Summer 1999

The fundamental and non-fundamental components of stock prices: The role of time-varying expected inflation

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THE FUNDAMENTAL AND NON-FUNDAMENTAL COMPONENTS OF STOCK PRICES: THE ROLE OF TIME-VARYING EXPECTED INFLATION

by

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A Dissertation Presented in Partial Fulfillment
of the Requirements for the Degree
Doctor of Business Administration

COLLEGE OF ADMINISTRATION AND BUSINESS
LOUISIANA TECH UNIVERSITY

August, 1999

UMI Number: 9935114

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LOUISIANA TECH UNIVERSITY

THE GRADUATE SCHOOL

August 14, 1999
Date

We hereby recommend that the dissertation prepared under our supervision

by Maosen Zhong

entitled Fundamental and Non-Fundamental Components of Stock Prices: The Role of Time-Varying Expected Inflation

be accepted in partial fulfillment of the requirements for the Degree of

Doctor of Business Administration

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Advisory Committee

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Dean of Graduate School

GS Form 13
8/96

ABSTRACT

I derive testable implications of fundamental and non-fundamental components of stock prices. In order to control for the role of time-varying expected inflation and to be able to perform reasonable empirical tests, I use a nominal (rather than a real) interpretation of the present-value model (PVM), whereby nominal interest rates approximate expected inflation. I conjecture that the fundamental and non-fundamental components represent the permanent and temporary components of stock prices, respectively. A series of cointegration analysis over the annual period 1871-1997 confirms my conjecture for the model with time-varying expected inflation. Various fundamental and non-fundamental exclusion tests indicate that both excess returns and expected inflation are price fundamentals. When both of these factors are present in the fundamental component, the non-fundamental component of stock prices exhibits little deviation from zero. However, the evidence in support of the inflation-augmented PVM seems somewhat sensitive to certain model specifications (notably, lag structures). The Hansen-Johansen recursive analysis reveals that the parameters in the non-fundamental component lack stability in the post-World War II period. Results from subsample analysis verify my suspicion of a significant regime shift. In particular, the inflation augmented-PVM holds only for the pre-WW II period. This implication of excess price volatility, as represented by the augmented PVM, stands up to alternative specifications such as measurements of variables and data frequency. Such evidence is clearly in line with

Shiller's (1981) belief in market irrationality and also consistent with Campbell's (1991) conclusion that evidence of market predictability is "overwhelming" only during the post-1950s.

DEDICATION

I would like to dedicate this dissertation to many wonderful people in my life, whose support–spiritually and physically–are more than I can possibly acknowledge. First, to my mother, Zhao Liangyu, and my father, Zhong Canzeng, who have instilled in me immeasurable love, support, and inspiration that made this accomplishment possible. To my grandparents whose expectations and encouragement always motivated me to work hard. To my many aunts, uncles, and cousins, whose care and support have made my life a lot more enriched. To Dr. Ali Darrat, my admirable academic mentor, who has been an excellent model for me, both in personality and in intellect. To my dissertation committee–Dr. Dwight Anderson (Chairman), Dr. Ali Darrat and Dr. Otis Gilley–who have provided me with invaluable criticisms, comments, and suggestions regarding the contents and exposition of this dissertation. Finally, to Dr. Marc Chopin, Maxwell Hsu, Thanomsak Suwannoi, and many other respected friends, teachers, and colleagues for their advice and support.

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CHAPTER 1

INTRODUCTION

Recent evidence that the present value model understates volatility in stock prices has ignited interest in possible market irrationality and stimulated further research in behavioral finance. Shiller (1981), for example, argues that actual stock prices are too volatile to be compatible with changes in dividends. Shiller's results are critically built on two main assumptions: a) that real (inflation-adjusted) dividends are stationary around a historical trend; and b) the real expected rate of return is constant. Subsequent research challenges both assumptions and finds them largely responsible for Shiller's conclusion.

For example, Kleidon (1986) and Marsh and Merton (1986) argue against the stationary assumption and provide an alternative dividend-smoothing process whereby dividends become less volatile than stock prices. Responding to this criticism, Campbell and Shiller (1987) incorporate the stationarity requirement as well as any possible cointegration between stock prices and dividends, but still report persistent deviations from the rational behavior implied by the present value model (PVM).

As to Shiller's second assumption, recent finance literature [e.g., French *et al.* (1987) and Fama (1991)] suggests that a time-varying discount rate could, at least partly, explain the observed variability of stock prices. Reacting to this possibility, Campbell and Shiller (1988a) outline a log-linear model allowing for the role of a time-varying discount rate. With alternative measures of time-varying *real* discount rates, they continue to report evidence of excess market volatility, arguing that there remain some "unexplained factors" in the determination of the dividend-price ratio. In their attempt to identify these "unexplained factors," researchers have pursued two main directions. The first is to evaluate different types of discount rates and examine their relationship with the economy [see, for example, Kandel and Stambaugh (1990), Abel (1993), Cecchetti, *et al.* (1990, 1993), and Campbell and Cochrane (1999)]. Other analysts, who believe in some degree of market irrationality, focus instead on investors' sentiment or psychology [see, for example, the "fads" model of Summers (1986), the "self-attribution" model of Daniel *et al.* (1998), and the "representativeness heuristic" model of Barberis, *et al.* (1998). See also Shiller (1998) for an insightful survey of the behavioral finance literature].

Clearly, behavioral models of stock prices are inconsistent with stock-market rationality and with any model that propagates it, including the PVM. If stock prices do deviate from the fundamental value predicted by the PVM, then the non-fundamental component of stock prices should be non-zero. In this study, I propose a new procedure to test whether stock prices have a significant non-fundamental component.

I first model stock prices as the sum of fundamental and non-fundamental components. The fundamental component is derived, in the context of Campbell and Shiller's (1988a) dividend ratio model, from the log-linear version of the PVM. The non-fundamental component of stock prices, being the difference between actual prices and their fundamental value, is any component unaccounted for by price fundamentals [such as the 'fads' component of Summers (1986) and Porterba and Summers (1988) or "rational bubbles" component of Blanchard and Watson (1982) and West (1988a)]. I then propose a likelihood ratio to test the statistical significance of the non-fundamental component of stock prices.

My testing methodology improves over Campbell and Shiller's (1988a) in several respects. In particular, they focus on the dividend-price ratio model (instead of the prices series itself) since the dividend-ratio model is independent of the deflator used to generate real variables and also because the ratio of the two variables is assumed stationary with a (1, -1) cointegrating relationship. By contrast, my test, while incorporating both desirable features, focuses on the prices themselves. Focusing on the dividend ratio in testing price volatility may be inappropriate. For example, rejecting the equality between the "rational dividend ratio" and the actual dividend ratio could be the outcome of dividend smoothing suggested by Marsh and Merton (1986) rather than the result of excess price volatility, *per se*. My test also uses the nominal, as opposed to the traditional real, version of the PVM to analyze price fundamental. In Chapter 3, I discuss five reasons for abandoning this tradition.

I model nominal stock prices, assumed to follow a non-stationary process, as the sum of fundamental and non-fundamental components. I show that these two components correspond respectively to the non-stationary and stationary components in the Gonzalo and Granger (1995)–hereafter GG--sense. In this context, it is possible to directly test whether the non-fundamental component of stock prices is significantly different from zero. Campbell and Shiller (1988a), on the other hand, only compare the volatility of the actual dividend-price ratio and the volatility of the forecasted present value of the real dividend growth rates. As such they do not explicitly decompose the prices or dividend-ratio into their fundamental and non-fundamental components and test the existence of non-fundamental component. Lee (1998) does decompose stock prices into several components, but without formally testing whether the non-fundamental component of stock prices is statistically significant. Besides this apparent weakness of Lee's model, his results are also derived from the Blanchard and Quah's (1989) approach which may not be appropriate, especially in a multivariate context [see Enders (1995, pp. 341-342)].

My nominal version of the PVM allows for the incorporation of an inflation premium into the fundamental value of stock prices, whereas the real version used in previous studies does not. It has been shown early in literature that the returns on stocks and bonds vary with expected inflation rates [Bodie (1976), Jaffe and Mandelker (1976), Fama and Schwert (1977)]. Furthermore, Fama (1991) argues that expected inflation may be one of the factors that account for time-varying expected returns which induce the excess price volatility. I hypothesize that an inflation premium, which has thusfar been ignored in testing PVM, could be one of the "unexplained factors" sought by Campbell and Shiller (1988a) to resolve the volatility issue of the PVM. The results I obtain provide strong evidence supportive of my hypotheses, but only during the pre-World War II period. However, for the post-World War II period, the incorporation of an inflation premium fails to rescue the inflation-augmented PVM, and stock prices continue to show significant deviations from the fundamental values. This apparent rejection of PVM in the post-World War II period stands up to alternative definitions of the variables and data horizons. Clearly, this evidence is supportive of Campbell's (1991) conclusion of stock-market predictability in the post-1950s period, and provides further credence to Shiller's thesis of "market irrationality".

Section 1: Statement of Problem

Fama (1991) argues that the documented excess volatility of stock prices may be due to time-varying discount rates expected by perfectly rational investors. The required rate of return can be decomposed into three factors: real interest rates, risk premiums for the excess risk of stocks compared to riskless short-term debt, and expected inflation. Real interest rates have been stable over time [Fama and Gibbons (1982)] and are found to explain very little variation of stock prices [Lee (1998)]. Time-varying expected excess returns are demonstrated to explain significant variation of stock prices but not all the variation [Campbell and Amber (1993) and Lee (1998)]. Nevertheless, a significant portion of the variance of stock prices is explained by a non-fundamental component [Lee (1998)], representing some degree of market inefficiency.

Here, I raise an interesting question: does the time–varying inflation premium play an important role in the present-value relation? To my best knowledge, the time-varying inflation premium has not been incorporated in testing the excess volatility of stock prices. This is perhaps because Shiller (1981) and LeRoy and Porter (1981) started the variance-bounds literature with real variables, assuming inflation is eliminated from both sides of the present-value equation. The real version of the present value model can avoid the trouble associated with non-stationarity [Campbell and Shiller (1988a)]. However, as Shiller and Beltratti (1992) note, both the nominal and real interpretations of the present value model are equally plausible, and either interpretation can be adopted for empirical convenience (most importantly, satisfying stationarity requirements). In the present study, I extend Campbell and Shiller's (1988a) dynamic Gordon model from a real version into a nominal version that allows for the role of expected inflation [see Chapter 3]. It is interesting to see whether the time-varying inflation premium plays a significant role along with the excess stock returns.

Another question of interest is whether the non-fundamental component of stock prices is significant. Due to econometric limitations, previous studies have not been able to estimate the non-fundamental component of stock prices and formally test its significance. I demonstrate that the non-fundamental component is a cointegrating vector of prices and some fundamental variables [e.g. dividends, expected inflation, and excess stock returns]. Then, I propose a chi-squared test procedure for testing the significance of the non-fundamental component of stock prices in a vector autoregression framework.

In the empirical tests of market rationality, I use a long period of data covering 120 years. Given the long period of my sample, some of the unexplained variation in stock prices is likely due to regime changes. For example, Fama and French (1988a) and Kim, Nelson, and Startz (1991) document that the mean reversion of stock prices is mostly due to the pre-World War II period. By contrast, Campbell (1991) argues that stock market predictability becomes clear only in the post-1950s. It is important to examine the possible regime shifts in the stock market rationality model with respect to World War II.

Section 2: Purpose of Study

The present study has five purposes: (1) to demonstrate that the fundamental component of stock prices is approximately a log-linear relationship among some relevant fundamental variables such as dividends, expected excess stock returns, expected inflation; (2) to estimate the fundamental component and the non-fundamental component using Gonzalo and Granger's (1995) methodology; (3) to investigate whether a time-varying inflation premium is one of the culprits of the rejection of the present-value relation; (4) to test whether the non-fundamental component of stock prices is significant after controlling for time-varying inflation premiums and risk premiums. (5) to examine the possible structural regime changes underlying the present-value relationship after World War II.

Section 3: Hypotheses and Propositions

Hypothesis 1: The fundamental component of stock prices is a random walk process, whereas the non-fundamental component of stock prices is a stationary temporary process.

Hypothesis 2: The fundamental and non-fundamental components of stock prices form the permanent and temporary decomposition in the Gonzalo and Granger (1995) sense.

Hypothesis 3: The non-fundamental component of stock prices is not significant. Thus, the present-value model still holds.

Hypothesis 4: There are no structural regime changes in the present-value relationship after World War II.

Proposition 1: The fundamental component of stock prices is approximately a log-linear relationship between the prices and fundamental variables such as dividends, expected inflation, and excess stock returns.

Proposition 2: Testing the significance of the non-fundamental component of stock prices can be conducted in the Gonzalo and Granger (1995) framework using Johansen and Juselius's (1990) weak exogeneity test of stock prices with respect to the cointegrated system.

Section 4: Limitations of the Study

Theoretically speaking, none of asset pricing models is testable. The reason is that financial assets are primarily priced according to investors' expectations rather than past or even current information, *per se*. Any test of market efficiency is a joint test of the efficient market hypothesis and the hypothesis that the asset pricing model used in the test captures all rational variations in the asset prices and/or returns. The inferences from the acceptance or rejection of the present-value model, run into this joint hypothesis problem.

My testing procedure uses *ex post* data to test an *ex ante* economic model and is subject to the limitation stressed by Roll (1977). In order to make an *ex ante* model testable, I have to impose some assumptions on stock price behavior (see Chapter 3 for elaboration). Although these assumptions survive the empirical tests, there is still a possibility that these assumptions may not represent the "true" behavior of stock prices.

Section 5: Organization Plan

Chapter 2 contains a thorough literature review for the following three related areas:

(1) Volatility tests and their implications

(2) Predictability tests of stock prices and their link with volatility tests

(3) An overview of the methodologies used by this line of research, including cointegration, vector-error-correction modeling, and permanent and temporary decomposition.

Chapter 3 derives the fundamental and non-fundamental components of stock prices.

Chapter 4 proposes an econometric model to decompose stock prices into fundamental and non-fundamental components and introduces the estimation and testing procedures.

Chapter 5 reports the empirical results and analyzes their implications.

Chapter 6 concludes the dissertation and highlights future areas of research.

CHAPTER 2

LITERATURE REVIEW

This chapter reviews the existing body of literature related to the present-value relationship, predictability of stock prices, and methodologies employed in these lines of research. In addition, I provide my own discussion of these studies. This chapter includes only the papers that I consider major contributions to the literature. Many other papers are also highlighted in the rest of this study.

A topic directly related to the present-value relationship is the variance-bounds debate or sometime called volatility tests, which I present in Section 1. It may be helpful, by way of motivation, to give at the outset a simple explanation indicating why excess volatility is fundamentally related to the predictability of multiperiod returns. Thus, I also review the literature of stock price predictability and compare it to the variance-bounds debate in Section 2. The variance-bounds literature is quite econometric oriented. It is necessary to outline the types of methodologies have been used in this area. In Section 3, I discuss the methodologies in the current literature in three categories: misspecification tests, cointegration and VAR approach, and the permanent/temporary decomposition. Finally, I also highlight the literature pertaining to the econometric methodologies that I use in my tests of the present-value model.

Appendix I provides a brief overview of the references cited in the present study.

Section 1: Does the Present-Value Relationship Hold? -- The Volatility Test Debate

Shiller (1981), along with LeRoy and Porter (1981), launched the variance-bounds literature trend. Shiller (1981) took the rational valuation formula as the basis for determining stock prices. Hence, stock prices are determined by economic fundamentals. When the variance of actual stock prices exceeds the maximum rational variation (the variance bound) in perfect foresight prices, stock prices are perceived to be too volatile to have been produced by rational investors. Volatility tests are joint tests of informational efficiency and that price equals the fundamental value that is represented by the present value model. Shiller (1981) finds that stock prices are too volatile to be justified by the fundamental values.

The simplest argument for excess volatility is given in the original LeRoy and Porter (1981) and Shiller (1981) papers. They argue that if, as the present-value model asserts, the actual price P_t should be the best expectation of *ex post* rational price P_t^*, the present value of actual future dividends, then the data must satisfy the variance inequality: $var(P_t^*) \geq var(P_t)$. The proof that the model implies this variance inequality is as follows. Since P_t is known at time t, I may write $P_t^* = P_t + u_t$, where u_t is a forecast error. If P_t is a sufficient statistic to forecast P^*, no information other than P_t can improve ones' forecast of P_t^*. This implies that the forecast error is a pure random error and independent of all information available at time t or earlier including P_t. That is, u_t must be uncorrelated with P_t. Therefore $var(P_t^*) = var(P_t) + var(u_t)$. Since variances cannot be negative, the variance inequality follows. This argument can be reversed to show that if the variance inequality is violated in U.S. data, then it must be that the forecast error $(P_t^* - P_t)$ is forecastable.

Shiller's (1981) implementation of an operational test of the inequality $var(P_t) \leq$

var(P_t^*) is simple and direct, more so than that of LeRoy and Porter. To correct for trend, Shiller divided through by constant growth rate trends. To understand Shiller's resolution of the problem that P_t^* is not observable, notice that the *ex post* rational price P_t^* is:

$$P_t^* = \beta(P_{t+1}^* + D_{t+1}) \tag{2.1}$$

where D_{t+1} is the one-period future cash flow over period t+1, and β is the discount factor. That is, β = 1/(1+K), where K is the discount rate.

Through recursion, the *ex post* rational price P_t^* can be solved as:

$$P_t^* = \left[\sum_{i=1}^{\infty} \beta^i D_{t+i}\right] + \left[\lim_{i\to\infty} \beta^i P_{t+i}^*\right] \tag{2.2}$$

Applying the terminal condition $\lim_{i\to\infty} \beta^i P_{t+i}^* = 0$ (2.3)

results in the formula for observable rational price P_t^* :

$$P_t^* = \left[\sum_{i=1}^{\infty} \beta^i D_{t+i}\right] \tag{2.4}$$

The terminal condition (2.3) says that the discounted value of the rational price P_t^* shrinks to zero as the horizon i increases. This condition will be valid unless the stock prices are expected to grow forever at the discount rate K or faster. In the "bubble" case, however, this assumption is relaxed, resulting in an infinite number of solutions to equation (2.2). Thus, any solution can be written in the form

$$\begin{aligned} P_t^* &= P_t^f + B_t \\ &= \left[\sum_{i=1}^{\infty} \beta^i D_{t+i}\right] + \left[\lim_{i\to\infty} \beta^i P_{t+i}^*\right] \end{aligned} \tag{2.5}$$

The first term P_t^f in equation (2.5) is the fundamental value, and the term B_t is often called a rational bubble. The word "bubble" recalls some of the famous episodes in financial history in which asset prices rose far higher than could easily be explained by fundamentals, and in which investors appeared to be betting that other investors would drive prices even higher in the future [For example, Mackay (1852) is a classic reference on early episodes such as the Dutch tulipmania in the 17th Century and the London South Sea Bubble and Paris Mississippi Bubble in the 18th Century.] However, many studies in the variance-bounds literature report consistent evidence against bubbles. See for example, West (1988b), Donaldson and Kamstra (1996), and Lee (1998).

Equation (2.4) is the solution to equation (2.1) that makes the rational price P_t^* observable. Using a century-long data set, Shiller (1981) finds that the standard deviation of actual stock prices exceeded that of the *ex post* rational stock prices by a factor of 5.59. Although no significance tests are reported, he interprets this result as constituting rejection of the variance-bounds inequalities.

LeRoy and Porter assume that dividends and stock prices, adjusted for trend as described below, are generated by a covariance-stationary bivariate linear process, with parameters restricted by:

$$\mathrm{var}(P^*) = \mathrm{var}(P) + \frac{\beta^2 \,\mathrm{var}(r)}{1-\beta^2} \tag{2.6}$$

where r represents excess returns [i.e., the difference between actual return and return based on last period's information set I_t: $r_t = d_t + p_t - E(d_t + p_t | I_{t-1})$].

This equation says that the variance of the *ex post* rational price equals the sum of the variance of actual prices and that of returns (where the latter is multiplied by a constant that depends on the discount rate). The variance of actual prices P_t are bounded by the variance of rational prices P_t^*. Hence, equation (2.6) underlies the logics of almost all the variance-bounds tests.

A simplified and intuitive version of the LeRoy-Porter implementation is as follows: (a) estimate a linear autoregressive model for dividends and estimate β as the reciprocal of 1 plus the average rate of return on a stock; (b) estimate var(P^*) by applying the present-value relation (2.1) directly to the model for dividends (thus avoiding the problem that P_t^* is unobservable); (c) estimate var(r_t)from the observable series of one-period returns; and (d) estimate var(P_t) from a linear model for P_t. LeRoy and Porter find that the point estimate of each of the two terms on the right-hand side of equation (2.6) by itself exceeds the term on the left-hand side, indicating rejection of the variance bounds.

Flavin (1983) criticizes Shiller's econometric tests from two aspects. The first is that both the variance of P_t and that of P_t^* are estimated with downward bias in small samples. Further, the effect is more severe for P_t^* than for P_t, implying a possible reversal of the empirical counterpart of the variance-bounds inequalities even if the present-value relation is true. The second is that Shiller's procedure for calculating an observable version of P_t^* also induces bias toward rejection.

Kleidon (1986) criticizes Shiller's (1981) contention that the smoothness of a time-series plot of P_t^* relative to P_t contradicts the variance-bounds theorems. To Kleidon, such a conclusion is completely unwarranted. While it is true that the variance-bounds inequality itself is model free, the properties of any econometric test of that inequality can only be investigated conditional on a particular dividends model. He argues that under reasonable specifications of the dividends model, variance-bounds tests will reject with high probability even if the present-value model is true. As Gilles and LeRoy (1991) comment, there can be no doubt that, at a minimum, the critics established that econometric problems with variance-bounds tests are potentially severe. Whether these problems are severe enough to account for the extent of the apparent excess volatility, however, remains controversial. In addition, Kleidon (1986) also criticizes Shiller's variance-bound test on the stationarity ground that the dividends may not follow a trend-stationary process as Shiller claims.

West (1988a) derives a variance-bounds test that (a) is valid even if dividends are nonstationary, and (b) does not require a proxy for the rational prices P_t^*. Even if dividends are generated by a linear process with a unit root (so that dividends and prices are cointegrated rather than stationary), the population return variances will be constant. Therefore their sample counterparts provide consistent estimates of population values. However, Gilles and LeRoy (1991) argue that a more natural treatment of trend is to specify a log-linear, rather than linear, dividend process, so that dividend growth rates are stationary. LeRoy and Parke (1992) adapt West's (1988a) variance-bounds test to the log-linear case. Campbell and Shiller's series of papers also incorporate the log-linear forms for the present value model.

Mankiw, Romer and Shapiro (1985, 1991) (hereafter MRS) claim to have provided an unbiased volatility test. After deriving this test, they provide no evidence of excess volatility.

Shea (1989) points out two major problems with MRS's tests. First, the outcome of the tests is very sensitive to the choice of terminal date. Second, because of the nonstationarity induced by the dependence of both population parameters and statistics, there is no prospect of using asymptotic theory to derive confidence intervals for the tests. Correcting for MRS's two problems, Shea provides results that are much more favorable to the variance-bounds theorems than MRS's.

Marsh and Merton (1986) and Merton (1987) observe that dividend smoothing by management could bias variance-bounds tests in general, and MRS's test in particular, toward rejection. If dividends are slow to reflect changes in underlying profitability, measured dividend volatility could give the impression that fundamentals had remained stable even when the opposite is the case. MRS (1991) respond that, empirically, Merton's criticism is of little practical importance.

Campbell and Shiller (1987) note that if the present value model is true, then (a) an optimal prediction of the present value of future expected dividends can be formed using current price alone; and (b) this optimal prediction coincides with current price. It follows that the present value model implies testable restrictions of the coefficients of a bivariate vector autoregression of stock prices and dividends.

Campbell and Shiller (1988a) introduce a method for trend correction that, although not perfect, is superior to anything that went before. They assume that dividends and whatever other variables predict dividends (e.g. time-varying real interest rates, time-varying excess stock returns) form a multivariate log-linear present value model. To reconcile the log-linearity of the dividend model with the linearity of the present-value relation, they log-linearized the expression defining the rate of return. After applying Taylor's approximation, they derive a log-linear present value model (sometime called the "dynamic Gordon model).[1]

[1] Similar derivation of the log-linear present value model can be found in Chapter 4 of this thesis.

Comparison of the actual rate of return with its log-linearized counterpart–the two are correlated almost perfectly–allows Campbell and Shiller to argue that the error introduced by the log-linearization is negligible. The present-value model that results from iterating the log-linearized version of the definition of the rate of return (r_t) expresses the log price–dividend ratio as the present value of the discounted expected dividend growth rates.

$$p_t - d_t = E_t \sum_{j=0}^{\infty} \rho^j [\ (1-\rho)\Delta d_{t+1+j} - r_{t+1+j}\ ] + \frac{k}{1-\rho} \tag{2.7}$$

where, p_t = log real stock prices at the end of period t,

d_t = log real dividends during time period t,

r_t = log real rate of return during period t, and

ρ and k are parameters of linearization.

All these variables are essentially free of trend, avoiding the econometric problems attending the earlier volatility tests by reason of the non-stationarity of the underlying series. Accounting for time-varying real interest rates and risk premiums (proxied by excess stock returns), Campbell and Shiller (1988a) report the results of a variety of tests of the equality of the log price–dividend ratio and the present value of future dividend growth rates, and find robust evidence of significant violation

Campbell and Shiller (1988b) add corporate earnings to the price-dividend vector autoregression. They find that earnings are a strong predictor of dividend growth (return on stock) even conditional on the current log price-dividend ratio. Their findings contradict the simple present-value model, which says that current price is a sufficient statistic for future dividend growth. They argue that a long moving average of earnings is a very natural proxy to represent fundamental value, and that there are not many competitors for this role. However, I argue that accounting earnings are not direct cash flows to the stocks, and only dividends are distributed cash flows that have direct effects on stock prices. Changes in earnings may influence the expectation of future cash flows, but never enter the present value as the numerator. Therefore, I do not incorporate earnings in the present value model, nor should it be included according to theory.

Gilles and LeRoy (1991) derive a variance bound test that is valid if dividends follow a geometric random walk and stock prices are non-stationary (but cointegrated). They therefore assume the dividend-price ratio is stationary and their variance inequality is $var(P_t|D_t) \leq var(P^*_t|D_t)$. The sample estimates of the variances [1871-1988 US aggregate index as used in Shiller (1981)] indicate excess volatility. However, they note that the sample variance of $var(P^*_t|D_t)$ is biased downward for two reasons. First because $(P^*_t|D_t)$ is positively serially correlated (Flavin, 1983) and secondly because at the terminal date the unobservable $E_tP^*_{t+n}$ is assumed to equal the actual (terminal) price P^*_{t+n}. (Hence, dividend innovations after the end of the sample are assumed to be zero, Merton (1987).) Using Monte Carlo experiments, they find that the first source of bias is most important and is very severe. Thus, Gilles and LeRoy conclude that the Shiller-type variance bounds test is "indecisive" (1991, p.986). They also develop a test based on the orthogonality of P_t and P^*_t (similar to West 1988) which is more robust. This "orthogonality test" uses the geometric random walk assumption for dividends [$\ln D_{t+1} = \ln D_t + \varepsilon_{t+1}$, with $E\varepsilon_{t+1} = \mu$, $var(\varepsilon_{t+1}) = \sigma^2$] and involves a test statistic with much less bias and less sample variability than the Shiller-type test. The orthogonality test rejects the present value model quite decisively.

Shiller and Beltratti (1992) analyze the relation between real stock prices and long-term interest rates within the dynamic Gordon model (this model is also called the rational expectations present value model) derived in Campbell and Shiller (1988a). They find that real stock prices fall when long-term interest rates rise (and rise when they fall) more than would be implied by the simple present value model. In view of the nature of the variability of discount rates and dividends in relation to information available in advance of this variability, there should indeed be generally a slight negative correlation between changes in real stock prices and changes in long-term interest rates, but the actual observed correlation is more negative in U.S. and U.K. data than it should be. This implies that stock prices "overreact" to bond yields–a similar conclusion of previous variance bounds literature. In this thesis, Shiller and Beltratti raise an important notion (in their footnote 3) that the nominal and the real interpretation of the present value model can be considered different ways of making the model suitable to empirical testing by turning nonstationary variables into stationary variables. For example, they use a real version of the model for stock prices but use the nominal version for bond yields. This opens the opportunity of using the nominal version of the present value model for stock prices (as used in this dissertation so that I can account for the effects of expected inflation).

In sum, the variance-bounds debate generally favors the finding of significant excess volatility, with the exception of MRS (1991). When the variance bounds tests are rejected, some studies provide a "fad" interpretation while others provide a "bubble" interpretation. The present value relation is derived based on an Euler equation combined with a transversality condition. When prices do not satisfy the transversality condition, they are thought to contain bubbles. Most studies, however, do not interpret the variance bounds rejection as evidence for bubbles partly because the tests are based on finite samples [see Shiller (1984) and West (1988b)]. Both in fads and bubbles, the stock price deviates from the present value of expected future dividends or fundamental values due to either noise trading, or feedback trading (trade based on past price changes), or irrational expectations (irrational waves of optimism and pessimism), or some other inefficiency. However, fad price deviations are expected to slowly decay to zero, whereas bubble price deviations are expected to last forever [see Cochrane (1991, p.471)]. Therefore, time-series behavior of price deviations is bound to shed some light on this debate.

Section 2: Volatility and Predictability: A Comparison

Volatility tests are a joint test of informational efficiency and that price equals fundamental value. Predictability tests include autocorrelation tests such as Fama and French (1988a) and Porterba and Summers (1988), and regression tests such as the Fama and French (1988b) test the relationship between actual price P_t and the perfect foresight price P_t^*.

Fama and French (1988a) estimate an autoregression model where the return over the interval t-N to t, called $R_{t-N,t}$, is correlated with $R_{t,t+N}$.

$$R_{t,t+N} = \alpha + \beta R_{t-N,t} + \varepsilon_t \tag{2.8}$$

They consider return horizons N from one to ten years. They find little or no autocorrelation, except for holding periods of between N=2 and N=7 years for which β is less than zero. That is, they find that the returns have negative autocorrelation in the long time horizon. There is a peak at N=5 years when β=-0.5, suggesting that a 10 percent negative return over five years is, on average, followed by a 5 percent positive return over the next five years. The value of R-squared in the regressions for the three to five-year horizons is about 0.35. Such a mean reversion ($\beta<0$) is consistent with that from the so-called "anomalies literature" where a "buy low, sell high" trading rule earns persistent positive profits. Fama and French (1988a) interpret this negative autocorrelation that stock prices are the sum of a random walk component and a stationary temporary component. However, Fama and French's findings of the temporary component of stock prices appear to be mainly due to the inclusion of a 1930s sample period (Fama and French 1988a). Autocorrelations for periods after 1940 are closer to 0.0, and they do not show the U-shaped pattern of the overall period of 1926-85. Due to the small sample size, the Fama-French regression testing approach may have little power.

In contrast to Fama and French (1988a) who find negative return autocorrelation, Lo and MacKinlay (1988) document a positive autocorrelation for weekly returns using their variance ratio test. This test capitalizes on the fact that the variance of the increments in a random walk is linear in the sampling interval. That is, if a series follows a random walk process, the variance of its q-differences would be q times the variance of its first difference. Hence, a variance ratio less than one should imply negative serial correlation, while a variance ratio greater than one implies positive serial correlation.[2] They find that the equally weighted NYSE index has as high as 30 percent autocorrelation! This finding casts doubt on the explanation that stock prices are the sum of a random walk component and a stationary mean-reverting component. If returns are in fact generated by such a process, then

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[2] Some researchers use the unit root tests to test for predictability of stock prices. The unit root tests only examine the permanent/transitory nature of shocks to a series. Even under the null hypothesis of unit root, the increments of the price series may be predictable. Indeed, there are also nonrandom walk alternatives in the unit root null hypothesis. Therefore, the unit root tests are clearly not designed to detect predictability.

their variance ratios should be less than unity when the interval q =2 (since negative correlation is implied by this process).

Porterba and Summers (1988) also investigate mean reversion by analyzing variances of holding period returns over different horizons. Their results suggest that stock returns show positive serial correlation over short periods and negative correlation over longer intervals. If stock returns are random, then variances of holding period returns should increase in proportion to the length of the holding period. They find that the variance of returns increases at a rate which is less than proportional to N, implying that returns are mean reverting (for 8>N>3 years). This conclusion is generally upheld when using a number of alternative stock price indexes, although the power of the tests is low when detecting persistent yet transitory returns. Using data on equally weighted and value-weighted NYSE returns over the 1926-1985 period and data from other nations and time periods, Poterba and Summers (1988) test the significance of a transitory price component using their point estimates. Results suggest that although individual data sets do not *consistently* permit rejection of the random-walk hypothesis at high significance levels, the transitory price component generally accounts for a substantial part of the variance in returns. Poterba and Summers (1988) also discuss the potentially important implications for financial practice. If stock price movements contain large transitory components, then for long-horizon investors the stock market may be less risky than it appears to be when the variance of single-period returns is extrapolated using the random-walk model. The presence of transitory price components also suggests the desirability of investment strategies, such as those considered by DeBondt and Thaler (1985), involving the purchase of securities that have recently declined in value ["buy loser, sell winner"].

Fama and French (1988b) extend their earlier univariate study on the predictability of expected returns over different horizons and examine the relationship between (nominal and real) returns and the dividend yield D/P.

$$R^{N}_{t+N} = \alpha + \beta(D/P)_t + \varepsilon_t \tag{2.9}$$

The equation is estimated for monthly and quarterly returns, and for annual returns of one to four years on the NYSE index. They also test the robustness of the model by estimating it over various sub-periods. For monthly and quarterly data, the dividend yield is often statistically significant (and $\beta<0$), but explains only about 5 percent of the variability in monthly and quarterly actual returns. For longer horizons, the explanatory power increased. The longer return horizon regressions also prove useful in forecasting "out-of-sample".

Although variance bound test literature and the predictability literature have different starting approaches, these two lines of research are closely related. Campbell and Shiller (1988b) provide a valuable discussion of the relation between the volatility tests and the return autocorrelation tests conducted by Fama and French (1988a), Poterba and Summers (1988) and others. They said "excess volatility and predictability of multiperiod return are not two phenomena, but one" (p.663).

The easiest way of seeing the relationship between the variance-bound tests and the predictability tests of Fama and French (1988a) is to note that in the regression of the rational price P_t^* [see equation (2.1)] and the actual price P_t.

$$P_t^* = a + bP_t + \varepsilon_t \tag{2.10}$$

The coefficient b is given by

$$b = \frac{\text{cov}(P_t, P_t^*)}{\text{var}(P_t)} \tag{2.11}$$

Since in the variance-bound tests I have $P_t^* = P_t + u_t$, cov(Pt, P_t^*) = cov(P_t, P_t+u_t) = cov(P_t, P_t) + cov(P_t, u_t) = cov(P_t, P_t) = var(P_t). Therefore, I obtain b=1 in equation (2.11). Hence, if the variance equality holds then I expect b=1 in the regression (2.10). Next, consider the long-horizon regression of Fama and French (1988a):

$$R_t^N = a + R_{t-N}^N + \eta_t \tag{2.12}$$

$$\ln P_{t+N} = a + (b+1)\ln P_t - b\ln P_t + \eta_t \tag{2.13}$$

where I have used $R_t^N = \ln P_{t+N} - \ln P_t$. Under the null hypothesis that expected returns are constant (a≠0) and independent of information at time t or earlier, the regression coefficient b is expected to be 0. If this is true, then from (2.13)

$$\ln P_{t+N} = a + \ln P_t + \eta_t \tag{2.14}$$

Hence under the null, H_0: b = 0, the Fama-French regressions are broadly consistent with the random walk model of stock prices.

Campbell (1991) argues that expected stock returns change through time in a fairly persistent fashion. The variability and persistence of expected stock returns account for a considerable degree of volatility in unexpected returns. The variance of news about future cash flows (dividends) accounts for only a third to a half of the variance of unexpected stock returns. The remainder of the stock return variance is due to news about future expected returns. Further, news about future returns is not independent of news about cash flows.

Increases in future expected cash flows tend to be associated with decreases in future expected returns. In addition, the variability of news about future excess stock returns is much greater than the variability of news about future real interest rates, and the latter has only a relatively small impact on stock returns. This casts doubt on explanations of variation in expected real stock returns which rely primarily on movements in real interest rates [e.g. Cecchetti *et al* (1990)]. A caveat is also worth mentioning in Campbell's (1991) study. Both asymptotic standard errors and the results of a small Monte Carlo experiment show that there is only weak evidence for stock return predictability in the prewar period. The evidence that returns are predictable is overwhelming only in the period after 1952.

Based on the empirical findings of stock return predictability, Reichenstein and Rich (1994) go further and discuss how investors would exploit this predictability in their own investment portfolios.

A similar question is addressed by Lander, Orphanides, and Douvogiannis (1997), and they find that converting a simple mean-reverting theory into a trading rule can yield significantly higher returns (in a statistical sense) than would be expected by pure chance alone.

In sum, these recent works generally suggest that financial asset returns are predictable to some degree. Thirty years ago this would be tantamount to an outright rejection of market efficiency. However, modern financial economics tells us that other perfectly rational factors may account for such predictability. The imperfect structure of securities markets and frictions in the trading process can generate predictability. Time-varying expected returns due to changing business conditions and risk factors can generate predictability. A certain degree of predictability may be necessary to reward investors for bearing certain dynamic risks. Nevertheless, the predictability, or the presence of a significant predictable component, of asset returns (or prices) may still reflect market inefficiency. It is necessary to examine the presence of the predictable (also called "temporary") component of asset prices using more powerful statistical techniques (Fama and French, 1988a).

Section 3: Methodological Advances

Econometric issues have dominated the variance-bounds debate. Testing volatility and predictability of stock returns is characterized by adopting newly developed econometric techniques. In addition to the literature that I have outlined in previous section, I present some other related studies which stress methodological issues.

Subsection 3.1: Misspecification Tests

Looking at regression equations that attempt to explain returns, an econometrician is typically interested in general diagnostic tests (e.g. are the residuals normal, serially uncorrelated, homoscedastic, explanatory variables weakly exogenous, etc), as well as in checking the outside sample forecasting performance of the equations and the temporal stability of the parameters. In many of the previous studies, this useful statistical information is not always fully presented, so it becomes difficult to ascertain whether the results are as "robust" as they are claimed to be.

Pesaran and Timmermann (1994) provide a study of stock returns that attempts to address the above criticisms of earlier work. They look at excess returns on the S&P 500 index and the Dow Jones index measured over one year, one quarter and one month for the period 1954-1971 and include the dividend yield, annual inflation, the change in the three-month interest rate, and the term premium. Their findings reinforce the earlier results that excess returns are predictable and can be explained quite well by a relatively small number of explanatory variables.

On the other hand, McQueen (1992) casts some doubt on the significance, even the existence, of long-horizon predictability. Using a generalized least squares (GLS) test for 1926-1987 period, McQueen fails to reject the hypothesis that monthly stock returns follow a random walk.

Subsection 3.2: Cointegration and VAR Approach

Campbell and Shiller (1987) incorporate the newly-developed cointegration technique into testing the long-run relationship between stock prices and dividends, finding a significant cointegrating vector during the years 1871-1986. They find that the spread between stock prices and dividends moves too much and that deviations from the present value model are quite persistent, although the strength of the evidence for this is sensitive to the discount rate assumed in the test.

Using stochastic simulation, Campbell and Shiller (1989) derive the small sample properties of parameter estimates and test statistics in the vector autoregressive dividend ratio model of Campbell and Shiller (1988a). They find that although there is some indication of small sample bias, the extent of the bias is not enough to reconcile the difference between the actual dividend ratio and the present value of future dividend growth. This suggests that the rejection of the present-value model cannot be justified by the small sample bias as raised by Flavin (1983), Kleidon (1986), Marsh and Merton (1986) and others.

Lee (1995, 1996b, 1998) also documents the presence of cointegration relationship among stock prices, dividends, and earnings. Particularly, Lee(1996b) investigates the comovements of earnings, dividends, and stock prices in a three-variable cointegrating system. He finds that the three series are cointegrated with a single cointegrating vector, suggesting that there is an equilibrium force that tends to keep these series together over time. Lee (1996b) also finds that a substantial fraction of stock price movement is driven by neither earnings changes nor dividend changes. Since dividends and earnings are thought to be fundamental variables of stock prices. Such a finding implies that stock prices deviate from the fundamental value [captured by the present value model].

Subsection 3.3: Permanent-
Temporary Decomposition

One motivation for using long-horizon returns is the permanent/transitory components alternative methodology, pioneered by Muth (1960) in a macroeconomic context. In this model, log prices are composed of two components: a random walk and a stationary process,

$$p_t = w_t + y_t \tag{2.15}$$

where, $w_t = \mu + w_t + \varepsilon_t$ $\varepsilon_t \sim \text{IID}(0, \sigma^2)$

y_t = any zero-mean stationary process,

and $\{w_t\}$ and $\{y_t\}$ are mutually independent. The common interpretation of the above equations modeling stock prices is that w_t is the "fundamental" component that reflects the efficient markets price, and y_t is a zero-mean stationary component that reflects a short-term or "transitory" deviation from the efficient-markets price w_t, implying the presence of "fads"[3] or other market inefficiencies. Since y_t is stationary, it is mean-reverting by definition and reverts to its mean of zero in the long run.

In a series of papers, Lee (1995, 1996a, 1998) elaborated on the permanent component and temporary decomposition of stock prices. Lee (1995) investigates the response of stock prices to permanent and temporary shocks to dividends. He relates the permanent and temporary components of dividends to stock prices. Having identified the permanent and temporary shocks to dividends by imposing the identifying restriction, Lee examines the relationship between prices and these two types of shocks. He finds that stock prices respond significantly to both the permanent and the temporary shocks to dividends. Furthermore, the initial response of stock prices to the temporary shocks is as strong as the initial response to the permanent shocks. Lee's findings add evidence to the mounting literature in support of the observed mean-reverting behavior of stock returns by incorporating a temporary component into stock prices.

After decomposing the dividends and earnings into two respective P-T components, Lee (1996a) documents that dividends respond strongly to permanent changes in earnings without any significant overreaction, whereas dividends respond little to transitory changes in earnings. His findings support the hypothesis that dividend changes are determined by changes in some measures of permanent earnings. Thus, the implication is that managers will perform better when the target dividend level is proportional to permanent earnings rather than to current earnings. This evidence seems to support the Marsh and Merton (1987)

[3] Fad implies the presence of a zero-mean stationary component that reflects a short-term or transitory deviation from the efficient market prices.

dividend smoothing hypothesis which is introduced to explain the seemingly excess volatility of stock prices that Shiller (1981) finds.

Lee (1998) attempts to answer the question whether the rejection of the simple present value relation and the mean reversion in stock returns can be explained by either time-varying discount rates or non-fundamental factors ("fads" or "bubbles"[4]) By allowing for time-varying discount factors in the model, Lee (1998) identifies various components of stock prices and examines the response of stock prices to different types of shocks: permanent and temporary changes in earnings and dividends, changes in discount factors, and non-fundamental factors. The identification of these innovations is achieved by imposing restrictions on the models of earnings, dividends, discount factors, and stock prices that take into account cointegrations among these variables. Lee finds that, although the long-term trend in stock prices is due to permanent changes in fundamentals, the short-term volatility is largely due to the discount factor changes reflected in excess stock return changes, but also partly due to non-fundamental factors. This suggests that the over-reaction of the stock market and the mean reversion in stock returns are primarily in response to excess return changes, and partly in response to non-fundamental factors (which is consistent with DeBondt and Thaler (1985) and Summers (1986)).

Lamoureux and Zhou (1996) argue that whether returns consist of a material stationary (temporary) component is questionable and perhaps due to inadequate data. Adopting a subjectivist analysis (treating the data as fixed), they employ a Bayesian

[4] Theoretically, stock prices equals a fundamental value plus a "fad" term or a "bubble", where the "fad" is a transitory component whereas the "bubble" term follows a persistent martingale process.

approach and let the data determine the permanent/temporary decomposition. Their findings suggest that stock prices follow a random walk.

According to the above major research literature, whether stock prices contain a significant predictable component is an enduring question. Earlier studies obviously suffer from low-powered tests [Porterba and Summers (1988)], small sample bias [Cecchetti et al (1990)] and misspecification problems [Pesaran and Timmermann (1994)]. More recent research [Lee (1995, 1996a, 1996b, 1998) and Lamoureux and Zhou (1996)] adopt the more sophisticated P-T decomposition technique of Blanchard and Quah (1989) and Quah (1992). Yet, this decomposition method is designed for obtaining orthogonal P-T decomposition in a *univariate* setting.

As noted by Gonzalo and Granger (GG, 1995), the GG decomposition method is the only multivariate permanent-temporary (P-T) decomposition method available thus far. This GG P-T decomposition method is conducted in the context of a cointegrated system using the vector error correction model (VECM) of Johansen (1988, 1991) and Johansen and Juselius (1992). Interestingly, Campbell, Lo and MacKinlay (1997) conjecture that there is only weak evidence for predictability of long-horizon stock returns in a univariate setting, but there could be stronger evidence for predictability of long-horizon returns once the model is enlarged to incorporate other relevant information. It is logical to speculate that the P-T decomposition ought to be done in a multivariate setting to avoid possible omission-of-variables bias. The information set should include relevant variables in the present-value model, i.e, dividends, time-varying discount rates, and inflation premiums.

Gonzalo and Granger (1995) propose a common-long-memory estimation technique to identify permanent and temporary components (P-T) in a multivariate context. Earlier P-T decompositions have been designed and used in a univariate framework. Stock and Watson (1988) propose a common-trend decomposition that basically extends the univariate decomposition proposed by Beveridge and Nelson (1981) to cointegrated systems. Nevertheless, the advantage of the Gonzalo and Granger (GG) decomposition with respect to the common-trends model of Stock and Watson is that with the GG method, it is easier to estimate the common long-memory components and to test hypotheses on these common long-memory components. In this thesis, I utilize the GG method to investigate whether stock prices have a significant non-fundamental component that is predictable. In contrast to the univariate P-T decomposition of Quah (1992), I employ the GG multivariate decomposition method to decompose stock prices into permanent and temporary components in the cointegrated system. I then construct a formal test in light of GG's methodology to test whether stock prices contain a significant temporary (non-fundamental) component. Finding a significant non-fundamental component indicates that stock prices deviate from their fundamental value and hence possess excess volatility.

CHAPTER 3

DERIVING FUNDAMENTAL AND NON-FUNDAMENTAL COMPONENTS OF STOCK PRICES

My analysis starts with the simple present value relation

$$P_t = E_t\left[\sum_{j=0}^{\infty}\frac{D_{t+1+j}}{(1+r)^{1+j}}\right] \tag{3.1}$$

where P_t is the nominal stock price at the end of period t; D_{t+1+j} is the nominal net cash flows in period t+1+j; r is the nominal required rate of return; E_t is the expectation operator. This present-value relation indicates that the current price is the discounted value of all future expected net cash flows discounted at a constant required rate of return. Below I generalize this simple present-value equation to allow for time-varying required rate of return in the denominator and accounting earnings in the numerator.

In this Chapter, I derive the fundamental and the non-fundamental components using Campbell and Shiller's (1988a) dividend-price ratio model. In my model, the fundamental component of stock prices allows for a role of time-varying inflation premiums. My model differs from Campbell and Shiller's (1988a) and Lee's(1998) in that I focus on the nominal, rather than real, interpretation of the model. Campbell and Amber (1993) and Lee (1998) report results denying any effect for real interest rates in explaining variations of stock prices.

However, unlike them, I postulate that using the nominal rate is more appropriate to investigate the validity of the present-value model for at least five reasons.

First, in practice, the standard PVM is commonly expressed in nominal, not real, terms. Secondly, the use of real interest rates in the PVM necessitates the use of unobservable data that could introduce possible biases due to the use of "extracted" data. Indeed, the real version (but not the nominal version) of PVM is susceptible to possible "price deflator bias." Thirdly, nominal interest rates largely reflect changes in expected inflation, particularly in annual (long-run) data [Sargent (1972)]. Thus, and following Fama and Schwert (1977), Geske and Roll (1983), and James *et al.* (1985), nominal interest rates can be provisionally taken to represent expected inflation. Fourthly, previous studies report that real interest rates explain little variation in stock prices [Campbell and Shiller (1988a), Campbell and Amber (1993), and Lee (1998)]. Therefore, if nominal interest rates are found to possess any explanatory power for stock prices, such effects most likely come from expected inflation. Finally, I attempt to uncover other explanatory variables for determining stock prices that have been ignored in previous studies. It is conceivable that expected inflation is one of such candidates [Fama (1991)]. All previous research thusfar employs the real interpretation of the PVM and hence ignores the role of inflationary expectations, although expected inflation is an important determinant of the required rate of return and thus of the present-value of future dividends.[5]

In addition to accounting for time-varying inflationary expectations, I directly estimate the functional form of the fundamental component of stock prices (i.e., the linear

[5]I should also note, following Shiller and Beltratti (1992), that both the nominal and real interpretations of the PVM are equally plausible, and either interpretation can be adopted depending on empirical convenience (most importantly, satisfying stationarity requirements).

present-value relationship)[6]. Specifically, following Fama and French (1988a)--but unlike Lee (1998)--I treat *nominal* stock prices as the sum of only two components, fundamental and non-fundamental elements.

Similar to Campbell and Shiller (1988a) and Campbell and Amber (1993), I derive a general log-linear model for *nominal* stock prices that allows for time-varying inflationary expectations (proxied by the nominal interest rates) and time-varying excess stock returns (risk premiums). The model can be written as:

$$p_t = E_t \sum_{j=0}^{\infty} \rho^j [\ (1-\rho)d_{t+1+j} - i_{t+1+j} - er_{t+1+j}] + \frac{k}{1-\rho} \tag{3.2}$$

where ρ and k are the parameters of linearization; E_t stands for the conditional expectations during period t, and

p_t = log nominal stock prices at the end of period t,

d_t = log nominal dividends during time period t,

i_t = log nominal interest rates during period t,

er_t = log excess nominal stock returns on a stock held during period t relative to the nominal return on short debt.

A log-linear framework has several advantages over the linear model: (1) Empirical literature indicates that stock prices and dividends are like many other macroeconomic time series in that they appear to grow exponentially over time rather than linearly [see Campbell, Lo and MacKinlay (1997)]; (2) A log-linear model is more convenient to use than a non-

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[6] Lee (1998) uses the permanent-temporary decomposition approach of Quah (1992). This approach, however, is incapable of estimating the functional forms of fundamental components. Further, in order to identify the structural VAR, arbitrary and perhaps inappropriate restrictions must be imposed [see Crowder and Wohar (1998)].

logged model because, with logged series, the first difference in dividend-adjusted stock prices will be the stock return and the spread will be the dividend yield [Lee (1995)]; (3) The estimated coefficients in a log-linear model represent elasticities and thus are easier to interpret.

To see how equation (3.2) is derived, consider the definition of the log or continuously compounded stock return r_{t+1} over period t to t+1. By convention, logs of variables are denoted by lowercase letters and non-logged variables denoted by capitals.

$$\begin{aligned} r_{t+1} &= \log(P_{t+1} + D_{t+1}) - \log(P_t) \\ &= \log(P_{t+1}(1 - D_{t+1} / P_{t+1}) - \log(P_t) \\ &= \log(P_{t+1}) - \log(P_t) + \log(1 - D_{t+1} / P_{t+1}) \\ &= \log(P_{t+1}) - \log(P_t) + \log\{1 - \exp[\log(D_{t+1}) - \log(P_{t+1})]\} \\ &= p_{t+1} - p_t + \log\{1 - \exp[d_{t+1} - p_{t+1}]\} \end{aligned} \tag{3.3}$$

The last term on the right-hand side of (3.3) is a non-linear function of the log dividend-price ratio, $f(d_{t+1} - p_{t+1})$. According to the first-order Taylor expansion theorem, any non-linear function $f(x_{t+1})$ can be approximated around the mean of x_{t+1}, $\bar{x}$.

$$f(x_{t+1}) \approx f(\bar{x}) + f'(\bar{x})(x_{t+1} - \bar{x}) \tag{3.4}$$

Let $x_{t+1} = d_{t+1} - p_{t+1}$. Substituting this approximation into (3.3), I obtain

$$\begin{aligned} f(x_{t+1}) &= \log[1 + \exp(x_{t+1})] \\ &\approx f(\bar{x}) + f'(\bar{x})(x_{t+1} - \bar{x}) \\ &= \log[1 + \exp(\bar{x})] + \frac{\exp(\bar{x})}{1 + \exp(\bar{x})}(x_{t+1} - \bar{x}) \end{aligned} \tag{3.5}$$

Thus,

$$
\begin{aligned}
r_{t+1} &\approx p_{t+1} - p_t + \log[1+\exp(\bar{x})] + \frac{\exp(\bar{x})}{1+\exp(\bar{x})}(x_{t+1} - \bar{x}) \\
&= p_{t+1} - p_t + \log[1+\exp(\bar{x})] + \frac{\exp(\bar{x})}{1+\exp(\bar{x})}(d_{t+1} - p_{t+1} - \bar{x}) \\
&= [1 - \frac{\exp(\bar{x})}{1+\exp(\bar{x})}]p_{t+1} - p_t + \frac{\exp(\bar{x})}{1+\exp(\bar{x})}d_{t+1} \\
&\quad + \log[1+\exp(\bar{x})] - \frac{\exp(\bar{x})}{1+\exp(\bar{x})}(\bar{x})
\end{aligned}
\tag{3.6}
$$

Let $\rho = \dfrac{1}{1+\exp(\bar{x})}$, $k = -\log(\rho) - (1-\rho)\log[(1/\rho) - 1]$. Then (3.6) becomes

$$r_{t+1} \approx \rho p_{t+1} - p_t + (1-\rho)d_{t+1} + k \tag{3.7}$$

Equation (3.7) is the log-linear definition of *ex post* stock returns. Rearrange (3.7) with the current price as the dependent variable.

$$p_t = k + \rho p_{t+1} + (1-\rho)d_{t+1} - r_{t+1} \tag{3.8}$$

Since p_t is non-stationary, the right-hand side of (3.8) is also non-stationary.

Equation (3.8) is measured *ex post*, and the next-period discount rate r_{t+1} is not observable in period t. To assign an economic meaning to expression (3.8), and following Campbell and Shiller (1988a), I impose some restrictions on the behavior of nominal discount rates. In particular,

$$E_t r_{t+1} = r_0 + E_t \pi^e_{t+1} + E_t er_{t+1} \tag{3.9}$$

where E_t denotes a rational expectation operator formed by using the information set I_t that is available to market participants at the end of period t, r_0 is a constant real riskless rate, π_{t+1}^e

is inflation expected at time t for time t+1, and er_{t+1} is the risk premium (excess stock returns) measured by the nominal gross return on a given stock during time t to t+1 relative to the nominal return on short debt i_{t+1}. Note that there is no inflationary premium component in excess stock returns, since the inflationary effect is canceled out after subtraction. Equation (3.9) implies that the *ex ante* return on stocks over the period t to t+1 equals a constant real riskless rate plus a inflation premium and a risk premium. I use a nominal interest rate on commercial paper i_{t+1} to approximate the variation in expected inflation: $\left(i_{t+1} = r_0 + \pi^e_{t+1}\right)$.

Hence

$$E_t r_{t+1} = E_t i_{t+1} + E_t er_{t+1} \tag{3.10}$$

Equations (3.8) and (3.10) give an economic model of the fundamental value of stock prices:

$$p_t^f = E_t\left[\rho p_{t+1} + (1-\rho)d_{t+1} - i_{t+1} - er_{t+1} + k\right] \tag{3.11}$$

Equation (3.11) states that the fundamental value of stock prices of this period is a log-linear combination of rational expectations of the next period's prices, dividends, nominal interest rates (a proxy for expected inflation), and a risk premium.

Note that expression (3.11) is equivalent to Campbell and Shiller's (1988a) dividend-price ratio model. Solving forward equation (3.11) and imposing a terminal condition that $\lim_{j\to\infty} \rho^j p_{t+j} = 0$ yields

$$p_t^f = \mathrm{E}_t\left[\sum_{j=0}^{\infty} \rho^j [\,(1-\rho)d_{t+1+j} - i_{t+1+j} - er_{t+1+j}]\right] + \frac{k}{1-\rho} \tag{3.12}$$

which is implied by Campbell and Shiller's (1988a) dividend-ratio model. However, this expression of the fundamental value eliminates the possibility for rational bubbles. Although I acknowledge that the rational bubbles hypothesis may not hold in practice [see, for example, West (1988b), Donaldson and Kamstra (1996)], I allow for the possibility of rational bubbles and use equation (3.11) to test the rational market hypothesis.

However, since the value at time t+1 is not yet known to investors at time t, equation (3.11) is not empirically testable. Thus, some restrictions must be placed on the behavior of stock prices to derive a testable economic model. To simplify the case, I only impose two common assumptions:

Assumption 1: *The fundamental component of stock prices follows a random walk (non-stationary) process.*

Assumption 2: *The non-fundamental component of stock prices (e.g., "fads" component or "rational bubble" component) follows an first-order auto regression [AR(1)] process.*

These two presumptions are frequently employed in finance literature [see, for example, Fama and French (1988a) and Porterba and Summers (1988)]. I empirically test the validity of both assumptions in a later section. However, unlike Porterba and Summers (1988), I do not impose the stationarity assumption on the non-fundamental component. I show in Lemma 1 below that these two assumptions together with Proposition 1 necessitate the stationarity of the non-fundamental component of stock prices.

Proposition 1: (see Appendix III for proof) *Under Assumptions 1 and 2, the fundamental component (p_t^f) and the non-fundamental component (p_t^{nf}) of stock prices are given as follows:*

$$E_t[p_{t+1}^f] = \rho p_{t+1} + (1-\rho)d_{t+1} - er_{t+1} - i_{t+1} + k \qquad (3.13)$$

$$E_t[p_{t+1}^{nf}] = (1-\rho)p_{t+1} - (1-\rho)d_{t+1} + er_{t+1} + i_{t+1} - k \qquad (3.14)$$

Proposition 1 provides a testable economic model of the fundamental value of stock prices. It states that the fundamental value of stock prices in period t+1 is a linear combination of that period's log of stock prices, the log of dividends, nominal interest rates and excess returns.

I also show from Lemma 1 below that the stock price behavior is better captured by a "fads" hypothesis than by the "bubble" hypothesis (consistent with Lee's (1998) empirical findings).

> **Lemma 1:** *Under Assumption 1 and 2, it holds that the non-fundamental component (p_t^{nf}) obtained from Proposition 1 is stationary.*

Since the non-fundamental component of stock prices is stationary in a "fads" scenario but non-stationary in a "bubble" scenario, results from Lemma 1 favor the "fads" hypothesis. Since p_{t+1}^{nf} is stationary, it follows that p_{t+1}^{nf} is represented by the cointegrating vector [p_{t+1}, d_{t+1}, i_{t+1}, er_{t+1}] with coefficients [(1-ρ), -(1-ρ), 1, 1] and a long-run intercept equal to (-k). Normalizing on prices, the cointegrating vector would be [1, -1, 1/(1-ρ), 1/(1-ρ), -k/(1-ρ)].

The system proposed in Proposition 1 extends the Campbell and Shiller (1988a) dividend-price ratio model in three directions: (a) it provides the stochastic models for both fundamental values and the non-fundamental components of stock prices, whereas the Campbell and Shiller model does not separate stock prices into such components; (b) it also allows for the possibility of both fads and rational bubbles and yet explicitly suggests that the fads model may better explain the price behavior; (c) using a nominal version of the model, I allow for a time-varying inflation premium to determine the fundamental value of stock prices.

CHAPTER 4

DATA AND METHODOLOGIES

Section 1: Data

Data for stock prices and dividends, starting in 1926, are taken from various issues of Standard and Poor's Statistical Service Security Price Index Record. The pre-1926 data counterparts are obtained from Cowles (1939). The interest rates are annual returns on four- to six- month commercial paper (six-month starting in 1980), rolled over in January and July. The interest rate data starting at 1938 are from the *Federal Reserve Bulletin*, and the pre-1938 data are culled from Macaulay (1938). Excess stock returns are the differentials between exact gross stock returns defined in equation (3.3) and nominal interest rates. In this dataset, the linear parameters of Taylor's approximation, ρ and k [see equation (3.7) and (3.8)] are computed to be 0.956 and 0.182, respectively.

The same data set has been used by numerous researchers in the volatility test literature [e.g. Shiller (1981), Campbell and Shiller (1988a, 1988b and 1989) and Lee (1996b, 1998)]. These data can be found in the Appendix II of this thesis.

Section 2: The Econometric Model of Stock Prices

I utilize the Gonzalo-Granger (1995) decomposition method and the Johansen cointegration technique to empirically identify the fundamental and non-fundamental components of stock prices. I show below that my econometric model coincides with the theoretical model in Proposition 1.

I decompose stock prices into a fundamental component (p_t^f), which is a linear combination of the non-stationary series, and a non-fundamental component (p_t^{nf}), which is a temporary I(0) process.

$$p_t = p_t^F + p_t^{NF} \tag{4.1}$$

Let Z_t be a n-dimensional vector of stock prices p_t and other fundamental variables such as d_t, i_t, and er_t.

$$Z_t = \begin{bmatrix} p_t \\ d_t \\ i_t \\ er_t \end{bmatrix}$$

If Z_t is a non-stationary vector, there might be a possibility that some linear combinations of Z_t become stationary (i.e., Z_t is cointegrated). Assume that the rank of the cointegration among Z_t is r [there exists a matrix of $\beta_{n \cdot r}$ of rank r, such that $\beta' Z_t$ is I(0)]. According to Granger's error-correction representation theorem, the vector Z_t has an ECM representation.

$$\Delta Z_t = \alpha\beta' Z_{t-1} + \sum_{i=1}^{\infty} \Gamma_i \Delta Z_{t-i} + \xi_t \tag{4.2}$$

where Δ=I-L, with L the lag operator; β is (n×r) coefficient matrix of cointegrating vectors; α is the (n×r) adjustment coefficient matrix; and Γ_i is (n×n) matrix of parameters reflecting the short-run structure; ξ_t is $IIN_n(0,\Lambda)$. The elements of Z_t can be explained in terms of (n-r) number of I(1) variables, f_t, called "common factors", plus some I(0) components.

$$\underset{n\times 1}{Z_t} = \underset{n\times k}{A_1} \underset{k\times 1}{f_t} + \underset{n\times 1}{\widetilde{Z}_t} \tag{4.3}$$

where k=n-r, A_1 is a loading matrix. Gonzalo and Granger (GG, 1995) have demonstrated that $A_1 f_t$ and $\widetilde{Z}_t$ form a permanent-temporary decomposition if the f_t's are linear combinations of the variables in Z_t, and the common factors f_t are identified. The common factors f_t can be identified in the ECM:

$$f_t = \alpha_\perp' Z_t \tag{4.4}$$

where $\alpha_\perp$ is (k×n) and $\alpha'_\perp \alpha=0$. Further, as discussed above, the fundamental component of stock prices is basically a permanent component and a linear combination of the fundamental variables. Once the common factors of f_t are identified, inverting the matrix $(\alpha_\perp, \beta)'$, I obtain the P-T decomposition of Z_t proposed by Gonzalo and Granger (1995).

$$\underset{n\times 1}{Z_t} = \underset{n\times k}{A_1} \underset{k\times 1}{\alpha'_\perp Z_t} + \underset{n\times r}{A_2} \underset{r\times n}{\beta'} \underset{n\times 1}{Z_t} \tag{4.5}$$

where the factor loadings $A_1 = \beta_\perp(\alpha'_\perp \beta_\perp)^{-1}$ and $A_2 = \alpha(\beta'\alpha)^{-1}$. Here $A_1 \alpha'_\perp Z_t$ is the permanent component in the system, and $A_2 \beta' Z_t$ is the I(0) temporary component, which can be interpreted as a deviation from the permanent trend.

As shown Proposition 1 in Chapter 3, the fundamental component is a linear combination of stock prices, dividends, expected inflation, and excess returns. Moreover, the fundamental and non-fundamental components are I(1) and I(0) processes, respectively. According to GG's (1995) Proposition 2, if the fundamental and non-fundamental components mimic the permanent and transitory components of stock prices, respectively, the Gonzalo and Granger permanent/transitory decomposition method may be legitimately used to identify the fundamental and non-fundamental components of stock prices in the ECM (4.2)[7].

Of course, the permanent component A_1f_t is not necessarily a random walk component of Z_t.[8] If the GG procedure is the appropriate method to decompose stock prices into fundamental and non-fundamental components, then the temporary component in Z_t should have the same estimated coefficients as predicted from Proposition 1 in Chapter 3. I do not attempt to test whether the coefficients of the common factors coincide with the theoretical parameters predicted Proposition 1 in Chapter 3. As long as the temporary component corresponds to the non-fundamental component of stock prices, one can infer that the fundamental component is embedded in the permanent components in equation (4.5). Thus, testing for the parameter restrictions in the temporary component is equivalent to testing the hypothesis that the GG permanent-temporary components empirically form the fundamental and non-fundamental components of stock prices.

Under this scenario, the non-fundamental component (p_t^{nf}) should correspond to the normalized cointegrating vector $\beta' Z_t$. I test two restrictions on the non-fundamental component of stock prices [equation (3.14) of Proposition 1]:

H_1: The log of prices and log of dividends have the parameter relationship (1, -1), and

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[7]Proposition 2 of Gonzalo and Granger (1995) states: In the factor model (4.2), the following conditions are sufficient to identify the common factors f_t:

(1) f_t are linear combinations of Z_t

(2) A_1f_t and $\tilde{Z}_t$ form a permanent/temporary decomposition.

[8]In fact, Gonzalo and Granger (1995) demonstrate that the random walk component of common factors f_t corresponds to Stock and Watson's (1988) common trend.

H_2: Nominal interest rates and excess stock returns have the parameter relationship (1, 1).

My tests examine the following three alternative present-value models, each with different elements in the fundamental component:

Model 1: a standard log linear present-value model with a constant discount rate, and dividends (d) are the only fundamental variables for stock prices: $Z_t = [p_t, d_t]'$.

Model 2: an expanded log linear model with time-varying excess returns (er): $Z_t = [p_t, d_t, er_t]'$.

Model 3: a further augmented model that adds a time-varying expected inflation (i): $Z_t = [p_t, d_t, er_t, i_t]'$.

Comparison of these models can identify the augmented effects of time-varying excess returns and time-varying expected inflation on stock prices. If news about future cash flows (dividends) does not explain all variations in stock prices [as documented by Shiller (1981) and Campbell (1991)], then stock prices should exhibit a significant non-fundamental component in Model 1. If excess stock returns account for the remainder of stock price variations, the non-fundamental component of stock prices should lose significance in Model 2. Otherwise, Model 3 investigates whether the time-varying expected inflation can reduce the non-fundamental component to approach insignificance. Of course, evidence for market irrationality (rejection of the present-value model) may be inferred if the augmented model (3) still fails to produce insignificance non-fundamental components.

Section 3: Estimation and Hypothesis Testing

Subsection 3.1: Testing for Cointegration and Estimating Common Factors

A brief review of techniques used to estimate cointegration properties of time series is provided to motivate the estimation of the common factors and to introduce notations.

The cointegrating matrix β, as has been shown by Johansen (1988, 1991) and Johansen and Juselius (1990), can be estimated as the eigenvectors associated with the r largest, statistically significant eigenvalues of the following equation:

$$|\lambda S_{kk} - S_{ko}S_{oo}^{-1}S_{ok}|=0 \tag{4.6}$$

where S_{oo} and S_{kk} are the residual moment matrices from the least square regressions of ΔZ_t and Z_{t-k} on ΔZ_{t-1}, ..., ΔZ_{t-k+1}, respectively, and S_{ok} is the cross-product moment matrix of the residuals. The maximum likelihood function is given by:

$$L_{\max}^{-2/T} = |S_{oo}|\prod_{i=1}^{r}(1-\hat{\lambda}_i) \tag{4.7}$$

where $\hat{\lambda}_i$ is the ith largest eigenvalue of equation (4.6).

The number of cointegrating vectors is determined by comparing the values of the likelihood function for the unrestricted model (r=n) and the restricted model ($r=r_0$). The resulting log-likelihood ratio is called the 'trace statistic' and is given by:

$$LR_{trace} = -T\sum_{i=r_0+1}^{n}\ln(1-\hat{\lambda}_i). \tag{4.8}$$

where T is the sample size. The distribution of the log-likelihood ratio test statistic is not given by the usual χ^2 distribution but rather a multivariate version of the Dickey-Fuller test statistic. A second test to determine the number of cointegrating vectors is to compare the likelihoods of the restricted models $r=r_0$ and $r=r_0+1$. This statistic is called the 'maximum eigenvalue statistic' and is given by[9]

$$LR_{\max} = -T\ln(1-\lambda_{r_0}) \tag{4.9}$$

Sometimes, these two Johansen statistics may give contradictory results as to how many cointegrating vectors exist. The Monte Carlo experiments reported in Cheung and Lai (1993) suggest that between Johansen's two LR tests for cointegration, the trace test shows more robustness to both skewness and excess kurtosis in the residuals than the maximal eigenvalue test. Since stock prices tend to have excess kurtosis and skewness as suggested by previous literature, I place greater weight on the trace test and only report trace statistics for cointegration implication and choose to report only Johansen's trace statistics to determine the number of cointegrating vectors.

Under the hypothesis of cointegration, the maximum likelihood estimator of $\alpha_\perp$ can be found by solving the equation (4.7). The choice of $\hat{\alpha}_\perp$ is the eigenvector associated with the (p-r) smallest eigenvalues.

Subsection 3.2: Hypothesis Testing
in the Cointegrated Framework

A. Testing Hypotheses in the Cointegrating Relationship

The non-fundamental component is the cointegrating relationship among the prices, dividends, expected inflation, and excess returns. I apply hypothesis tests to empirically

[9]Reimers (1992) suggests replacing T in (4.8) and (4.9) with (T-pL) to adjust for small sample bias, where L is the number of VAR lag length.

identify the non-fundamental component of stock prices. The tests are conducted by imposing restrictions and then testing whether the cointegrating vector is empirically identified. Results show that the empirically identified non-fundamental component is compatible with that suggested by the theoretical derivation in the previous section.

Usually economic theory implies certain long-run relations between variables. Johansen (1988) and Johansen and Juselius (1990) show how to estimate the cointegrating matrix under linear restrictions. A restriction on β can be formed as

$$\beta = H\,\varphi \tag{4.10}$$

where H is a (p×s) restriction matrix and φ is a (s×r) matrix of coefficients. Under this hypothesis the maximum likelihood estimator of φ can be found as the eigenvectors associated with the r largest, statistically significant eigenvalues of the equation:

$$\left|\tilde{\lambda} H' S_{kk} H - H' S_{ko} S_{oo}^{-1} S_{ok} H\right| = 0 \tag{4.11}$$

with the likelihood function:

$$L_{\max}^{-2/T} = |S_{oo}| \prod_{i=1}^{r} (1 - \tilde{\lambda}_i)^{-1} \tag{4.12}$$

The likelihood ratio statistic of the hypothesis (4.10) is

$$T \sum_{i=1}^{r} \ln \frac{(1 - \tilde{\lambda}_i)}{(1 - \hat{\lambda}_i)} \tag{4.13}$$

and distributed as standard χ^2 with r(p-s) degrees of freedom.

B. Testing Hypotheses in Common Factors

Restrictions on the common factors are formed and tested. Let G be (n×m) restriction matrix and ϕ be (m×(n-r)). Then the hypotheses on $\alpha_\perp$ can be formed as:

$$\underset{n\times(n-r)}{\alpha_\perp} = \underset{n\times m}{G}\ \underset{m\times(n-r)}{\phi} \qquad \text{with } (n\text{-}r) \le m \le n \tag{4.14}$$

The estimates of ϕ are the eigenvectors associated with the m smallest eigenvalues of the following problem:

$$\left|\theta G' S_{oo} G - G' S_{ok} S_{kk}^{-1} S_{ko} G\right| = 0 \tag{4.15}$$

with the likelihood function:

$$L_{\max}^{-2/T} = \left|S_{oo} - S_{ok} S_{kk}^{-1} S_{ko}\right| \{\prod_{i=r+1}^{p} (1 - \tilde{\theta}_{i+m-p})\}^{-1} \tag{4.16}$$

The likelihood ratio statistic of the hypothesis (4.14) is given by:

$$T \sum_{i=r+1}^{p} \ln \frac{(1 - \tilde{\theta}_{i+m-p})}{(1 - \hat{\lambda}_i)} \tag{4.17}$$

and distributed again as standard χ^2 with (p-r)(p-m) degrees of freedom.

C. Testing the Significance of the Non-Fundamental Component

In order to test whether the non-fundamental component of stock prices is significantly different from zero, I re-write ECM (4.2) as follows:

$$\begin{bmatrix} \Delta p_t \\ \Delta d_t \\ \Delta i_t \\ \Delta er_t \end{bmatrix} = \begin{bmatrix} \alpha_1 \\ \alpha_2 \\ \alpha_3 \\ \alpha_4 \end{bmatrix} \beta' Z_{t-1} + \sum_{i=1}^{\infty} \Gamma_i \Delta Z_{t-i} + \xi_t \tag{4.18}$$

Since the error-correction term ($\beta' Z_{t-1}$) is embedded in the transitory non-fundamental component, then testing whether the temporary component is significant amounts to testing the null of $\alpha_1 = 0$. If $\alpha_1 = 0$, then the temporary component of Z_t will have no short-run effect on stock prices[10]. Thus, I may conclude that stock prices do not have a significant temporary (or, equivalently, non-fundamental) component. The test procedure is described in the following proposition (see the Appendix III for a proof).

> **Proposition 2:** *Under the assumption of one cointegrating vector (r=1), testing the significance of the temporary component of a series in Z_t in the Gonzalo and Granger (1995) framework is equivalent to testing weak exogeneity of the series with respect to α and β in the Johansen and Juselius (1990) framework.*

The test for weak exogeneity of stock prices p_t, in the system is conducted by placing zero restrictions on α_1 to give a new restricted model, and then using a likelihood ratio test involving the restricted and unrestricted models to ascertain whether the restrictions are valid. The form of the restrictions is determined by specifying a (n×m) matrix A of linear restrictions where (n-m) equals the number of row restrictions imposed on α, such that the null hypothesis amounts to testing whether $\alpha = A\alpha_0$. Imposing the restrictions reduces α to a (m×n) matrix α_0. It is also useful to note that these same restrictions in A could be imposed by specifying a (n×(n-m)) matrix B such that $B'\alpha = 0$. Clearly, B must be orthogonal to A, that is, $B'A = A_\perp' A = 0$. Both matrices A and B are used in the mechanics of restricting the

[10]Note that the temporary component can only have short-run effects on Z_t.

Johansen reduced rank regression model, thereby obtaining (n-1) new eigenvalues $\hat{\lambda}_i^{*}$ for the restricted model which are used in the following LR test statistic:

$$-2\log(Q) = T\sum_{i=1}^{r}\log\left\{\frac{(1-\hat{\lambda}_i^{*})}{(1-\hat{\lambda}_i)}\right\} \tag{4.19}$$

This test statistic is compared with the χ^2-distribution with (r×(n-m)) degrees of freedom in order to obtain the significance level for rejecting the null hypothesis.

CHAPTER 5

EMPIRICAL RESULTS

Section 1: Unit Root Test Results

I test for the presence of unit roots in all four variables in my models, and Table 1 reports the results from the Augmented Dickey-Fuller (1979, ADF) test and the Weighted Symmetric (WS) test. Pantula *et al.* (1994) argue that the WS procedure is the most powerful unit root test against several alternatives, including the ADF test. I allow up to 12 lags in the testing equations, choosing the proper lags based on the Akaike Information Criterion (AIC) with the requirement of white noise residuals.

As can be seen in Table 1, the null of non-stationarity is rejected for all four variables in levels, but not in first-differences. Therefore, each variable in Z is ~I(1) (i.e. first-difference stationary). Engle and Granger (1987) demonstrate that it is possible for the levels of these variables to cointegrated.[11]

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[11] One may notice the difference between my finding and Lee's (1998) in the stationarity status of excess returns. Lee (1998) finds excess returns to be stationary in levels. To justify this, I test the stationarity of nominal gross stock returns and find them to be level-stationary. The excess stock returns, which are the difference between the stationary gross returns and non-stationary nominal interest rates, cannot be stationary in levels.

Table 1 Unit Root Test Results
(Cowles/S&P 500 Annual Data: 1871-1997)

Levels				
	p_t	d_t	i_t	er_t
Weighted Symmetric Test	-0.64 [6]	-0.60 [6]	-2.04 [4]	-2.65 [10]
Augmented Dickey-Fuller Test	-0.30 [6]	-1.24 [6]	-2.12 [4]	-2.41 [10]
First-Differences				
	Δp_t	Δd_t	Δi_t	Δer_t
Weighted Symmetric Test	-4.79** [6]	-5.86 ** [5]	-4.59 ** [7]	-4.81** [11]
Augmented Dickey-Fuller Test	-5.77 ** [5]	-5.93 ** [5]	-4.54 ** [7]	-4.81** [11]

Notes:
Variables are defined as follows: p_t is the log stock price, d_t is the log dividend, i_t is the nominal interest rate, and er_t is excess stock returns relative to short-term debt. Annual data over the period of 1871-1997 are used. A time trend is included in the unit root test regression. The numbers in brackets are proper lags generally selected by AIC. The ** indicates rejection at the 5% level.

Section 2: Cointegration Relationships

In a cointegrating framework, I can test whether the non-fundamental component of stock prices [equation (3.14) in Proposition 1] contains parameters similar to those suggested by the data in the cointegrating relationship. I employ the Johansen (1988) test to examine cointegration in the three alternative models outlined previously. Unlike Lee (1998), I do not impose *a priori* restrictions on the cointegrating parameters (such as the price-dividend ratio). Rather, I allow the data to determine the cointegrating parameters and then perform formal tests of parameter restrictions. Table 2 displays the results from the Johansen test. The orders of VARs in these tests are jointly determined by the AIC and the requirement of white-noise residuals.

The results suggest that there is one non-zero cointegrating vector in each model[12]. As discussed earlier, I hypothesize that the GG decomposition method should fit the data adequately. If true, the parameters in the cointegrating vector $\beta' Z_t$ should be consistent with the theoretical restrictions described in H_1 and H_2. Thus, I proceed to test the restrictions in the cointegrating vector using the Johansen and Juselius (1992) likelihood ratio test. The restrictionhypothesis can be formulated as in (4.12), with a properly formulated H-restriction matrix. Table 3 reports the formalized H matrices and the χ^2 test results for the three alternative models.

As the table shows, except for Model 3, the implied coefficient restrictions are rejected across the other two models. Such a finding suggests that the stationary price-dividend ratio holds only when expected inflation is also allowed to be time-varying. This

[12]Following Cheung and Lai (1993), I rely on the trace (as opposed to the maximal eigenvalue) test since it is relatively insensitive to skewness and excess kurtosis in the residuals. However, results from the maximal eigenvalue test do not alter the conclusions from the trace test.

suggests that in order to satisfy the stationarity condition of the dividend-price ratio model for stock prices, expected inflation must be augmented into the model.

In addition to the price-dividends ratio restriction, Proposition 1 also requires equality between the coefficients of excess returns and expected inflation (hypotheses H_2). Table 3 also displays the joint test results for this restriction together with the price-dividends ratio restriction (see Model 3). Looking closely at the restricted β-coefficients, I find that the estimated values coincide with the theoretical coefficients of the non-fundamental component in equation (3.14) where the parameters ρ and k take the values 0.956 and 0.181, respectively. More specifically, the normalized cointegrating vector in equation (3.14) is expected to be $[p_t - d_t + 22.73i_t + 22.73er_t - 4.14]$, and this theoretical vector is very close to what is estimated in the last three rows (Model 3) of Table 3. This also corroborates my earlier result that the fundamental and non-fundamental components represent respectively the permanent and temporary components of stock prices. However, Models 1and 2 exhibit sufficient departure from the theoretical parameters of equation (3.14). Without the time-varying expected inflation, the non-fundamental (and hence the fundamental) component would be incorrectly specified. Thus, Model 3 appears to provide the best fit to the data.

Recall that my derivation of fundamental and non-fundamental components is based on two assumptions; namely, that (a) the fundamental component of stock prices follows a random walk, and (b) the non-fundamental component follows an AR(1) process. Both assumptions are required for the GG decomposition method to work. To test the empirical validity of the first assumption, I use Model 3 as the benchmark model for these tests.

Table 2 Johansen Cointegration Tests
(Cowles/S&P 500 Annual Data 1871-1997)

	Trace Statistics: $LR_{trace} = -T \sum_{i=r_0+1}^{p} \ln(1 - \hat{\lambda}_i)$				**Cointegrating Vector (β'):**
	H_0: r=0	**H_0: r≤1**	**H_0: r≤2**	**H_0: r≤3**	**(normalized β')**
Model 1 $Z_t = [p_t, d_t, k]'$	22.04**	5.11			[1, -1.29, -3.47]
Model 2 $Z_t = [p_t, d_t, er_t, k]'$	32.05*	15.93	5.40		[1, -3.37, 94.37, -5.00]
Model 3 $Z_t = [p_t, d_t, er_t, i_t, k]'$	52.37*	23.72	11.32	1.85	[1, -1.02, 25.12, 24.46, -4.25]

Notes:
The trace statistics are compared to the critical values from Osterwald-Lenum (1992, Table 1*). An * indicates rejection of the null of no-cointegration at the 10% significance level, while ** indicates rejection at the 5% level. The proper (AIC-selected/white-noise) lags in the VARs are 5, 2, and 3 for Models 1-3, respectively.

Table 3 H-Restriction Matrices and Johansen's (1991) χ^2-Tests
(Cowles/S&P 500 Annual Data 1871-1997)

	H- Restriction Matrices (transposed)	**χ^2 tests (p-values)**
Model 1 $Z_t = [p_t, d_t, k]'$	$\begin{bmatrix} 1 & -1 & 0 \\ 0 & 0 & 1 \end{bmatrix}$	$\chi^2(1) = 6.86$ **
Model 2 $Z_t = [p_t, d_t, er_t, k]'$	$\begin{bmatrix} 1 & -1 & 0 & 0 \\ 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 1 \end{bmatrix}$	$\chi^2(1) = 4.84$ **
Model 3 $Z_t = [p_t, d_t, i_t, er_t, k]'$ with price-dividends ratio restriction only	$\begin{bmatrix} 1 & -1 & 0 & 0 & 0 \\ 0 & 0 & 1 & 0 & 0 \\ 0 & 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 0 & 1 \end{bmatrix}$	$\chi^2(1) = 2.29$ $\beta' = [1, -1, 26.18, 25.72, -4.30]$
Model 3 $Z_t = [p_t, d_t, i_t, er_t, k]'$ with combined restrictions of price-dividends ratio and the equality of the two discount factors.	$\begin{bmatrix} 1 & -1 & 0 & 0 & 0 \\ 0 & 0 & 1 & 1 & 0 \\ 0 & 0 & 0 & 0 & 1 \end{bmatrix}$	$\chi^2(2) = 3.38$ $\beta' = [1, -1, 25.33, 25.33, -4.26]$

Notes: The ** indicates rejection at the 5% level.

First, I apply the Lo and MacKinlay's (1988) variance ratio test on the derived fundamental component, p_t^f over 1 to 32 intervals (results are reported in Table 4). Both the asymptotic normal Z test and the heteroscedasticity-consistent Z* test results suggest that the fundamental component of stock price follows a random walk.

I check for the stationarity of the two components of stock prices. The ADF and WS tests are employed, and the results (reported in Table 5) suggest that the fundamental component is first-difference stationary, and the non-fundamental component is level-stationary.

Finally, to test the second assumption, I apply a Ljung-Box Q-test on the AR(1) residuals of the non-fundamental component. The Ljung-Box Q-statistics generally suggest absence of autocorrelation up to six annual lags. These results (reported in Table 6) support my conjecture that the non-fundamental component obeys an AR(1) process.

Table 4 Variance Ratio Tests on the Fundamental Component of Stock Prices in Model 3 (Cowles/S&P 500 Annual Data: 1871-1997)

Intervals	Z-tests	p-values for Z tests	Z* tests	p-values for Z* tests
2	0.73	0.46	0.43	0.67
3	-0.25	0.80	-0.17	0.87
4	-0.31	0.76	-0.23	0.82
5	-0.49	0.62	-0.39	0.70
6	-0.87	0.39	-0.73	0.46
7	-1.07	0.28	-0.95	0.34
8	-1.03	0.30	-0.97	0.33
9	-1.10	0.27	-1.09	0.28
10	-0.95	0.34	-0.98	0.33
11	-0.77	0.44	-0.83	0.41
12	-0.69	0.49	-0.77	0.44
13	-0.76	0.45	-0.88	0.38
14	-0.78	0.43	-0.93	0.35
15	-0.70	0.48	-0.86	0.39
16	-0.78	0.44	-0.97	0.33
17	-0.84	0.40	-1.08	0.28
18	-0.69	0.49	-0.90	0.37
19	-0.77	0.44	-1.04	0.30
20	-0.78	0.43	-1.08	0.28
21	-0.57	0.57	-0.80	0.43
22	-0.60	0.55	-0.86	0.39
23	-0.70	0.48	-1.03	0.31
24	-0.73	0.46	-1.08	0.28
25	-0.62	0.53	-0.93	0.35
26	-0.61	0.54	-0.93	0.35
27	-0.61	0.54	-0.95	0.34
28	-0.64	0.52	-1.01	0.31
29	-0.63	0.53	-1.01	0.31
30	-0.61	0.54	-0.99	0.32
31	-0.53	0.60	-0.87	0.39
32	-0.73	0.46	-1.22	0.22

Notes: The Z test statistics assume homoscedasticity, whereas the Z* test statistics are heteroscedasticity-consistent.

Table 5 Unit Root Tests on the Derived Fundamental and Non-Fundamental Components of Stock Prices (Cowles/S&P 500 Annual Data: 1871-1997)

Levels		
	Fundamental Component	Non-Fundamental Component
Weighted Symmetric Test	-0.78 [6]	-5.08** [5]
Augmented Dickey-Fuller Test	-0.47 [6]	-5.22** [5]
First-Differences		
	Fundamental Component	
Weighted Symmetric Test	-5.18** [5]	
Augmented Dickey-Fuller Test	-5.34** [5]	

Notes:
The fundamental and non-fundamental components of stock prices are derived from Proposition 1 where ρ and k take the values 0.956 and 0.182, respectively. Annual data over the period of 1871-1997 are used. A time trend is included only in the unit root test regression for the levels of the fundamental component. The numbers in brackets are proper lags generally selected by AIC. An ** indicates rejection at the 5% level.

Table 6 Ljung-Box Q-Test on the AR(1) Residuals of
the Non-Fundamental Component of Stock Prices in Model 3
(Cowles/S&P 500 Annual Data: 1871-1997)

Lags	**Ljung-Box Q-statistic**	**p-values**
1	0.05	0.82
2	4.29	0.12
3	7.33	0.06
4	7.50	0.11
5	9.07	0.11
6	10.17	0.12

Section 3: Exclusion Tests Within the Fundamental and Non-Fundamental Components

My results thus far suggest that the present-value model of stock prices should be augmented by a time-varying expected inflation. In this section, I further investigate whether previous rejections of the present-value model are caused by their neglect of this important variable. Based on Model 3, I examine the significance of each variable in the fundamental and non-fundamental components of stock prices. To explore the importance of each series in the fundamental components, I perform exclusion tests on the common factors (f_t) using the GG (1995) approach. The null hypothesis is that a particular series can be excluded from the fundamental component [see (4.16) in Section 3 of Chapter 4]. The exclusion test on the non-fundamental component ($\beta' Z_t$) again follows the Johansen and Juselius (1992) procedure. Table 7 reports the results of these exclusion tests for Model 3.

Table 7 Exclusion Tests in Fundamental and Non-Fundamental Components in Model 3 (Cowles/S&P 500 Annual Data 1871-1997)

Fundamental Component: $\chi^2(3)$ test				**Non-Fundamental Component: $\chi^2(1)$ test**			
p_t	d_t	i_t	er_t	p_t	d_t	i_t	er_t
7.23*	12.07**	6.94*	7.34*	12.10**	12.06**	15.27**	14.75**

Notes: An * indicates rejection of exclusion at the 10% significance level, while ** indicates rejection at the 5% level.

As can be seen from the table, none of the variables can be excluded from either component. This suggests that all four variables (including expected inflation) should be maintained in the model to avoid a serious loss of information about fundamentals. This finding provides further support to my earlier contention that Model 3 is the best time-series model for stock prices among alternative models in the sense that it contains the necessary information underlying the fundamentals.

Section 4: Testing the Significance of the Non-Fundamental Component

I test the significance of the non-fundamental component of stock prices using Proposition 2 in Chapter 4. Table 8 displays the test results for the significance of the non-fundamental component in the three models. To check the robustness of Model 3, I also report the test results jointly with the two restrictions on the cointegrating vector. Models 1 and 2 exhibit significant non-fundamental components. In contrast, when I control for time-varying expected inflation (i_t) in Model 3, the non-fundamental component loses significance. This implies that previous findings of significant deviations from the fundamental value appear mainly due to overlooking expected inflation. Thus, I postulate that the "unexplained factors" discussed in Campbell and Shiller (1988a) could very well be the time-varying inflation premium.

I must caution, however, that these results for the full sample period (1871-1997) appear to be sensitive to the particular lag specifications employed in the test. For example, using lag 5 (instead of 3) for Model 3, I find that the non-fundamental component is significant at least at the 10% level with and without theoretical restrictions [$\chi^2(3)$=8.70 and $\chi^2(1)$= 3.37]. Such a lack of model robustness is of course disappointing and could very well be due to possible regime changes, particularly after World War II. For instance, Fama and French (1988) and Kim, Nelson and Startz (1991) report that the mean reversion of stock prices is mostly due to the pre-WW II period, and Campbell (1991) argues that stock market predictability becomes clear only in the post-1950s. Consequently, I turn my attention next to testing the structural stability of the cointegrating vector (non-fundamental component).

Table 8 Testing the Significance of Non-Fundamental Components
(Cowles/S&P 500 Annual Data 1871-1997)

	Non-Fundamental Component ($\beta' Z_t$)	**χ^2 tests of significance (p-values)**
Model 1 $Z_t = [p_t, d_t, k]'$	$p_t - 1.29d_t - 3.48$	$\chi^2(1) = 11.82$ **
Model 2 $Z_t = [p_t, d_t, er_t, k]'$	$p_t - 3.37d_t + 94.37er_t - 5.00$	$\chi^2(1) = 4.37$ *
Model 3 $Z_t = [p_t, d_t, i_t, er_t, k]'$ (with unrestricted β)	$p_t - 1.02d_t + 25.06i_t + 24.37er_t - 4.25$	$\chi^2(1) = 0.12$
Model 3 $Z_t = [p_t, d_t, i_t, er_t, k]'$ (with restricted β)	$p_t - d_t + 25.50i_t + 25.50er_t - 4.27$	$\chi^2(3) = 4.83$

Notes: An * indicates rejection of the null of zero non-fundamental component of stock prices at the 10% significance level, while ** indicates rejection at the 5% level.

Section 5: Stability of the Cointegrating Vector

In light of the apparent sensitivity of my results to reasonable model modifications (such as lag lengths), it is important to check whether the non-fundamental component of stock prices contains stable parameters. In particular, I explore whether the inferences I make for the cointegration ranks and cointegrating vectors are stable in pre- and post-WW II periods. To address this issue, I employ the Hansen-Johansen (1993, 1998) recursive analysis for testing whether the cointegration inferences are sample dependent. I test Model 3 and focus on three related hypotheses:

H_3: *The trace test statistics (cointegration ranks) are sample independent.*

H_4: The eigenvalue corresponding to the non-zero cointegrating vector is stable over time.

H_5: *The estimated coefficients of the non-fundamental component (restricted β-vector) are sample independent.*

I use 1872-1919 (48 observations) as the base period for the recursive analysis and extend the sample by adding the succeeding observations one by one giving sample paths of 1920-1997 (78 observations). Figure 1 plots the trace test statistics obtained from recursive estimation against time. All statistics are scaled by the (asymptotic) 10% critical value whereby values greater than unity imply rejection of the null hypothesis of stable cointegration ranks at the 10% significance level.

Figure 2 shows the time paths of the non-zero eigenvalue with 95%-confidence bands calculated using the Hansen and Johansen (1998) method. I observe a decrease in point estimates of the eigenvalue during 1947-1952, an evidence of weak constancy of the non-zero eigenvalue. The non-constancy of the eigenvalue affects the constancy of cointegration parameters.

Figure 3 plots the test statistics for constancy of cointegration parameters [derived in Hansen and Johansen (1993)] against time. The test statistics are scaled by the 5% significance level, whereby values greater than unity imply rejection of the null hypothesis of constant cointegration space. It can be seen that the values of the statistics begin to show noticeable increases towards the significance (unity) line only in the post-WW II period. Taken together, then, the recursive analysis suggests that the cointegration rank may not be sample dependent, but there seems to be a structural shift in the cointegrating parameters over the post-WW II period.

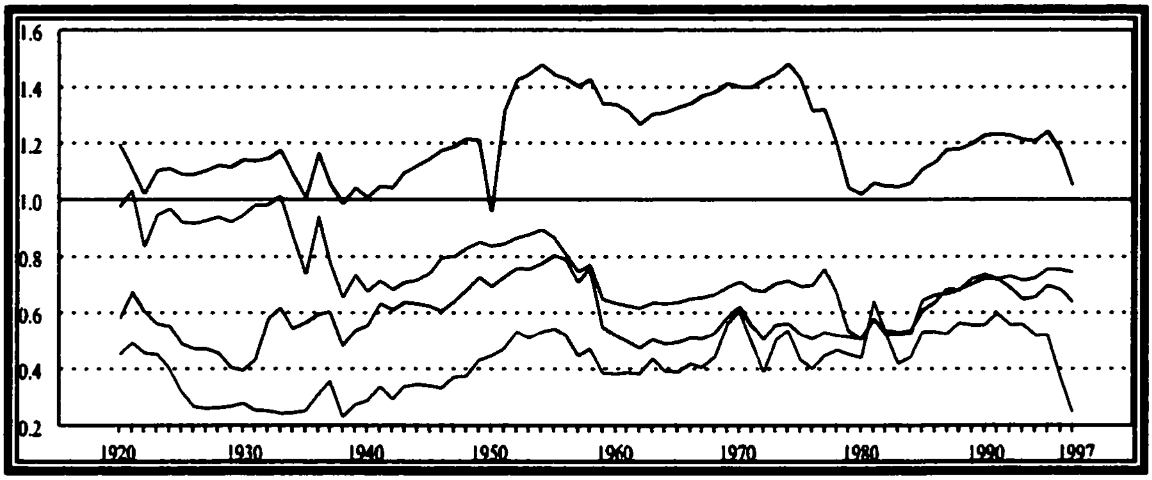

Figure 1: Test of Sample Dependence of the Cointegration Rank (Trace Test)

Notes: Figure 1 plots the trace test statistics obtained from recursive estimation against time. All statistics are scaled by the (asymptotic) 10% critical value whereby values greater than unity imply rejection of the null hypothesis of stable cointegration ranks at the 10% significance level.

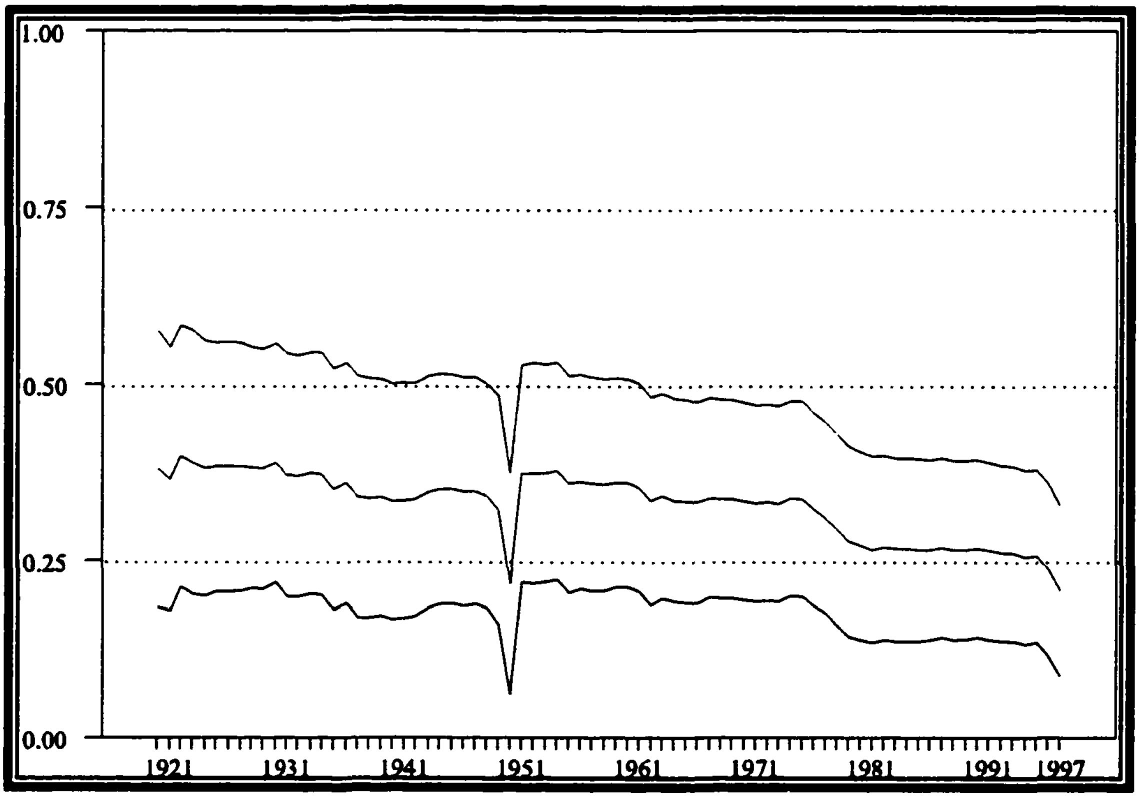

Figure 2: Test for Stability of the Non-Zero Eigenvalue

Notes: Figure 2 plots the time paths of the non-zero eigenvalue with 95% confidence bands calculated using Hansen and Johansen (1998) method.

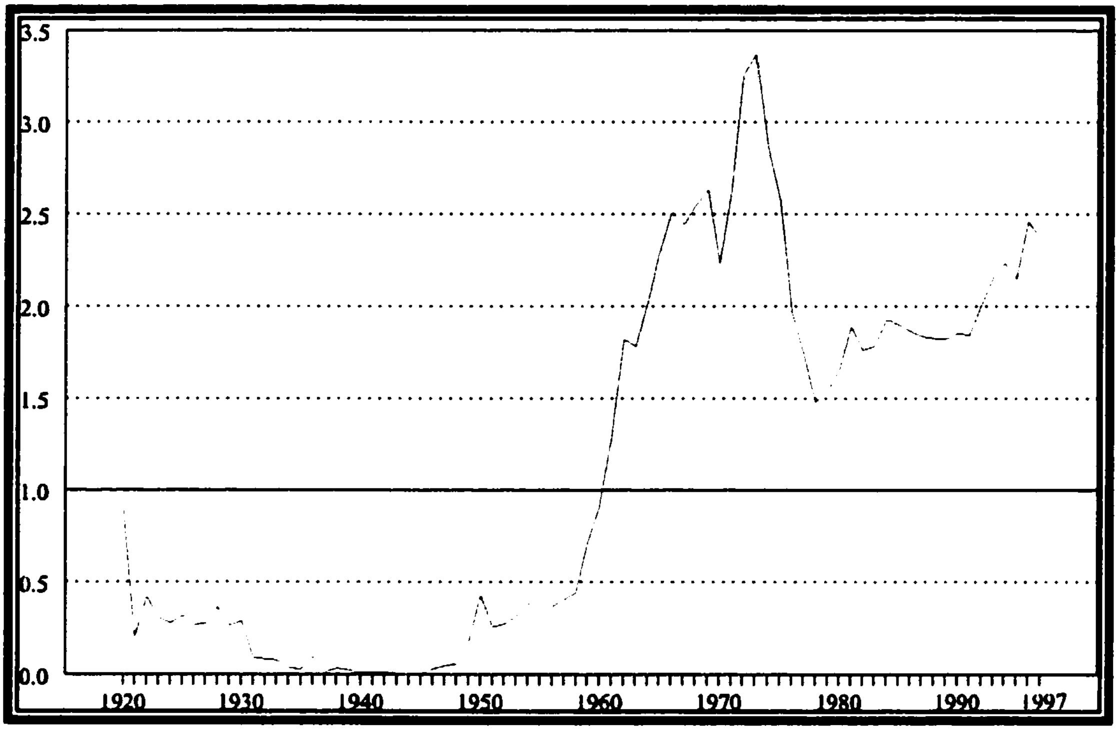

Figure 3: Test for Constancy of the Cointegrating Vector

Notes: Figure 3 plots the test statistics for constancy of cointegration vector against time. [see Hansen and Johansen (1993)].

Section 6: Sub-sample Analyses

I partition the sample into the pre-WW II period (1871-1940) and post-WW II period (1947-1997), and perform similar cointegration tests on Model 3. According to unit root tests, the four variables continue to be first-difference stationary in the two separate regimes. As is the case for the full sample period, there appears to be one non-zero cointegrating vector in the two separate samples. The Johansen trace test results are displayed in Table 9.

Table 9 Johansen Cointegration Tests of Model 3 for the Pre- and Post- WW II Periods (Cowles/S&P 500 Annual Data)

	Trace Statistics: $LR_{trace} = -T\sum_{i=r_0+1}^{p}\ln(1-\hat{\lambda}_i)$				**Cointegrating Vector (β):**
	H_0: r=0	**H_0: r≤1**	**H_0: r≤2**	**H_0: r≤3**	**(normalized β')**
(1871-1940) Model 3 $Z_t = [p_t, d_t, er_t, i_t, k]'$	50.18*	21.49	9.87	2.15	[1, -1.01, 20.09, 20.39, -4.00]
(1947-1997) Model 3 $Z_t = [p_t, d_t, er_t, i_t, k]'$	52.70*	22.73	10.45	2.46	[1, -1.07 20.85, 19.07, -4.18]

Notes: The trace statistics are compared to the critical values from Osterwald-Lenum (1992, Table 1*). An * indicates rejection of the null of no-cointegration at the 10% significance level. The VAR lags, selected to ensure the absence of serial correlation of the VAR, are 3 for both sub-periods.

Next, I test for the theoretical restrictions as outlined in hypotheses H_1 and H_2. For the sub-period 1871-1940, the joint restrictions are not rejected even at the 10% significance level [$\chi^2(2) = 1.46$, p-value = 0.48]. However, for the post-WW II period (1947-1997), the restriction test results provide the exact opposite conclusion [$\chi^2(2) = 6.11$ (p-value = 0.05)], suggesting the rejection of the joint restriction hypotheses H_1 and H_2. I use alternative lags and the rejection verdict is quite unwavering. When I test hypotheses H_1 and H_2 separately, both can be rejected at the 5% significance level. Specifically, the test statistic for H_1 is only $\chi^2(1) = 5.69$ (p-value = 0.02) and for H_2 is $\chi^2(1) = 5.55$ (p-value = 0.02). These results suggest that the fundamental and non-fundamental components may not be well-represented by my preferred four-variable econometric model 3 in the post-WW II period perhaps due to failure to incorporate other "unexplained factors". Only in the pre-WW II period does Model 3 explain well the fundamental and non-fundamental components of stock prices.

Finally, I formally test the statistical significance of the non-fundamental component of stock prices in the pre- and post-WW II periods. As to the first sub-sample, and with the unrestricted cointegrating vector, the test statistic is $\chi^2(1) = 0.01$. With the cointegrating vector restricted by hypotheses H_1 and H_2, the test statistic is $\chi^2(3) = 1.53$. Therefore, the null hypothesis that the non-fundamental component is zero could not be rejected for the pre-WW II period (alternative lags yielded similar results). Turning to the post-WW II period, the results clearly reject the null of an insignificant non-fundamental component of stock prices [$\chi^2(1) = 8.87$ and 10.32, respectively].

In summary, the results consistently suggest that stock prices are too volatile to be justified by the fundamentals including dividends, time-varying risk premiums and expected inflation for the post-WW II period. My augmented four-variable present-value model of fundamental values captures stock price variation quite well but only in the pre-WW II period. For the more recent post-WW II period, the augmented present-value model fails to fully account for price volatility. This finding, of course, is inconsistent with stock market rationality in recent years. Due to the substantial conviction on the part of most popular press and financial analysts regarding market rationality, one who argues otherwise does so at great risk and must provide sufficient proof. Therefore, in the next section, I further check the sensitivity of my conclusion to alternative measures of the basic variables and data frequency.

Section 7: Alternative Results

In this section, I explore whether my finding against market rationality in the post-WW II period is sensitive to the alternative measures of variables and data frequency. In particular, I measure stock prices, as in Campbell and Shiller (1988a), by the value-weighted NYSE index (instead of the Cowles/S&P 500) and measure interest rates by the 3-month Treasury bill rate (instead of the 4-6 month Commercial Paper rate)[13]. Much of the evidence brought to bear on the variance-bound tests of market efficiency is culled from annual data perhaps in order to avoid the bid-ask spread effect and non-synchronous trading [Shiller (1981), Campbell and Shiller (1988a), and Lee (1998)]. However, it may be argued that a shorter data frequency is preferred in tests of market efficiency. More importantly, the limited number of observations in the annual dataset for the post-WW II period may be more seriously subject to Flavin's (1983) critique of "small sample bias" of the variance bounds tests. To address this concern, I use quarterly data over the post-WW II period (1947:1 - 1997:4) which yields 204 observations.

The quarterly data series are compiled from the monthly figures whose sources are outlined as follows. The value-weighted NYSE index is obtained from the CRSP stock index file. Following Lee's (1995) method of compiling dividend series, I first denote nominal stock prices and dividends series as P_t and D_t, respectively. The value-weighted

[13]When extracting expected inflation from nominal interest rates, previous researchers typically used 3-month Treasury bills [see Fama and Schwert (1977) among others].

index returns including dividends are $RD_t = \{[(P_t + D_t)/P_{t-1}] - 1\}$, where dividends, D, is also obtained from the CRSP data file, and the value-weighted index return excluding dividend $R_t = [(P_t / P_{t-1}) - 1]$. Therefore, the dividend series is defined as $D_t = (RD_t - R_t)*P_{t-1}$. The three-month Treasury bill rates are culled from the DRI database. The quarterly excess stock returns are computed in the same manner as in the annual dataset. In the NYSE dataset, ρ and k are computed to be 0.990 and 0.054, respectively.

I begin by testing for unit roots in the four variables over the quarterly period. The results from the ADF and WS tests (reported in Table 10) suggest, similar to the annual series, that all four variables are ~I(1). Also similar to the case of annual observations, results from the Johansen test (reported in Table 11) continue to indicate the presence of one non-zero cointegrating vector in the quarterly sample. Accordingly, I proceed to test the theoretical restrictions H_1 and H_2. Results in the first column of Table 12 suggests that hypothesis H_1 is not rejected for any of the three models. This implies that the dividend-price ratio is stationary in the quarterly NYSE dataset. As the second column of Table 12 reports, both restrictions H_1 and H_2 cannot be jointly rejected for Model 3. This means that the decomposed elements from the GG method empirically represent the fundamental and non-fundamental components of NYSE prices.

The last two columns of Table 12 display the test statistics for the significance of the non-fundamental component of stock prices. Consistent with the results from the annual data, the null hypothesis of a zero non-fundamental component of stock prices is soundly rejected for all three models, including the preferred Model 3. This finding provides further support to my inference that stock prices significantly deviate from the fundamental value in the post-WW II period. Although both time-varying discount factors (excess returns and inflation premium) are incorporated in the fundamental component of the stock prices in my augmented present-value model, I still observe a significant non-fundamental component. All this seems to suggest that variations in stock prices (however measured) are adequately accounted for by the inflation-augmented present-value model in the pre-WW II period, but not in the post-WW II period.

Table 10 Unit Root Test Results
(NYSE Dataset Quarterly: 1947:1 - 1997:4)

Levels				
	p_t	d_t	i_t	er_t
Weighted Symmetric Test	-1.30 [18]	-2.71 [18]	-2.25 [10]	-1.53 [18]
Augmented Dickey-Fuller Test	-1.28 [18]	-2.11 [18]	-2.07 [10]	-1.08 [18]
First-Differences				
	Δp_t	Δd_t	Δi_t	Δer_t
Weighted Symmetric Test	–8.46** [2]	-4.45 ** [8]	-5.42 ** [8]	-7.80** [8
Augmented Dickey-Fuller Test	-8.35 ** [2]	-4.32 ** [8]	-5.29 ** [8]	-7.76** [8]

Notes:
Variables are defined as follows: p_t is the log stock price, d_t is the log dividend, i_t is the nominal interest rate, and er_t is excess stock returns relative to short debt. A time trend is included in the unit root test regression. The numbers in brackets are proper lags generally selected by AIC. The ** indicates rejection at the 5% level.

Table 11 Johansen Cointegration Tests
(NYSE Dataset Quarterly: 1947:1 - 1997:4)

	Trace Statistics: $LR_{trace} = -T \sum_{i=r_0+1}^{p} \ln(1-\hat{\lambda}_i)$				**Cointegrating Vector (β):**
	H_0: r=0	**H_0: r≤1**	**H_0: r≤2**	**H_0: r≤3**	**(normalized β')**
Model 1 $Z_t = [p_t, d_t, k]'$	66.27**	1.89			[1, -1.09, -5.62]
Model 2 $Z_t = [p_t, d_t, er_t, k]'$	73.99**	9.963	2.60		[1, -1.18, -3.09, -5.59]
Model 3 $Z_t = [p_t, d_t, er_t, i_t, k]$	84.77**	22.94	10.02	1.99	[1, -1.73, 315.50, -325.34, -5.41]

Notes: See notes to Table 2. The VAR lags. The proper (AIC-selected/white-noise) lags in the VARs are 4 for all three models.

Table 12 Tests for Theoretical Restrictions (H_1 and H_2) and
Significance of the Non-Fundamental Component of Stock Prices
(NYSE Dataset Quarterly: 1947:1 - 1997:4)

	Test for H_1	**Joint Test of H_1 and H_2**	**Test for significance of non-fundamental component without restriction**	**Test for significance of non-fundamental component with H_1 and/or H_2 restrictions**
Model 1	$\chi^2(1)$=0.44	--	$\chi^2(1)$=15.50**	$\chi^2(2)$=15.51**
Model 2	$\chi^2(1)$=1.42	--	$\chi^2(1)$=16.12**	$\chi^2(2)$=16.12**
Model 3	$\chi^2(1)$=1.18	$\chi^2(2)$ = 1.35	$\chi^2(1)$=16.81**	$\chi^2(3)$=17.42**

Notes: The ** indicates rejection at the 5% level.

CHAPTER 6

CONCLUSION AND DISCUSSIONS

In this study, I derive some testable implications of fundamental and non-fundamental components of stock prices. In order to control for the role of time-varying expected inflation and be able to perform reasonable empirical tests, I use a nominal (rather than a real) interpretation of the present-value model. I demonstrate theoretically that the fundamental and non-fundamental components represent the permanent and temporary components (in the Gonzalo and Granger (1995) sense) of stock prices. The empirically identified cointegrating vector confirms my conjecture for the model with time-varying expected inflation. Various fundamental and non-fundamental exclusion tests performed with annual data over 1871-1997 indicate that expected inflation is an important fundamental. The inflation-augmented present-value model seems to adequately account for stock price variations with an insignificant non-fundamental component. Thus, results deduced from a 120-year sample suggest that the present-value model is an adequate representation of the data, provided that the model is augmented by time-varying expected inflation. This suggests that the observed over-reaction of stock prices and the mean-reversion behavior in stock returns may very well be due to the important, but neglected, inflation premium.[14]

However, my evidence in support of the inflation-augmented present-value model seems somewhat fragile and highly sensitive to certain model details, most notably lag structures. Indeed, the Hansen-Johansen recursive analysis reveals that the parameters in the non-fundamental component lack stability starting with the World War II period. This finding suggests the need to perform my tests separately on the pre- and post- WW II periods. Results from the two sub-samples confirmed my suspicion of a significant regime shift in the post-WW II period. In particular, the results for the pre-WW II period continued to support the inflation-augmented present-value model. However, this model failed to adequately account for price variations in the post-WW II period and the non-fundamental component did not show a tendency towards zero. This inference against market rationality, as represented by the augmented present-value mode, stands up to alternative specifications (such as measurement of variables and data frequency). Such a finding is clearly in line with Shiller's belief in market irrationality and also consistent with Campbell's (1991) conclusion that market predictability is "overwhelming" only in the recent period since 1950s.

[14] Of course, a question that naturally arises is whether expected inflation is also a fundamental variable of real (as opposed to nominal) stock prices. Since the real version of the Campbell and Shiller model does not allow for the role of expected inflation, it is difficult to address this question. However, extensive research has shown that there is a negative relation between changes in real stock prices and changes in inflation. The culprit behind this negative relation may involve many factors including real economic activity [Fama (1981)], alternative monetary regimes [Kaul (1990)] and fiscal deficits [Geske and Roll (1983) and Darrat (1990)]. This dissertation argues that a time-varying inflation premium has important influence on stock prices. Overlooking this variable is capable of producing "unexplained" excess volatility in stock prices.

Future research should further modify the PVM by incorporating other potential discount factors (e.g. unexpected inflation) and/or using alternatively measured cash flows (instead of directly using dividends). Clearly, behavioral models for financial asset pricing will be a very promising research area in finance.

APPENDIX I

SUMMARY OF MAJOR STUDIES

Exhibit 2-1 Major Studies on Excess Volatility of Stock Prices

Shiller (1981)	Argues that the stock prices are too volatile to be justified by the rational expectation present-value model. Thus, the stock prices have excess volatility and considered to be irrational.
LeRoy and Porter (1981)	Develop a version of variance bounds test similar to Shiller's (1981) and find that the volatility of actual stock prices is higher than the volatility of rational prices implied by the present-value model by a large degree.
Flavin (1983)	Criticizes Shiller's econometric tests on two aspects: (i) small sample bias; (ii) biased computation procedure of the rational-expectation prices.
Kleidon (1986)	Argues that Shiller's (1981) "excess-volatility" finding is completely unwarranted, depending on the model specification of dividend forecasts. In addition, the trend-stationarity assumption in Shiller's (1981) may also induce econometric problems.
Marsh and Merton (1986) and Merton (1987)	Observe that dividend smoothing by management could bias variance-bounds test toward rejection. If dividends are slow to reflect changes in underlying profitability, the measured dividend volatility could give an impression that fundamentals had remained stable even when the opposite is the case.
Campbell and Shiller (1987)	Test the present-value model in the bivariate vector autoregression of stock prices and dividends and reject the model.

Campbell and Shiller (1988a)	Test the equality of price-dividend ratio and the present value of future dividend growth rates implied by the log-linear present value model, and they find robust evidence of significant violation.
Campbell and Shiller (1988b)	Add corporate earnings to the price-dividend vector autoregression and find that earning is a strong predictor of dividend growth (return on stock). This finding is against the present-value model, according to which price is a sufficient summary statistic for future dividend growth.
West (1988a)	Derives a variance-bounds test that is valid even if dividends are non-stationary and does not require a proxy for the rational prices.
Shea (1989)	Points out two major problems with Mankiw, Romer and Shapiro's (MRS) test: (i)the outcome of MRS' test is sensitive to the choice of terminal date; (ii) neglect of stationarity property of the series.
Mankiw, Romer and Shapiro (1991)	Provide an unbiased volatility test which shows no evidence of excess volatility of stock prices.
Gilles and LeRoy (1991)	Derive a variance-bound test that is valid if dividends follow geometric random walk and stock prices are non-stationary (but cointegrated). Their test rejects the present value model quite decisively.
LeRoy and Parke (1992)	Adapt West's (1988a) variance-bounds test from a linear case to a log-linear case, which may be appropriate for dividends process.
Shiller and Beltratti (1992)	Analyze the relation between real stock prices and long-term interest rates within the dynamic Gordon model (this model is also called the rational expectations present value model) derived in Campbell and Shiller (1988a) and find that real stock prices drop when long-term interest rates rise (and rise when they fall) more than would be implied by the simple present value model.

Exhibit 2-2 Major Studies on Stock Prices Predictability

Fama and French (1988a)	Estimate an autoregression model of N-horizon return and find that long-horizon (3- to 5-year) returns are negatively autocorrelated and hence can be predictable.
Fama and French (1988b)	Examine the relationship between (nominal and real) stock returns and the dividend yields and find the dividend yields to be a significant predictor for stock returns.
Lo and MacKinlay (1988)	Document a positive autocorrelation for weekly returns using their variance ratio tests
Porterba and Summers (1988)	Find that the variance of returns increases at a rate which is less than proportional to holding period N, implying that returns are mean reverting (for 8>N>3 years).
Campbell (1991)	Finds that the variance of unexpected stock returns is explained primarily by the variance of expected stock returns and variance of cash flows (dividends). Both asymptotic standard errors and the results of a small Monte Carlo experiment show that there is only weak evidence for stock return predictability in the prewar period. The evidence that returns are predictable is overwhelming only in the period after 1952.
Reichenstein and Rich (1994)	Discuss how investors would exploit the return predictability in their own investment portfolios.
Lander, Orphanides and Douvogiannis (1997)	Find that converting a simple mean-reverting theory into a trading rule can yield significantly higher returns (in a statistical sense) than would be expected by pure chance alone.

Exhibit 2-3 Major Studies on Methodological Issues in Stock Price Volatility Tests

Campbell and Shiller (1989)	Study the small sample bias of the dividend-price ratio model of Campbell and Shiller (1988a) and argue that the possible bias does not justify the rejection of the present value model.
McQueen (1992)	Uses generalized least squares (GLS) test for 1926-1987 period and fails to reject the random walk (unpredictability) hypothesis of stock returns.
Pesaran and Timmermann (1994)	Provide a study of stock returns that attempts to address the above criticisms of misspecification in the predictability tests and yet reinforce the earlier results that excess returns are predictable and can be explained quite well by a relatively small number of explanatory variables.
Lee (1995)	Investigates the response of stock prices to permanent and temporary shocks to dividends and finds that stock prices respond significantly to both the permanent and the temporary shocks to dividends. His findings add evidence to the mounting literature in support of the observed mean-reverting behavior of stock returns by incorporating a temporary component into stock prices.
Lee (1996a)	Documents that dividends respond strongly only to permanent changes in earnings but not to transitory changes in earnings. This finding supports the permanent earnings hypothesis and the Marsh and Merton (1987) dividend smoothing hypothesis which is introduced to explain the seemingly excess volatility of stock prices found in Shiller (1981).
Lee (1996b)	Finds that stock prices, dividends, and earnings are cointegrated with a single cointegrating vector, suggesting that there is an equilibrium force that tends to keep these series together over time. He also finds that a substantial fraction of stock price movement is driven by neither earnings changes nor dividend changes, implying that stock prices deviate from the fundamental value.

Lamoureux and Zhou (1996)	Argue that whether returns consist of a material stationary (temporary) component is questionable and perhaps due to inadequate data. Adopting a subjectivist analysis (treating the data as fixed), they employ a Bayesian approach and find that stock prices follow a random walk (and hence unpredictable).
Lee (1998)	Identifies various components of stock prices and examines the response of stock prices to different types of shocks: permanent and temporary changes in earnings and dividends, changes in discount factors, and non-fundamental factors. Results suggest that the over-reaction of the stock market and the mean reversion in stock returns are primarily in response to excess return changes, and partly in response to non-fundamental factors.

APPENDIX II

DATA

Cowles/S&P 500 DATASET
(Annual Data: 1871-1997)

YEAR	PRICES	DIVIDENDS	4-6 MONTH COMMERCIAL PAPER RATES
1871	4.44	0.26	6.35
1872	4.86	0.30	7.81
1873	5.11	0.33	8.35
1874	4.66	0.33	6.86
1875	4.54	0.30	4.96
1876	4.46	0.30	5.33
1877	3.55	0.19	5.03
1878	3.25	0.18	4.90
1879	3.58	0.20	4.25
1880	5.11	0.26	5.10
1881	6.19	0.32	4.79
1882	5.92	0.32	5.26
1883	5.81	0.33	5.35
1884	5.18	0.31	5.65
1885	4.24	0.24	4.22
1886	5.20	0.22	4.26
1887	5.58	0.25	6.11
1888	5.31	0.23	5.02
1889	5.24	0.22	4.68
1890	5.38	0.22	5.41
1891	4.84	0.22	5.97
1892	5.51	0.24	3.93
1893	5.61	0.25	8.52
1894	4.32	0.21	3.32
1895	4.25	0.19	3.09
1896	4.27	0.18	5.76
1897	4.22	0.18	3.44

YEAR	PRICES	DIVIDENDS	4-6 MONTH COMMERCIAL PAPER RATES
1898	4.88	0.20	3.55
1899	6.08	0.21	3.36
1900	6.10	0.3	4.64
1901	7.07	0.32	4.30
1902	8.12	0.33	4.72
1903	8.46	0.35	5.50
1904	6.68	0.31	4.34
1905	8.43	0.33	4.17
1906	9.87	0.40	5.47
1907	9.56	0.44	6.23
1908	6.85	0.40	5.32
1909	9.06	0.44	3.65
1910	10.08	0.47	5.26
1911	9.27	0.47	4.00
1912	9.12	0.48	4.35
1913	9.30	0.48	5.65
1914	8.37	0.42	4.64
1915	7.48	0.43	3.65
1916	9.33	0.56	3.64
1917	9.57	0.69	4.25
1918	7.21	0.57	5.98
1919	7.85	0.53	5.56
1920	8.83	0.51	7.30
1921	7.11	0.46	7.44
1922	7.30	0.51	4.58
1923	8.90	0.53	4.96
1924	8.83	0.55	4.34
1925	10.58	0.60	3.87
1926	12.65	0.69	4.28
1927	13.40	0.77	4.26
1928	17.53	0.85	4.64
1929	24.86	0.97	6.01
1930	21.71	0.98	4.15
1931	15.98	0.82	2.43
1932	8.30	0.50	3.36
1933	7.09	0.44	1.46
1934	10.54	0.45	1.01
1935	9.26	0.47	0.75
1936	13.76	0.72	0.75
1937	17.59	0.80	0.88
1938	11.31	0.51	0.88

YEAR	PRICES	DIVIDENDS	4-6 MONTH COMMERCIAL PAPER RATES
1939	12.50	0.62	0.56
1940	12.30	0.67	0.56
1941	10.55	0.71	0.53
1942	8.93	0.59	0.63
1943	10.09	0.61	0.69
1944	11.85	0.64	0.72
1945	13.49	0.66	0.75
1946	18.02	0.71	0.76
1947	15.21	0.84	1.01
1948	14.83	0.93	1.35
1949	15.36	1.14	1.58
1950	16.88	1.47	1.32
1951	21.21	1.41	2.12
1952	24.19	1.41	2.39
1953	26.18	1.45	2.58
1954	25.46	1.54	1.80
1955	35.60	1.64	1.81
1956	44.15	1.74	3.21
1957	45.43	1.79	3.86
1958	41.12	1.75	2.54
1959	55.62	1.83	3.74
1960	58.03	1.95	4.28
1961	59.72	2.02	2.91
1962	69.07	2.13	3.39
1963	65.06	2.28	3.50
1964	76.45	2.50	4.09
1965	86.12	2.72	4.46
1966	93.32	2.87	5.44
1967	84.45	2.92	5.55
1968	95.04	3.07	6.17
1969	102.04	3.16	8.05
1970	90.31	3.14	9.11
1971	93.49	3.07	5.66
1972	103.3	3.15	4.62
1973	118.42	3.38	7.93
1974	96.11	3.60	11.03
1975	72.56	3.68	7.24
1976	96.86	4.05	5.70
1977	103.81	4.67	5.28
1978	90.25	5.07	7.78
1979	99.71	5.65	10.88

YEAR	PRICES	DIVIDENDS	4-6 MONTH COMMERCIAL PAPER RATES
1980	110.87	6.16	11.37
1981	132.97	6.63	17.63
1982	117.28	6.87	14.60
1983	144.27	7.09	9.37
1984	166.39	7.53	11.11
1985	171.61	7.90	8.35
1986	208.19	8.28	7.31
1987	264.51	8.81	6.25
1988	250.48	9.73	7.63
1989	285.41	11.05	9.29
1990	339.97	12.10	8.43
1991	325.50	12.20	6.92
1992	416.08	12.38	3.91
1993	435.23	12.58	3.44
1994	472.99	13.18	4.35
1995	465.25	13.79	6.45
1996	614.42	14.90	5.68
1997	766.22	15.33	5.78

NYSE DATASET
(Quarterly Data: 1947:1 1997:4)

Date	PRICES	DIVIDENDS	3-MONTH T-BILLS
Mar-47	14.1690	0.1532	0.3800
Jun-47	13.2650	0.1745	0.3800
Sep-47	13.8917	0.1700	0.7367
Dec-47	13.8500	0.2682	0.9067
Mar-48	13.2640	0.1805	0.9900
Jun-48	14.8027	0.1911	1.0000
Sep-48	14.1020	0.1924	1.0500
Dec-48	13.6633	0.3078	1.1400
Mar-49	13.2177	0.2046	1.1700
Jun-49	12.7667	0.2084	1.1700
Sep-49	13.5533	0.1924	1.0433
Dec-49	14.5610	0.3315	1.0767
Mar-50	15.3453	0.2196	1.1033
Jun-50	16.0837	0.2278	1.1533
Sep-50	16.4780	0.2765	1.2200
Dec-50	17.5447	0.4345	1.3367
Mar-51	19.1573	0.2544	1.3667

Date	PRICES	DIVIDENDS	3-MONTH T-BILLS
Jun-51	19.2037	0.2593	1.4900
Sep-51	20.4480	0.2660	1.6033
Dec-51	20.4050	0.3450	1.6100
Mar-52	20.9307	0.2689	1.5667
Jun-52	20.7763	0.2724	1.6467
Sep-52	21.2890	0.2692	1.7833
Dec-52	21.5880	0.3365	1.8933
Mar-53	21.9937	0.2708	1.9800
Jun-53	20.9000	0.2703	2.1533
Sep-53	20.3667	0.2710	1.9567
Dec-53	21.1183	0.3538	1.4733
Mar-54	22.6550	0.2776	1.0600
Jun-54	24.6730	0.2843	0.7867
Sep-54	26.2363	0.2872	0.8833
Dec-54	28.5697	0.3940	1.0167
Mar-55	30.9950	0.2984	1.2267
Jun-55	32.9517	0.3082	1.4833
Sep-55	34.9460	0.3251	1.8567
Dec-55	35.4540	0.4616	2.3400
Mar-56	37.0167	0.3377	2.3267
Jun-56	38.1623	0.3544	2.5667
Sep-56	38.6080	0.3560	2.5833
Dec-56	37.3393	0.4473	3.0333
Mar-57	36.4293	0.3704	3.1000
Jun-57	38.8843	0.3756	3.1367
Sep-57	37.1700	0.3790	3.3533
Dec-57	33.3453	0.4294	3.3033
Mar-58	34.0533	0.3714	1.7600
Jun-58	36.2617	0.3627	0.9567
Sep-58	39.7030	0.3516	1.6800
Dec-58	43.5827	0.3934	2.6900
Mar-59	46.0213	0.3620	2.7733
Jun-59	48.4417	0.3740	3.0000
Sep-59	48.9647	0.3784	3.5400
Dec-59	48.6450	0.4455	4.2300
Mar-60	46.3177	0.3930	3.8733
Jun-60	46.3923	0.3931	2.9933
Sep-60	46.2377	0.3842	2.3600
Dec-60	46.3987	0.4269	2.3067
Mar-61	53.1377	0.3964	2.3500
Jun-61	55.2187	0.3926	2.3033
Sep-61	56.5393	0.4219	2.3033
Dec-61	59.1617	0.4678	2.4600

Date	PRICES	DIVIDENDS	3-MONTH T-BILLS
Mar-62	58.0853	0.4139	2.7233
Jun-62	49.7567	0.4139	2.7133
Sep-62	47.9863	0.4075	2.8400
Dec-62	50.1347	0.5076	2.8133
Mar-63	54.3363	0.4156	2.9067
Jun-63	57.7733	0.4419	2.9367
Sep-63	58.7270	0.4269	3.2933
Dec-63	60.7243	0.5493	3.4967
Mar-64	63.7327	0.4584	3.5300
Jun-64	65.7430	0.4878	3.4767
Sep-64	67.8197	0.4723	3.4967
Dec-64	69.2263	0.5813	3.6833
Mar-65	71.6343	0.5004	3.8900
Jun-65	71.7563	0.5402	3.8733
Sep-65	72.0440	0.5066	3.8633
Dec-65	76.2080	0.6524	4.1567
Mar-66	75.9970	0.5542	4.6033
Jun-66	72.9813	0.5773	4.5800
Sep-66	66.1063	0.5585	5.0300
Dec-66	67.2333	0.6386	5.2000
Mar-67	74.3647	0.5772	4.5133
Jun-67	77.7433	0.5963	3.6567
Sep-67	81.5663	0.5808	4.2933
Dec-67	81.4733	0.6476	4.7433
Mar-68	77.9930	0.5999	5.0400
Jun-68	85.4827	0.6230	5.5133
Sep-68	86.4420	0.6044	5.1967
Dec-68	91.8320	0.6794	5.5800
Mar-69	87.7960	0.6251	6.0867
Jun-69	87.5847	0.7053	6.1900
Sep-69	80.0463	0.6310	7.0100
Dec-69	81.6027	0.6769	7.3467
Mar-70	75.9053	0.6398	7.2100
Jun-70	64.9893	0.6343	6.6667
Sep-70	68.4567	0.6262	6.3267
Dec-70	73.6690	0.6602	5.3500
Mar-71	83.1697	0.6239	3.8367
Jun-71	86.2323	0.6285	4.2400
Sep-71	83.4420	0.6206	5.0033
Dec-71	82.7183	0.6422	4.2300
Mar-72	91.1310	0.6263	3.4367
Jun-72	92.9473	0.6424	3.7700
Sep-72	93.4060	0.6183	4.2200

Date	PRICES	DIVIDENDS	3-MONTH T-BILLS
Dec-72	97.9390	0.6867	4.8633
Mar-73	94.2593	0.6398	5.7000
Jun-73	86.3497	0.6619	6.6033
Sep-73	88.7767	0.6599	8.3233
Dec-73	83.4410	0.7842	7.5000
Mar-74	79.5067	0.6987	7.6167
Jun-74	72.0110	0.7299	8.1533
Sep-74	58.5567	0.7462	8.1900
Dec-74	58.2907	0.8127	7.3600
Mar-75	66.6227	0.7528	5.7500
Jun-75	75.6053	0.7661	5.3933
Sep-75	71.8283	0.7610	6.3300
Dec-75	74.3033	0.7866	5.6267
Mar-76	83.9770	0.7774	4.9167
Jun-76	84.7033	0.8330	5.1567
Sep-76	86.3547	0.8429	5.1500
Dec-76	86.9833	0.9526	4.6733
Mar-77	84.6670	0.9169	4.6300
Jun-77	83.6693	0.9719	4.8400
Sep-77	83.0427	0.9418	5.4967
Dec-77	80.7400	1.0871	6.1100
Mar-78	76.6040	1.0130	6.3933
Jun-78	84.0037	1.0744	6.4767
Sep-78	89.4490	1.0533	7.3133
Dec-78	81.9590	1.1720	8.5700
Mar-79	86.5513	1.1376	9.3833
Jun-79	89.0067	1.2234	9.3767
Sep-79	95.1823	1.1669	9.6733
Dec-79	93.4467	1.3171	11.8433
Mar-80	97.5123	1.3079	13.3533
Jun-80	98.0010	1.3287	9.6167
Sep-80	110.1483	1.3400	9.1533
Dec-80	120.4903	1.5698	13.6133
Mar-81	118.1617	1.4151	14.3900
Jun-81	119.1240	1.4385	14.9067
Sep-81	110.9910	1.4303	15.0533
Dec-81	111.4237	1.5193	11.7500
Mar-82	103.2877	1.5108	12.8133
Jun-82	100.9380	1.5230	12.4200
Sep-82	103.4130	1.5342	9.3167
Dec-82	123.5803	1.5714	7.9067
Mar-83	133.8533	1.5936	8.1067
Jun-83	148.6767	1.5454	8.3967

Date	PRICES	DIVIDENDS	3-MONTH T-BILLS
Sep-83	148.4713	1.5844	9.1400
Dec-83	148.5080	1.6634	8.8000
Mar-84	143.1637	1.6382	9.1700
Jun-84	138.1870	1.7059	9.7967
Sep-84	144.0877	1.6536	10.3200
Dec-84	148.3217	1.7911	8.8033
Mar-85	162.2633	1.7524	8.1833
Jun-85	168.4360	1.7772	7.4600
Sep-85	168.6497	1.8033	7.1067
Dec-85	180.3163	1.8565	7.1667
Mar-86	202.4017	1.8581	6.8967
Jun-86	218.0507	2.2690	6.1400
Sep-86	213.9137	1.8920	5.5233
Dec-86	216.8370	2.2357	5.3533
Mar-87	248.3427	1.9719	5.5367
Jun-87	254.9423	2.1421	5.6567
Sep-87	278.3780	2.1640	6.0433
Dec-87	209.3097	2.2706	5.8633
Mar-88	225.7490	2.4679	5.7233
Jun-88	230.0680	2.1714	6.2100
Sep-88	232.5457	2.4044	7.0100
Dec-88	238.1063	2.9431	7.7267
Mar-89	252.0737	2.4818	8.5400
Jun-89	270.1050	2.6731	8.4100
Sep-89	296.6657	2.7522	7.8433
Dec-89	293.0093	2.5681	7.6533
Mar-90	281.0997	2.5173	7.7600
Jun-90	292.3890	2.8321	7.7467
Sep-90	274.0620	2.6099	7.4767
Dec-90	265.3233	2.7669	6.9900
Mar-91	301.0100	2.5374	6.0233
Jun-91	314.8677	2.7110	5.5600
Sep-91	325.2877	2.6011	5.3767
Dec-91	330.1173	2.6778	4.5400
Mar-92	342.2210	2.5504	3.8933
Jun-92	343.9703	2.6854	3.6800
Sep-92	348.7453	2.6076	3.0833
Dec-92	358.1266	2.7149	3.0700
Mar-93	372.5670	2.6884	2.9600
Jun-93	375.1987	2.6943	2.9667
Sep-93	385.0687	2.6717	3.0033
Dec-93	390.8467	2.7301	3.0600
Mar-94	390.1867	3.3690	3.2433

Date	PRICES	DIVIDENDS	3-MONTH T-BILLS
Jun-94	376.4653	3.8869	3.9867
Sep-94	387.5520	2.7613	4.4767
Dec-94	381.2717	3.1475	5.2800
Mar-95	399.3967	2.8433	5.7367
Jun-95	431.3310	2.9512	5.5967
Sep-95	462.5403	3.2501	5.3667
Dec-95	485.3270	3.1474	5.2600
Mar-96	518.7683	3.3780	4.9300
Jun-96	536.9630	3.1364	5.0200
Sep-96	531.8737	3.2110	5.0967
Dec-96	583.9274	3.3167	4.9767
Mar-97	613.2794	3.1919	5.0600
Jun-97	659.4187	3.2065	5.0467
Sep-97	730.5947	3.4228	5.0467
Dec-97	745.2197	3.4684	5.0900

APPENDIX III

PROOF OF PROPOSITIONS AND LEMMA

Proof of Proposition 1

Under Assumption 1,

$$p^f_{t+1} = p^f_t + \varepsilon_{t+1} \text{ , where } \varepsilon_{t+1} \text{ is white-noise.} \tag{A1}$$

or, $$E_t[p^f_{t+1}] = p^f_t \tag{A2}$$

Equation (3.11) and (A2) yield the expression of the fundamental component (3.13) in Proposition 1.

To derive the non-fundamental component, I obtain, by definition:

$$\begin{aligned} E_t[p^{nf}_{t+1}] &= p_{t+1} - E_t\left[p^f_{t+1}\right] \\ &= p_{t+1} - [\rho p_{t+1} - (1-\rho)d_{t+1} - er_{t+1} - i_{t+1} + k] \end{aligned} \tag{A3}$$

I obtain (3.14) by re-arranging (A3).

Proof of Lemma 1

Assumption 2 can be expressed as follows:

$$p^{nf}_t = ap^{nf}_{t-1} + v_t \text{ , where } v_t \text{ white-noise.} \tag{A4}$$

Then, $$\Delta p^{nf}_t = \frac{a-1}{a} p^{nf}_t + v_t \tag{A5}$$

Given $\Delta p_t = \Delta p_t^f + \Delta p_t^{nf}$ and $\Delta p_t^f = p_t^f - p_{t-1}^f = \varepsilon_t$, I obtain:

$$\Delta p_t = \varepsilon_t + \Delta p_t^{nf} = \frac{a-1}{a} p_t^{nf} + (\varepsilon_t + v_t) \tag{A6}$$

where (ε_t+v_t) are also white-noise. Since Δp_t is (presumably) stationary, it follows that p_t^{nf} is also stationary.

Proof of Proposition 2

In the Gonzalo and Granger (1995)–hereafter GG– framework, Z_t can be decomposed into a permanent and a temporary component:

$$\underset{n\times 1}{Z_t} = \underset{n\times(n-r)}{A_1} \underset{(n-r)\times 1}{\alpha'_\perp} Z_t + \underset{n\times r}{A_2} \underset{r\times n}{\beta'} \underset{n\times 1}{Z_t} \tag{A7}$$

with the factor loadings $A_1 = \beta_\perp(\alpha'_\perp\beta_\perp)^{-1}$, $A_2 = \alpha(\beta'\alpha)^{-1}$. Here, $A_1\alpha'_\perp Z_t$ is the permanent component in the system and $A_2\beta' Z_t$ is the I(0) temporary component which can be interpreted as a deviation from the permanent trend, and α and β are the coefficient matrices in the following error-correction model (ECM):

$$\Delta Z_t = \alpha\beta' Z_{t-1} + \sum_{i=1}^{\infty} \Gamma_i \Delta Z_{t-i} + \xi_t \tag{A8}$$

To test whether a series in Z_t contains a zero temporary component, I formulate the hypothesis:

$$\text{H:}\ \underset{n\times r}{\alpha} = \underset{n\times s}{W} \underset{s\times r}{\Psi} \tag{A9}$$

where W is the restriction matrix and Ψ is the (s×r) matrix of coefficients.

Under hypothesis (A9), the temporary component becomes

$$\begin{aligned} A_2\beta' Z_t &= \alpha(\beta'\alpha)^{-1}\beta' Z_t \\ &= (W\Psi)(\beta' W\Psi)^{-1}\beta' Z_t \end{aligned} \tag{A10}$$

When r = 1, $(\beta' W\Psi)$ is a non-zero scalar. Let Q = $(\beta' W\Psi)$, then

$$\underset{}{A_2\beta' Z_t} = \underset{n\times s}{(W}\ \underset{s\times r}{\Psi)}\ \underset{1\times 1}{Q^{-1}}\ \underset{1\times n}{\beta'}\ \underset{n\times 1}{Z_t} \tag{A11}$$

To test whether a particular row in $A_2\beta'Z_t$ is zero, I specify the W restriction matrix as a $(n\times 1)$ vector composed of a zero element and ones for the rest of the elements. For example, to test whether stock prices P_t in Z_t contains a zero temporary component, I specify W as follows:

$\underset{n\times s}{W} = \begin{bmatrix} 0 \\ 1 \\ 1 \\ 1 \end{bmatrix}$, and Ψ becomes a (1×1) scalar.

It follows that

$$\underset{n\times 1}{A_2\beta' Z_t} = (\underset{1\times 1}{\Psi}\ \underset{1\times 1}{Q^{-1}})\ W\ \underset{n\times 1}{(\beta' Z_t)} \tag{A12}$$

The first element of the W vector (i.e., 0) picks up the first element of the temporary component $A_2\beta' Z_t$. Therefore, if hypothesis (A9) is valid, equation (A7) characterizes the temporary component $A_2\beta' Z_t$. Hypothesis (A9) can be formally tested in the ECM by specifying H: $\alpha_1 = 0$, that is, the error-correction term is not significant in the ECM. Johansen and Juselius (1990) provide a likelihood ratio statistic to test this hypothesis (also called weak exogeneity test with respect to α and β). This weak exogeneity test is briefly describe below.

First, consider the following notations: S_{oo} and S_{kk} are the residual moment matrices from the least-square regressions of ΔZ_t and Z_{t-k} on $\Delta Z_{t-1}, \ldots, \Delta Z_{t-k+1}$, respectively; and S_{ok} is the cross-product moment matrix of the residuals, and,

$$S_{kk.b} = S_{kk} - S_{kb}S_{bb}^{-1}S_{bk}$$

$$S_{ka.b} = S_{ka} - S_{kb}S_{bb}^{-1}S_{ba}$$

$$S_{aa.b} = S_{aa} - S_{ab}S_{bb}^{-1}S_{ba}$$

Let B = W ', such that B'W = 0. $S_{kb} = S_{0k}$B, S_{bk} = B'S_{0k}, S_{bb} = B'S_{00}B, S_{ab} = W'S_{00}B.

According to Johansen and Juselius' (1990) Theorem 6.1 (p. 200), under the hypothesis $\alpha = W\Psi$, the maximum likelihood estimator of Ψ can be solved as the eigenvector associated with

$$|\lambda H' S_{kk.b} H - H' S_{ka.b} S_{aa.b}^{-1} S_{ka.b} H| = 0 \qquad \text{(A13)}$$

for $\tilde{\lambda}_1 > \tilde{\lambda}_2 > \ldots > \tilde{\lambda}_{s+1} = \ldots = \tilde{\lambda}_n$, and $V = (\hat{v}_1 \ldots \hat{v}_n)$ normalized by $\hat{V}' S_{kk.b} \hat{V} = I$.

Now take $\hat{\beta} = (\hat{v}_1 \ldots \hat{v}_r)$ that yields the estimates $\Psi = (W'W)^{-1} S_{ak.b} \hat{\beta}$, and

$\alpha = W\hat{\psi} = W(W'W)^{-1} W'(S_{ok} - S_{oo}B(B'S_{oo}B)^{-1} B'S_{ok}) \hat{\beta}$. The maximized-likelihood function is:

$$L_{\max}^{-2/T}(H) = |S_{oo}| \prod_{i=1}^{r} (1 - \hat{\lambda} i) \tag{A14}$$

The likelihood-ratio statistic of Hypothesis (A9) is:

$$-2\ln(Q;H) = T\sum_{i=1}^{r} \ln\{(1-\tilde{\lambda}_i)/(1-\hat{\lambda}_i)\} \tag{A15}$$

Asymptotically, this statistic is χ^2 distributed with r×(p-m-s) degrees of freedom.

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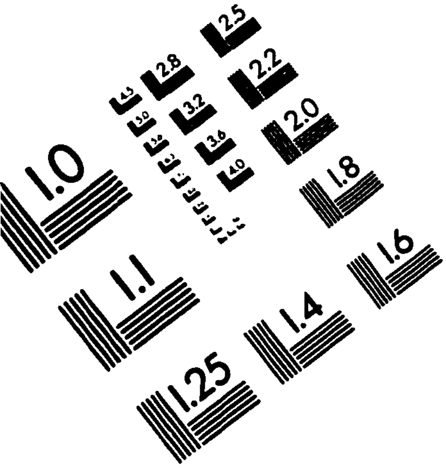
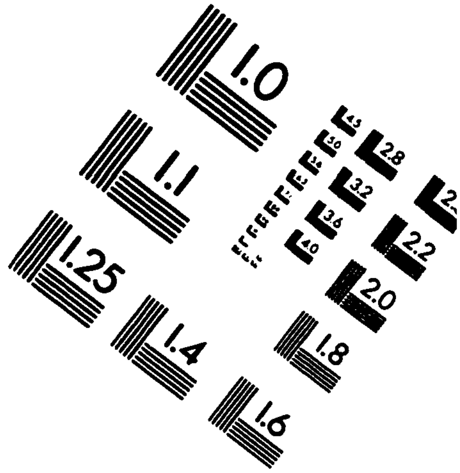
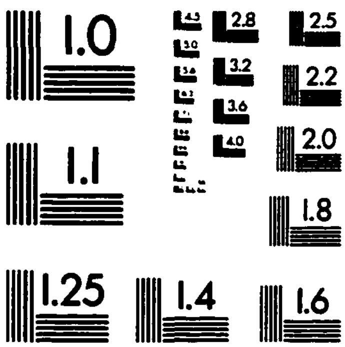
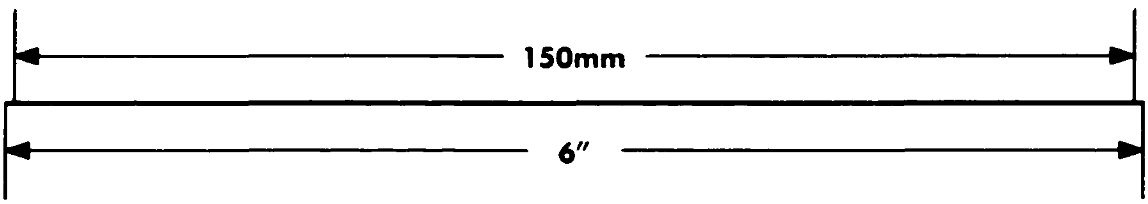
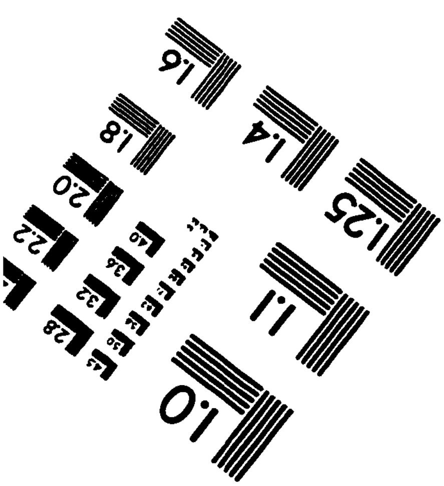
IMAGE EVALUATION TEST TARGET (QA-3)

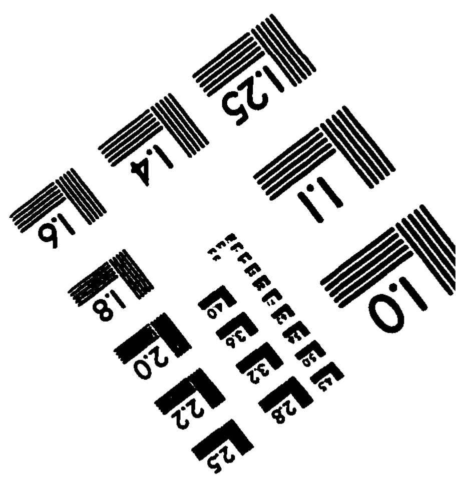












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