0.000) | 328.35<br>(0.000) | 698.91<br>(0.000) |
| Q(16) | 1063.3<br>(0.000) | 501.68<br>(0.000) | 481.98<br>(0.000) | 1143.9<br>(0.000) |
| Q(24) | 1354.8<br>(0.000) | 514.1<br>(0.000) | 651.75<br>(0.000) | 1540.7<br>(0.000) |

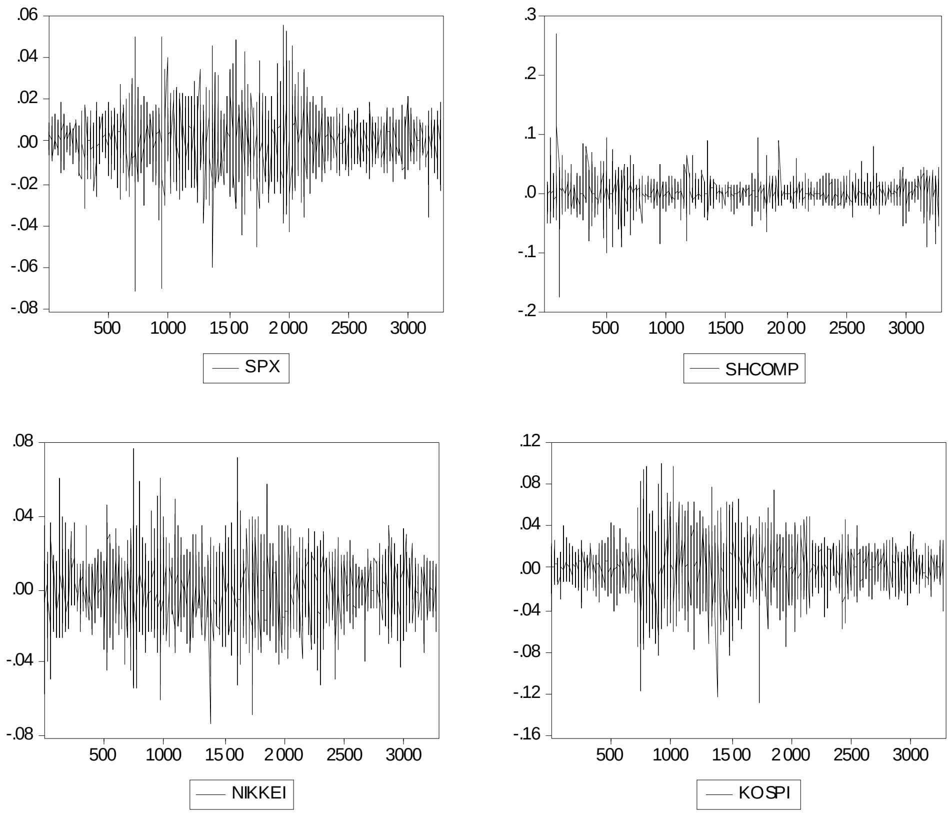


FIGURE 4.1. − Movements in daily returns

Figure 4.1 shows dynamic movements in daily stock returns in each market.[29] As evidenced by the diagram, the daily stock returns of the Korean market showed wide fluctuations during and after the financial crisis. The pattern of successive periods of large volatility followed by successive periods of low volatility is pronounced. Thus, the GARCH appears to be suitable for modeling such volatility clustering.

[30] The origin on the X-axis starts with January 3, 1995. The 1,000th observation corresponds to November 2, 1998, the 2,000th observation to September 2, 2002, and the 3,000th observation to July 3, 2007.

## 4.5. EMPIRICAL RESULTS

In order to examine price and volatility spillovers from the U.S., Chinese, and Japanese markets to the Korean market, I have estimated five sets of the EGARCH model: (1) SPX (U.S.) - KOSPI (Korea), (2) SHCOMP (China) – KOSPI (Korea), (3) NIKKEI (Japan) - KOSPI (Korea). For each set of the EGARCH model, I have estimated 14 coefficients.

### 4.5.1. *GARCH Effects*

The persistence of volatility (GARCH effect) in each market is measured by $\gamma_i$, and the estimated values of $\gamma_i$ for each market are as follows:

| | | | |
|---|---|---|---|
| U.S.: | $0.976^{***}$ (sample 1); | $0.911^{***}$ (sample 2); | $0.986^{***}$ (sample 3) |
| China: | $0.729^{***}$ (sample 1); | $0.441^{***}$ (sample 2); | $0.981^{***}$ (sample 3) |
| Japan: | $0.989^{***}$ (sample 1); | $0.979^{***}$ (sample 2); | $0.973^{***}$ (sample 3) |
| Korea: | $0.935^{***}$ (sample 1); | $0.895^{***}$ (sample 2); | $0.993^{***}$ (sample 3) |

where *** indicates significance at the 1 percent level, ** significance at the 5 percent level, and * significance at the 10 percent level.

The GARCH effect is confirmed in most markets. Thus, there was a tendency for volatility to persist in all markets, which renders support to the GARCH specification. Second, the magnitude of the GARCH effect is more or less of the same order in most markets, ranging from 0.729 to 0.997.

### 4.5.2. *Asymmetric Price and Volatility Spillovers*

The coefficients which pertain to price and volatility spillovers are as follows:[30]

(a) $\beta_{2,1}$ measures the effect of past innovations in the U.S., Chinese, and Japanese markets on the price of the Korean market at *t*.

(b) $\alpha_{2,1}$ determines the overall volatility effect from the U.S., Chinese, and Japanese markets to the Korean market at *t*. It includes the size effect and the asymmetric effect of U.S., Chinese, and Japanese innovations on the volatility of the Korean market. In this section, I evaluate the overall volatility spillover effect ($\alpha_{i,j}$) without separating the asymmetric effect of a positive $\alpha_{ij}\left(\left|1+\delta_j\right|\right)$ or negative shock $\alpha_{ij}\left(\left|-1+\delta_j\right|\right)$ (I will address the leverage effect in Sections 4.3 and 4.4).

As a whole, the U.S. and Japanese markets showed similar pattern in the price and spillover; the Chinese market revealed several distinctive features. First, in the Korean market, the price spillover effect from the U.S. market gained strength for all periods but the volatility spillover effect became much stronger after the crisis. Second, the price spillover effect from the Japanese market to the Korean market became stronger from the crisis period but the volatility spillover effect became much stronger after the crisis. Third, the Chinese market the price spillover effect on the Korean market remained quite small and stable except the crisis period and the volatility effect remained also quite stable on the Korean market for all periods.

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[30] Empirical results among the U.S., Chinese, and Japanese markets are not reported here, but available from the authors upon request.

In this regard, Feenstra, Huang and Hamilton (2003) have presented an interesting proposition. They have noted that Korea has vertically and horizontally-integrated industry groups, which could lead to different responses to different financial shocks. According to Feenstra et al., Korea has some of the largest and most vertically-integrated industry groups (V-groups) and much smaller and concentrated in upstream sectors (U-groups). The responses of industry groups to large external shocks such as the 1997 financial crisis can be different between V-groups and U-groups. They suggest that the equilibria of U-groups are stable, so that a temporary shock will not have permanent effects on markets. However, the equilibria of V-groups are unstable, so that the effects of a competitive shock will be much more severe. My test results give some empirical content to their proposition concerning which different business group will experience financial difficulty in the presence of large shocks.

TABLE 4.4

SPILLOVERS FROM THE U.S. MARKET TO THE KOREAN MARKET

| | Coefficients | z-statistics | Coefficients | z-statistics | Coefficients | z-statistics |
|---|---|---|---|---|---|---|
| $\beta_{1,0}$ | 0.00087 | 3.467*** | -0.00009 | -0.179 | 0.00006 | 0.392 |
| $\beta_{1,1}$ | 0.09447 | 2.323** | -0.00128 | -0.018 | -0.0244 | -1.084 |
| $\alpha_{1,0}$ | -0.29772 | -4.930*** | -0.87174 | -5.258*** | -0.17988 | -8.349*** |
| $\alpha_{1,1}$ | 0.09407 | 4.342*** | 0.10499 | 1.303 | 0.06893 | 6.371*** |
| $\gamma_1$ | 0.97686 | 182.857*** | 0.91125 | 47.086*** | 0.98638 | 539.886*** |
| $\delta_1$ | -0.87712 | -3.202*** | -3.20809 | -1.243 | -1.47627 | -5.549*** |
| $\beta_{2,0}$ | -0.00068 | -1.332 | -0.00308 | -1.269 | 0.00084 | 3.090*** |
| $\beta_{2,1}$ | 0.10578 | 1.669* | 0.53951 | 2.844*** | 0.59187 | 17.982*** |
| $\beta_{2,2}$ | 0.13711 | 3.421*** | 0.10319 | 1.693* | -0.0326 | -1.500 |
| $\alpha_{2,0}$ | -0.72327 | -3.147*** | -9.65491 | -2.460*** | -0.17507 | -7.873*** |
| $\alpha_{2,1}$ | 0.01294 | 0.587 | 0.00774 | 0.360 | 0.04217 | 8.873*** |
| $\alpha_{2,2}$ | 0.1256 | 3.500*** | -0.12567 | -1.035 | 0.10885 | 8.358*** |
| $\gamma_2$ | 0.92953 | 37.831*** | -0.43883 | -0.762 | 0.9881 | 463.768*** |
| $\delta_2$ | -0.65822 | -2.998*** | -0.17668 | -0.320 | -0.01534 | -0.057 |
| Diagnostics on standardized residuals | | | | | | |
| | S&P 500 | KOSPI | S&P 500 | KOSPI | S&P 500 | KOSPI |
| Mean | 0.00849 | -0.00092 | 0.02671 | 0.00018 | -0.00121 | -0.00173 |
| SD | 1.00526 | 1.00148 | 1.01546 | 1.00238 | 1.0003 | 1.00175 |
| Skewness | -0.28812 | 0.03330 | -0.60587 | 0.25127 | -0.36397 | -0.31494 |
| Kurtosis | 4.12239 | 3.65346 | 4.24721 | 3.90447 | 4.6596 | 4.97565 |
| Q(12) | 11.77 | 11.948 | 8.52 | 13.646 | 16.318 | 10.173 |
| | (0.381) | (0.368) | (0.666) | (0.253) | (0.130) | (0.515) |
| Q2 (12) | 9.616 | 10.08 | 5.624 | 22.898 | 10.469 | 14.71 |
| | (0.565) | (0.523) | (0.897) | (0.018) | (0.489) | (0.196) |

Note: *** indicates significance at the 1 percent level, ** significance at the 5 percent level, and * significance at the 10 percent level.

First of all I check price and volatility spillovers from the U.S. market to the Korean market in Table 4.4. For sample 1 (period prior to the crisis), I am not able to reject the null hypotheses that $\alpha_{2,1} = 0$. Thus, no evidence of volatility spillovers from the U.S.

market to the Korean market is found. However, the price spillover effect from the U.S. market was present at the 10% significance of level before the financial crisis. During the financial crisis (sample 2), the spillover effects from the U.S. market to the Korean market are mixed: I have found a positive spillover effect for stock prices, but no evidence is found for volatility spillovers. After the financial crisis (sample 3), the transmission of U.S. shocks to the prices and volatility of the Korean market picked up strongly ($\beta_{2,1}$: 0.106 before crisis $\rightarrow$ 0.592 after crisis; $\alpha_{2,1}$: 0.013 before crisis $\rightarrow$ 0.042 after crisis).

Second, I check price and volatility spillovers from the Chinese market to the Korean market in Table 4.5. I have obtained somewhat different results for the Korean market from those of the Chinese market. For sample 1 (period prior to the crisis), I am able to reject the null hypotheses that $\beta_{2,1} = 0$ and $\alpha_{2,1} = 0$. Thus, no evidence of price and volatility spillovers from the Chinese market to the Korean market is found. During the financial crisis (sample 2), the spillover effects from the Chinese market to the Korean market are mixed: I have found a positive spillover effect for volatility, but no evidence is found for price spillovers. The magnitude of spillover effect remained quite small and stable between the prior- and post-crisis periods ($\beta_{2,1}$: 0.045 before crisis $\rightarrow$ 0.062 after crisis), and the volatility spillover effect also remained stable and significantly after the financial crisis ($\alpha_{2,1}$: 0.056 before crisis $\rightarrow$ 0.048 after crisis).

# TABLE 4.5

## SPILLOVERS FROM THE CHINESE MARKET TO THE KOREAN MARKET

| | Coefficients | z-statistics | Coefficients | z-statistics | Coefficients | z-statistics |
|---|---|---|---|---|---|---|
| $\beta_{1,0}$ | 0.00099 | 0.988 | 0.00031 | 0.439 | 0.00055 | $2.505^{**}$ |
| $\beta_{1,1}$ | 0.00139 | 0.030 | 0.0384 | 0.423 | 0.00218 | 0.112 |
| $\alpha_{1,0}$ | -2.14557 | $-4.259^{***}$ | -5.35707 | $-8.228^{***}$ | -0.25502 | $-10.690^{***}$ |
| $\alpha_{1,1}$ | 0.24365 | $4.933^{***}$ | 0.61352 | $7.948^{***}$ | 0.14064 | $19.667^{***}$ |
| $\gamma_1$ | 0.72943 | $11.224^{***}$ | 0.44174 | $5.977^{***}$ | 0.98189 | $370.440^{***}$ |
| $\delta_1$ | 0.33148 | $3.034^{***}$ | -0.37882 | $-2.931^{***}$ | -0.10430 | $-2.404^{**}$ |
| $\beta_{2,0}$ | -0.00056 | -1.171 | -0.00386 | -1.728 | 0.00091 | $2.821^{***}$ |
| $\beta_{2,1}$ | 0.04536 | $2.742^{***}$ | -0.16912 | -1.151 | 0.06176 | $3.165^{***}$ |
| $\beta_{2,2}$ | 0.13901 | $3.649^{***}$ | 0.09958 | 1.268 | 0.03671 | $1.713^{*}$ |
| $\alpha_{2,0}$ | -0.48116 | $-3.149^{***}$ | -0.85166 | $-1.925^{*}$ | -0.08703 | $-6.509^{***}$ |
| $\alpha_{2,1}$ | 0.05609 | $2.709^{***}$ | 0.09472 | $1.735^{*}$ | 0.04813 | $6.080^{***}$ |
| $\alpha_{2,2}$ | 0.07350 | $2.606^{***}$ | 0.19707 | $2.251^{**}$ | 0.09054 | $9.022^{***}$ |
| $\gamma_2$ | 0.95138 | $58.454^{**}$ | 0.89488 | $14.309^{***}$ | 0.99703 | $821.061^{***}$ |
| $\delta_2$ | -0.89023 | $-2.560^{**}$ | 0.04547 | 0.173 | -0.24062 | $-4.099^{***}$ |
| Diagnostics on standardized residuals | | | | | | |
| | SHCOMP | KOSPI | SHCOMP | KOSPI | SHCOMP | KOSPI |
| Mean | -0.00457 | -0.00308 | 0.02057 | 0.02459 | -0.01903 | -0.00679 |
| SD | 1.00039 | 1.00138 | 1.00110 | 0.99859 | 1.00057 | 1.00367 |
| Skewness | 2.05687 | 0.03550 | 0.17725 | 0.17946 | 0.24534 | -0.33784 |
| Kurtosis | 29.8225 | 3.47876 | 5.19208 | 3.36692 | 7.76199 | 5.14062 |
| Q(12) | 4.892 | 11.001 | 3.496 | 11.345 | 21.056 | 8.060 |
| | (0.936) | (0.443) | (0.982) | (0.415) | (0.033) | (0.708) |
| Q2 (12) | 1.429 | 10.524 | 8.335 | 11.121 | 5.814 | 13.946 |
| | (1.000) | (0.484) | (0.683) | (0.433) | (0.885) | (0.236) |

Note: *** indicates significance at the 1 percent level, ** significance at the 5 percent level, and * significance at the 10 percent level.

TABLE 4.6

SPILLOVERS FROM THE JAPANESE MARKET TO THE KOREAN MARKET

| | Coefficients | z-statistics | Coefficients | z-statistics | Coefficients | z-statistics |
|---|---|---|---|---|---|---|
| $\beta_{1,0}$ | 0.00018 | 0.583 | -0.00185 | -1.785* | 0.0002 | 0.839 |
| $\beta_{1,1}$ | -0.05366 | -1.292 | -0.03538 | -0.607 | -0.01226 | -0.511 |
| $\alpha_{1,0}$ | -0.11149 | -4.963*** | -0.18726 | -2.002** | -0.34032 | -7.202*** |
| $\alpha_{1,1}$ | 0.01795 | 1.348 | 0.0421 | 1.287 | 0.14417 | 9.445*** |
| $\gamma_1$ | 0.98903 | 369.931*** | 0.97997 | 84.062*** | 0.97359 | 206.622*** |
| $\delta_1$ | -5.30875 | -1.338 | -4.00542 | -1.170 | -0.41826 | -5.798*** |
| $\beta_{2,0}$ | -0.00072 | -1.405 | -0.00386 | -1.728* | 0.00097 | 3.899*** |
| $\beta_{2,1}$ | -0.01695 | -0.592 | 0.38123 | 4.106** | 0.63316 | 31.108*** |
| $\beta_{2,2}$ | 0.14401 | 3.553*** | 0.04457 | 0.568 | 0.02409 | 1.120 |
| $\alpha_{2,0}$ | -0.87011 | -3.244*** | -0.70535 | -2.549** | -0.13315 | -8.272*** |
| $\alpha_{2,1}$ | 0.00377 | 1.203 | -0.01852 | -1.194 | 0.05249 | 7.067*** |
| $\alpha_{2,2}$ | 0.13419 | 3.456*** | 0.19613 | 2.336** | 0.11045 | 8.816*** |
| $\gamma_2$ | 0.91378 | 31.397*** | 0.91795 | 24.500*** | 0.99404 | 722.527*** |
| $\delta_2$ | -0.65436 | -3.180*** | -0.15358 | -0.533 | -0.01956 | -0.342 |
| Diagnostics on standardized residuals | | | | | | |
| | SHCOMP | KOSPI | SHCOMP | KOSPI | SHCOMP | KOSPI |
| Mean | -0.03979 | -0.00308 | 0.02073 | 0.03412 | -0.0077 | -0.0094 |
| SD | 1.01672 | 1.00138 | 1.00932 | 0.99934 | 1.00109 | 1.00319 |
| Skewness | -0.20098 | 0.0355 | -0.04959 | 0.11551 | -0.17651 | -0.08972 |
| Kurtosis | 7.41934 | 3.47876 | 3.309 | 3.62189 | 4.27702 | 4.10654 |
| Q(12) | 11.918 | 12.808 | 11.499 | 10.125 | 7.428 | 10.191 |
| | (0.370) | (0.306) | (0.402) | (0.519) | (0.763) | (0.513) |
| Q2 (12) | 6.617 | 9.074 | 9.274 | 10.557 | 29.197 | 18.07 |
| | (0.862) | (0.615) | (0.597) | (0.481) | (0.002) | (0.080) |

Note: *** indicates significance at the 1 percent level, ** significance at the 5 percent level, and * significance at the 10 percent level.

Third, I check price and volatility spillovers from the Japanese market to the Korean market in Table 4.6. I have obtained somewhat similar results for the Korean market from those of the U.S. market in case of volatility spillovers. For sample 1 (period prior to the crisis), I am not able to reject the null hypotheses that $\beta_{2,1} = 0$ and $\alpha_{2,1} = 0$. Thus, no evidence of price and volatility spillovers from the Japanese market to the Korean market is found. During the financial crisis (sample 2), the spillover effects from the Japanese market to the Korean market are mixed: I have found a positive spillover effect for stock prices, but no evidence is found for volatility spillovers. After the financial crisis (sample 3), the transmission of Japanese shocks to the prices and volatility of the Korean market became strongly ($\beta_{2,1}$: -0.017 before crisis $\rightarrow$ 0.633 after crisis; $\alpha_{2,1}$: 0.004 before crisis $\rightarrow$ 0.053 after crisis). The Japanese market is the very similar to U.S. market where volatility spillover effect from the Japanese market was present after the financial crisis.

Fourth, I check price and volatility spillovers from the Korean market to the U.S. market. Figure 4.1 reveals that four markets also seemed to be subjected to above-average volatility during the 1997 Asian financial crisis. It is possible to conjecture that the transmission of prices and volatility from the Korean market to the U.S., Chinese, and Japanese markets occurred during the crisis period. First, I test exogeneity of the regressors by the Granger causality procedure.

Results from the Granger causality tests clearly support one-way causation from the U.S. market to the Korean market. This supports the assumption of weak exogeneity in the U.S. market relative to the Korean market, and limited feedback from the Korean market to U.S., market. For the U.S. market and the Korean market, the U.S. market explains movements in the Korean market, but not vice-versa. However, I have different result for the Korean market from that of the Chinese market and the Japanese market.

Granger causality tests show no causation from the Chinese market to the Korean market. In addition, there is mixed causation between the Japanese market and the Korean market. There is an endogenous problem from the Japanese market to the Korean market.

U.S. not cause Korea: $3.762^{*}$ (sample 1);$9.812^{***}$ (sample 2);$291.995^{***}$ (sample 3)

Korea not cause U.S. : 0.435 (sample 1); 0.485 (sample 2);0.018 (sample 3)

China not cause Korea: 0.050 (sample 1); 0.044 (sample 2);0.439 (sample 3)

Korea not cause China: 0.640 (sample 1); 0.239 (sample 2);2.031 (sample 3)

Japan not cause Korea: 0.898 (sample 1); $4.525^{**}$ (sample 2);0.464 (sample 3)

Korea not cause Japan: 0.872 (sample 1); 1.108 (sample 2);$12.152^{***}$ (sample 3)

where *** indicates significance at the 1 percent level, ** significance at the 5 percent level and * significance at the 10 percent level.

Interestingly enough, I have found that there was only a significant price spillover from the Korean market to the Chinese market prior- and post-crisis periods. After the financial crisis (sample 3), the transmission of Korean shocks to volatility exists. The magnitude of the price spillover effect declined significantly and is small between the prior- and post-crisis periods and the volatility spillover effect also became meager. I also have found that there was only a significant price spillover from the Korean market to the Japanese market from the Asian financial crisis period. Although I accept the result of Granger test after the crisis period, the transmission of Korean shocks to the prices and volatility of the Japanese market increases.

Korea $\rightarrow$ China:

$\alpha_{i,j}$ : -0.016 (sample 1), 0.394$^{***}$(sample 2), - 0.010$^{**}$(sample 3)

$\beta_{i,j}$ : 0.174$^{**}$, 0.006(sample 2), 0.022$^{**}$(sample 3)

Korea $\rightarrow$ Japan:

$\alpha_{i,j}$ : 0.025 (sample 1), 0.054(sample 2), 0.058$^{***}$(sample 3)

$\beta_{i,j}$ : 0.001 (sample 1), 0.106$^{***}$(sample 2), 0.339$^{+}$(sample 3)

where *** indicates significance at the 1 percent level, ** significance at the 5 percent level and * significance at the 10 percent level, and + denotes significant but has a endogenous problem.

Thus, I can argue that the Chinese and Japanese markets were affected to some extent by the Korean financial shock during the crisis period.

4.5.3. *Own Asymmetric Volatility and the Leverage Effect*

In this section I discuss how shocks occurred in each market affect its own volatility. An own asymmetric volatility effect is measured by $\alpha_{i,i}$ and $\delta_i$

$$\frac{\partial \ln \sigma_{it}^2}{\partial z_{it}} = \alpha_{it} \frac{\partial f_i}{\partial z_{it}}$$

$$= \alpha_{ii} \left( \left| 1 + \delta_i \right| \right) \text{ for a positive shock}$$

$$= \alpha_{ii} \left( \left| -1 + \delta_i \right| \right) \text{ for a negative shock}$$

(a) $\alpha_{1,1}$ measures the effect of past innovations originated in the U.S., Chinese, and Japanese markets on the volatility of the U.S., Chinese, and Japanese markets market at *t*, and $\alpha_{2,2}$ measures the effect of past innovations originated in the Korean market on its own volatility at *t*.

(b) $\delta_1$ determines the own asymmetric effect of a shock on market volatility in the U.S., Chinese, and Japanese markets, and $\delta_2$ determines the own asymmetric effect of a shock on market volatility in the Korean market.

As I discussed, if $\delta_i < 0$, a negative shock has a greater effect on market volatility than a positive shock. If $\delta_i > 0$, a positive shock has a larger effect on market volatility than a negative shock. Thus, if $\alpha_{1,1}$ is significant, and $\delta_1$ is significant and negative, asymmetry in volatility exists in the U.S., Chinese, and Japanese markets. Similarly, if $\alpha_{2,2}$ is significant, and $\delta_2$ is significant and negative, asymmetry in volatility is present in the Korean market under consideration. The $\delta_i$ coefficient is negative in all markets with only one exception that occurred before the period of the financial crisis: Chinese $\delta_i$ is 0.331 which is significant at the 5 percent level, and $\delta_i$ from the Chinese market to the Korean market in the sample period 2 is 0.045 which is insignificant at any reasonable level of significance. Thus, the dominant evidence shows that bad news in each market could have a greater impact on its own market volatility than good news. Also, there is one episode in which $\delta_i$ is significant and smaller than - 1 ($\delta_i < -1$): The estimate of $\delta_i$ in the U.S. market was - 1.476 during the sample period 3. In all other markets and samples, $\delta_i$ lies between - 1 and zero ($-1 < \delta_i < 0$).

I am particularly concerned with the leverage effect which is given by $|-1+\delta_i|/|1+\delta_i|$. It measures how large the effect of a negative shock on volatility is relative to the effect of a positive shock. For example, if the size of the leverage effect is 2, then the effect of a negative shock on volatility is twice as large as the effect of a positive shock on volatility. The estimated leverage effect for each market is presented in Table 4.7.

TABLE 4.7

OWN LEVERAGE EFFECT

| | Before Crisis (Period 1) | Crisis (Period 2) | After Crisis (Period 3) |
|---|---|---|---|
| S&P 500 | 15.27578## | 1.90576 | 5.19932## |
| SHCOMP | 0.50208## | 2.21965## | 1.23287## |
| NIKKEI | 1.46417 | 1.66547 | 2.43797## |
| KOSPI | 5.23795## | 0.97143** | 1.23244## |

Note: **: Only $\alpha_{jj}$ is significant at the 5 percent level.
##: Both $\alpha_{jj}$ and $\delta_j$ are significant at the 5 percent level.

The own leverage effect in each market tends to have tapered off substantially after the Asian financial crisis only except for the U.S. and Korean markets: China: 0.502 (before the crisis) $\rightarrow$ 1.233 (after the crisis); Japan: 1.464 $\rightarrow$ 2.438. However, a significant own leverage effect where both $\alpha_{2,2}$ and $\delta_2$ are significant at the 5 percent level was present in the U.S., Chinese, Japanese markets before the crisis, but after the crisis, a significant leverage effect showed up in all the markets.

4.5.4. *Asymmetric Volatility Spillovers from the U.S., Chinese, and Japanese Markets to the Korean Market: Cross Leverage Effect*

Finally, I have figured out the asymmetric spillover effect of a shock originated in the U.S., Chinese, and Japanese markets on the conditional volatility of the Korean market based on my empirical results. The coefficients which pertain to such asymmetric volatility spillovers are $\alpha_{2,1}$ and $\delta_1$. Following equation (4.7), I can calculate the magnitude of the spillover effect of good news (1% market advances) and bad news (1% market declines) from the U.S. market on each market's volatility as follows:

Positive Shock

(4.15) $$\frac{\partial \ln \sigma_{2t}^2}{\partial z_{1t}} = \alpha_2 \left( \left| 1 + \delta_1 \right| \right)$$

Negative Shock

(4.16) $$\frac{\partial \ln \sigma_{2t}^2}{\partial z_{1t}} = \alpha_2 \left( \left| -1 + \delta_1 \right| \right)$$

TABLE 4.8

EFFECT OF + 1% INNOVATIONS IN U.S., CHINESE, AND JAPANESE MARKETS ON THE VOLATILITY OF THE KOREAN MARKET (IN %)

| | Before Crisis (Period 1) | Crisis (Period 2) | After Crisis (Period 3) |
|---|---|---|---|
| S&P 500 | 0.00795 | 0.08548 | 0.10043## |
| SHCOMP | 0.37342## | 0.29421 | 0.21555## |
| NIKKEI | 0.08122 | 0.27825 | 0.15267## |

Note: **: Only the coefficient of volatility spillovers ($\alpha_{2,1}$)is significant at the 5 percent level.
##: The coefficients of both volatility spillovers ($\alpha_{2,1}$) and asymmetry ($\delta_1$) are significant at the 5 percent level.

TABLE 4.9

EFFECT OF - 1% INNOVATIONS IN U.S., CHINESE, AND JAPANESE MARKETS ON THE VOLATILITY OF THE KOREAN MARKET (IN %)

| | Before Crisis (Period 1) | Crisis (Period 2) | After Crisis (Period 3) |
|---|---|---|---|
| S&P 500 | 0.1215 | 0.1629 | 0.52218## |
| SHCOMP | 0.18749## | 0.65305 | 0.26575## |
| NIKKEI | 0.11892 | 0.46342 | 0.37222## |

Note: **: Only the coefficient of volatility spillovers ($\alpha_{2,1}$)is significant at the 5 percent level.
##: The coefficients of both volatility spillovers ($\alpha_{2,1}$) and asymmetry ($\delta_1$) are significant at the 5 percent level.

In Table 4.8 and Table 4.9 report the effect of Innovations in U.S., Chinese, and Japanese markets on the volatility of the Korean market. First, I note that the spillover effect of a negative shock (market declines) from the U.S. and Japanese markets outweighed the spillover effect of a positive shock (market advances). This asymmetric spillover effect appears strongly after the financial crisis. In addition, the spillover effect of a positive shock (market advances) from the Chinese market is much larger than the spillover effect of a negative shock (market declines) before the financial crisis but a negative innovation is slightly larger than a positive shock after the financial crisis.

Second, the negatively asymmetric spillover effect increased from the Chinese market to the Korean market before and after the crisis and from the U.S. and Japanese markets to the Korean market after the crisis. For example, a - 1% shock in the Chinese market increased the conditional volatility of the Korean market from 0.187% to 0.266%, while a - 1% shock in the U.S. market decreased the conditional volatility of the Korean market from 0.122% to 0.526%. These findings indicate that the Korean market became most vulnerable to negative shocks generated in the U.S., Chinese, and Japanese markets after the financial crisis. Third, the positively asymmetric spillover effect increased from all big markets except the Chinese market to the Korean market after the crisis and from the U.S. and Japanese markets to the Korean market after the crisis.

It is also worthwhile noting that when the asymmetric effect of a shock originated in big markets on its own volatility (own leverage effect) is significantly strong, then the asymmetric spillover effect of a big market shock on the Korean market's volatility (cross leverage effect) appears, and vice versa. (This phenomenon does not occur only in the U.S. market before the financial crisis.) For example, before the crisis, a significant own leverage effect was present in the Chinese market and then a significant cross leverage effect was present in the Korean market. After the crisis, a significant own leverage effect was present in all markets and then a significant cross leverage effect was present in the Korean market. Thus, the effect of a domestic shock (big markets' shock) on market volatility increases when the effect of a foreign shock on market volatility dominates. Conversely, when the effect of a foreign shock (big markets' shock) gains ground, then the effect of a domestic shock on market volatility loses strength.

## 4.6. SUMMARY AND CONCLUDING REMARKS

The main purpose of this study has been to investigate price and volatility spillovers from the U.S., Chinese, and Japanese markets to the Korean market before and after the 1997 Asian financial crisis. The existing literature on the effects of the 1997 financial crisis from U.S., Japan, and Chinese to small country is few and far between, despite the fact that the importance and influence of these markets in the world financial market have continued to increase.

The Korean market that was hit most hard by the financial turmoil underwent major changes in the behavior of price and volatility spillovers from the U.S. and the Japanese markets. Interestingly enough, these two markets showed similar patterns of transmission in volatility spillovers after the financial crisis. Especially, the price spillover effect from the U.S. market was present significantly for all periods but the price spillover effect from the Japanese market exist significantly from the financial crisis. On the other hand, the magnitude of spillover effect from the Chinese market to the Korean market remained quite small and stable between the prior- and post-crisis periods and the volatility spillover effect also remained stable significantly between the prior- and post-crisis periods.

Price and volatility spillover effects from the U.S. and the Japanese markets showed up most strongly in the Korean market after the crisis. This finding supports the Feenstra, Huang and Hamilton's (2003) proposition that the effect of an external shock will be much more severe in V-groups. Several factors might coalesce for such shifts. First, the Korean market was most closed in the region before the crisis, but the Korean government took a series of drastic actions to remove many restrictions on capital transactions in the wake of the financial crisis. The consequence of such actions was massive inflows of foreign funds into the Korean market. Currently foreigners' portfolio investment accounts for more than 50 percent of the market value of stocks (10 largest business groups) traded in the Korean market. In addition to the financial factors, the real sector of the Korean economy is heavily dependent on the United States and Japan with the United States and Japan being the largest trading partners of Korea.

The Korean market also different from the U.S. and the Japanese markets as far as the transmission of shocks originated in the Chinese market are concerned. The price spillover effect existed except during the financial crisis and the volatility spillover effect was evident in all samples. This result is also consistent with Feenstra, Huang and Hamilton's conjecture that an external shock will have strong effects on markets for U-groups. Here are two explanations: One possible explanation for the presence of the volatility spillover effect from the Chinese market is that the Korean economy has operated small- and medium-sized firms and that the interdependence of the Korean economy with the Chinese economy is that strong. The other is the market openness. The Chinese market has opened to foreign investors gradually from 2002. Therefore, the spillover effect could be remained stable and small in my sample periods.

To sum up, new information on stock prices originated in the U.S. market was more rapidly and continuously transmitted to the Korean market for all periods, but the transmission of volatility from the U.S. and Japanese markets to the Korean market was also considerably increased after the crisis. The price spillover effect from the Japanese market to the Korean market became stronger from the crisis period. The Chinese market the price spillover effect on the Korean market remained quite stable except the crisis period and the volatility effect remained also quite stable on the Korean market for all periods. Asymmetry in the spillover effect of U.S., China, and Japan shocks on market volatility was pronounced in the Korean market after the financial crisis.

# CHAPTER V

# CONCLUSIONS

First of all, I investigate a relationship between inflation volatility and stock returns. Expected real stock returns are determined by market fundamentals such as the expected values of inflation, money growth, real output growth, and monetary and real shocks. The relation between real stock returns and expected inflation can be of either sign depending on the degree of inflation volatility. My empirical analysis suggests that real stock returns are negatively related to expected inflation in periods of low volatility of inflation and positively related to expected inflation in periods of high volatility of inflation.

In order to test for empirical analysis, I have employed an GARCH(1,1)-M model and conducted an empirical investigation using quarterly data for 16 countries. The data set includes 13 stable-price countries and three volatile-price countries. Empirical results have confirmed that the relationship between real stock returns and expected inflation was negative in all stable-price countries and positive in two volatile-price countries. The only exception was Israel where the volatility of inflation measured by the standard deviation of the inflation rate was relatively high during the sample period, but a negative relation between real stock returns and expected inflation was found, albeit the coefficient was not significant.

Second, I investigate rational bubbles in asset prices. In this study, I formulate an information error model which allows one to derive the measure of non-fundamentals in sotkc prices in a straightforward manner. This study also provides a new method of revealing preliminary evidence that there is the increasing bursting rate at a decreasing rate for extraneous or instrinsic bubbles as the Weibull distribution. This study is the first attempt to apply the Weibull distribution to the test of rational bubbles.

As a result, my empirical analysis As a result, my empirical analysis is the first step in applying survival analysis to bubbles, and it reveals preliminary evidence that there is the increasing bursting rate at a decreasing rate for extraneous or instrinsic bubbles in the U.S. stock market. I have also conducted two different cointegration tests for market efficiency. In the Granger hypothesis I have accepted the null hypothesis of no linear cointegration between stock prices and economic variables. In the Campbell-Shiller's hypothesis, there are no cointegration relationships between the dividend-price ratio and fundamentals except long-term interest.

Third, I investigate price and volatility spillover effects from the U.S., Chinese, and Japanese markets. The main purpose of this study has been to investigate price and volatility spillovers from the U.S., Chinese, and Japanese markets to the Korean market before and after the 1997 Asian financial crisis. The existing literature on the effects of the 1997 financial crisis from U.S., Japan, and Chinese to small country is few and far between, despite the fact that the importance and influence of these markets in the world financial market have continued to increase.

New information on stock prices originated in the U.S. market was more rapidly and continuously transmitted to the Korean market for all periods, but the transmission of volatility from the U.S. and Japanese markets to the Korean market was also considerably increased after the crisis. The price spillover effect from the Japanese market to the Korean market became stronger from the crisis period. The Chinese market the price spillover effect on the Korean market remained quite stable except the crisis period and the volatility effect remained also quite stable on the Korean market for all periods. Asymmetry in the spillover effect of U.S., China, and Japan shocks on market volatility was pronounced in the Korean market after the financial crisis.

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# APPENDIX A

## THE SOURCE OF STOCK RETURNS IN *INTERNATIONAL FINANCIAL STATISTICS* (*IFS*)

| Countries | Source |
|---|---|
| AUSTRALIA | Australian Stock Exchange (ASX), base December 31, 1979. Through March 2000, data refer to the All Ordinaries Index. Beginning in April 2000, index refers to the S&P/ASX 200. |
| CANADA | The Toronto Stock Exchange for a composite of 300 shares, base 1975. |
| FRANCE | The index covers the common shares of the 40 enterprises having the largest capitalization. Price data refer to averages of end-of-week quotations for each month. Prior to 1987, the index was calculated from the sample of 180 shares on the Paris exchange. |
| ISRAEL | December 2000, refers to quotations on the 23rd of each month and covers all ordinary shares quoted on the Tel Aviv Exchange. |
| ITALY | Data refer to the MIB index calculated by the Milan Stock Exchange and are based on the quoted prices of all stocks traded on that exchange. |
| JAPAN | The index, base January 4, 1968, refers to the average of daily closing prices for all shares listed on the Tokyo exchange. |
| KOREA | Beginning 1983, comprises stock prices weighted by total market values. Prior to 1983, the Dow-Jones Average Index is used. |
| MEXICO | General share price index covering shares quoted on the Mexico City Stock Exchange, base October 1, 1978. |
| NETHERLAND | The index, base 1985, comprised a sample of 127 shares. |
| NEW ZEALAND | Beginning in June 1986, gross index calculated by the New Zealand Stock Exchange, which has a base of June 30, 1986 = 100. All shares of all public companies listed on the New Zealand Stock Exchange are contained within the index. |
| NORWAY | The index refers to midmonth prices of manufacturing and mining shares quoted on the Oslo Exchange. |
| PERU | General share price index covering industrial and mining shares quoted in the Lima Stock Exchange, base December 1991. |
| PHILIPPINES | Index of the Manila Stock Exchange on commercial and industrial shares, base 1965. Beginning in December 1972, stock price index of the Manila and Makati stock exchanges, base 1972. Beginning in January 1978, stock price index of the Manila and Makati stock exchanges, base 1985. Beginning in April 1994, stock price index of the Philippine Stock Exchange, base 1985. |
| SPAIN | Index of Madrid Stock Exchange share prices, base December 1970. Beginning January 1986, data refer to base December 1985. |
| UK | Data refer to the average of daily quotations of 500 industrial ordinary shares, base 1985. |
| US | Price-weighted monthly average covering 30 blue chip stocks quoted in the Dow Jones Industrial Average (DJIA). The NASDAQ Composite Index (base February 5, 1971) is a market capitalization-weighted index covering domestic and international-based common stocks, ordinary shares, American Depository Receipts (ADRs), shares of beneficial interest, REITs, Tracking Stocks and Limited Partnerships and excluding exchange traded funds, structured products, convertible debentures, rights, units, warrants and preferred issues. The S&P Industrials (base 41-43=10) is a Laspeyres-type index based on daily closing quotations for companies in the Industrials on the New York Exchange. The AMEX Average (base August 31, 1973) is a total-market-value-weighted index that covers all common shares, warrants, and (ADRs) listed. |

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