Stock Market Valuation :
the Role of the Macroeconomic Risk Premium

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Abstract: Using annual and quarterly data since 1952, we estimate a fundamentals-based empirical model for the earning-price ratio of US stocks. The key fundamental-variable is a time-varying discount rate, decomposed into a time-varying measure for the real interest rate and the equity risk premium. Applying the Johansen procedure, we implicitly estimate the equity risk premium with cointegration test in an error correction model. This equity risk premium is determined by GDP volatility and price inflation. In a lesser extent, the share of U.S. equities held by institutional investors can explain the risk premium. Demographic variables explain the earning-price ratio but only as a short-run phenomenon. Our results suggest that change in the macroeconomic equity risk premium has driven much of the recent run-up in stock prices.

JEL Classification : G19, C32, E32.

Keywords : Johansen Procedure, Valuation Ratios, Equity Risk Premium, Present Value Model.

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1. Introduction

The Standard & Poor's (S&P) 500 stock index closed at an all-time high of 1527 on 24 March, 2000. Since then, the index has declined by about 40% to 917 as of April 30, 2003, roughly where it was about six years ago. Falling stock prices have been accompanied by even larger percentage declines in corporate earnings. In April 2003, the S&P 500 price-earnings (P/E) ratio of 33.23 exceeds the P/E ratio of 27.77 that prevailed at the market peak and is about two times higher than the average P/E ratio of 15.55 going back to 1926. Since April 1980 the P/E ratio – the value that investors assign to each dollar of reported earnings – has been multiplied by about five (from 6.79 to 33.23). This expansion produces an extraordinary annual return on stock prices (excluded dividends) of 9.83% over the period (5.66% since 1926)[1].

According to Ibbotson Associates (2002), the average compound annual return (included dividends) on the S&P 500 was 10.7% from 1926 to 2001. The corresponding return on long-term U.S. government bonds was 5.3%. Stocks delivered an annual excess return over bonds of 5.4% during this period. This observed excess returns on the stock market over the risk-free rate is an order of magnitude higher than the premium predicted by theoretical model (Mehra and Prescott, 1985). This is the so called "equity premium puzzle"[2].

Jagannathan, McGrattan and Scherbina (2001) demonstrate that the U.S. equity premium has declined significantly during the last three decades. The authors calculate that the equity premium is close to zero during the 1990s. Their results suggest that the premium is now about where the standard model says it should be. However, tautologically, there is always some equity risk premium that sets value equal to price. The issue is whether the implied risk premium is in fact expected by investors.

Why would investors be willing to pay more for each dollar of corporate earnings that they have in the past? How can we explain recent stock market valuations? There are several candidate explanations in the literature. These include: (1) a bubble (e.g. Shiller, 2000); (2) an excess demand for stocks by baby-boom cohort (Poterba, 2001); (3) a lower required rate of returns on equities that is a shrinking equity premium (e.g. Siegel, 1999).

The topic of the present paper is to investigate, under the assumption of no bubble, whether fundamentalist factors could explain the valuation ratios over the sample period 1952Q1-2001Q2. According to the standard present value model, stock prices are

[1] From April 1980 to August 2000, this annual stock prices return was 11.19%.
[2] See Kocherlakota (1996) and Siegel and Thaler (1997) for reviews of the literature on the equity premium puzzle before the debate of a possible change in the premium in the last half of the 1990s.

fundamentally determined by the discounted value of its expected future dividends, which in turn derive their value from future expected earnings (e.g., see Campbell, Lo and MacKinlay, 1997). The required rate of return (the discount rate) is composed of the risk free rate and an equity premium. If the discount rate is stationary[3] and the transversality condition holds (no bubble), then the dividend-price and earning-price ratios (D/P and E/P) are stationary (e.g. Campbell and Shiller, 1987). Nevertheless, stationary tests tend to not reject the unit root hypothesis (e.g. Craine, 1993; Lamont, 1998) that implies the presence of a rational bubble (Diba and Grossman, 1988; Hamilton and Whiteman, 1985). However, as shown in Timmermann (1995), when expected returns vary over time, the present value model does not generally imply the stationary of these valuation ratios (D/P and E/P).

Consequently, our empirical investigation is based on the present value model with time-varying expected returns. We implement an error correction model (ECM) that makes it possible to use series with different orders of integration. We estimate, applying the Johansen procedure, the implicit equity risk premium of the model in the cointegration relation, from I(1) variables likely to determine this risk premium required by investors. Real interest rates and earning growth are present in the short-term specification. Several determinants of the equity premium are tested in the long-term specification: demographic variables, macroeconomic risk, share of institutional investors.

The rest of the paper is organized as follows. Section 2 reviews recent literature of the U.S stock market valuation. Section 3 provides the theoretical framework of our analysis. Section 4 reports the empirical results from the unit root tests. Section 5 reviews factors likely to compose the cointegration vector of the ECM. In section 6, we estimate the ECM. Section 7 provides dynamic simulations. Section 8 concludes.

2. Literature survey

This section attempts to provide a survey of the recent run-up stock prices explanations. Different types of explanations are proposed in the literature. First, few authors focus on the "new economy" to explain the recent U.S. stock market valuation by considering intangible

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[3] A time series is called weakly, or covariance, stationary if it satisfies the following criteria (Hamilton, 1994):
$E(y_t) = E(y_{t+m}) = \mu \;\; \forall\, t$ et $\forall\, m$, constant mean;
$\text{var}(y_t) < \infty \;\; \forall\, t$, finite variation;
$\text{cov}(y_t, y_{t+k}) = E[(y_t - \mu)(y_{t+k} - \mu)] = 0 \;\; \forall\, t$ et $\forall\, k$, each covariance only depends on k.
A time series is denoted I(0) when it is stationary already in levels, and non-stationary or integrated of order d (I(d)) when it must be differenced d times in order to achieve stationarity.

capital (Hall, 2000) and the arrival of the information and communication technologies as a technological revolution (e.g. Greenwood and Jovanovic, 1999). Second, the rise in equity prices give evidence of a speculative bubble as advocated by Shiller (2000). Third, some commentators suggest that the rise in U.S. stock prices during the 1990s was partly attributable to the growing demand for financial assets as Baby Boomer cohort began to save for retirement. Finally, few authors suggest the decline in the equity risk premium as a result of lower transaction costs, increased diversification, increased participation (e.g. Heaton and Lucas, 1999 ; Siegel, 1999). In addition, we propose a shrinking equity premium as a result of lower macroeconomic risk.

Intangible Capital

A first possible explanation is that changes in the stock market value of the firm largely reflect changes in the quantity of its capital and not its price. Hall (2000) decomposes the market value of the firm's capital (which can be observed in the stock market) into the quantity of capital and its price. Though physical capital can be observed, much of organizational capital consists of unobservable, such as intellectual property and the quality of employees. Hall's assumptions about the absence of monopoly and the speed at which the capital stock adjusts to its desired level, help him find a close correlation between changes in the value of capital and in its quantity in the data. The recent run-up in the stock market can then be interpreted as an increase in the amount of organizational capital in the economy.

Therefore, Hall (2000) argues that earnings have become increasingly understated in recent years because much of the investment in the new economy is in intangible capital, which is, for conventional accounting procedures, treated as a current expense and deducted from earnings. So, the denominator of the E/P ratio has become biased downward.

However, Bond and Cummins (2000) address whether the increase in the stock market reflects the growing role of intangible capital in generating earnings or a persistent and broadly-based increase in the market valuation of companies relative to their fundamental value. The starting point of their analysis is a measure of firms' value based on analysts' earning forecasts. This is used to construct a measure of average Tobin' Q based on expected earnings, rather than markets' valuation of those earnings. They measure the flow of intangible capital from firm's advertising expenditures and R&D expenses. They consider that this intangible capital is proportional to the stock of physical capital itself endogenous in their model. The authors identify a limited role for intangible investment, and they find no evidence that it accounts for the spectacular rise in the stock market valuation of firms. They conclude that the intangible explanation for equity prices is a “fiction” and that persistent deviations of equity values from firm’s fundamental valuations are an important feature of US stock markets in the past 17 years.

Technology shocks and the stock market

For a few years, literature on stock market models with major technological change has developed (e.g., Greenwood and Yarukoglu, 1997; Greenwood and Jovanovic, 1999; Hobijn and Jovanovic, 2001; Manuelli, 2000). The key idea underlying these models is that a major technical innovation lowers the value of existing firms, causing a reduction in the value of the stock market that will persist until shares in the new firms that can use the new technology make their way to the market. These models are used to explain both the decline in the stock market in the early 1970s and its rise since the mid-1980s with the arrival and the development of the Information Technologies (IT).

These models are based on four assumptions. First, the arrival of the new technology makes the old capital fully obsolete. Second, the success of the IT revolution became evident in the early 1970s. Third, the IT revolution favored new firms, that incumbents resisted it because they may not have the skills needed to adopt new technology, and that this caused their values to fall. And, fourth, new firms cannot immediately compensate for the resulting decline in the stock market, because they will not be in a position to issue tradable securities. Only after these firms have done IPOs will the value of the new technology be reflected in the market.

A stock market bubble

Few authors argue that a stock market bubble explains the ascent in stock prices in the 1990s. Zeira (1999) provides another link between technological change and stock market fluctuations. The author shows that “informational overshooting” occurs when the market expands to a new capacity (increased productivity or entry of new investors), which is unknown until it is reached. If market fundamentals change for an *unknown* period, prices experience a boom, which ends in a crash, due to informational dynamics. This model offers a rational explanation to the phenomenon of stock market booms and crashes. It claims that if fundamentals change for an unknown period, stock prices overshoot.

Shiller (2000) argues that stock prices in the 1990s displayed the classic features of a speculative bubble. Shiller notes that, throughout history, occurrences of major speculative bubbles have generally coincided with the emergence of some superficially plausible "new era" theory. Even with a pickup in trend productivity growth, investors may have overreacted by heedlessly extrapolating the temporary surge in earnings growth of the late 1990s far into the future (See Zeira (1999). High prices are sustained, temporarily, by investor enthusiasm rather than real fundamental factors. Investors, according to Shiller, believe it is safe to purchase stocks, not because of their intrinsic value, but because they can be sold to someone else at a higher price. Stock prices are driven by a self-fulfilling prophecy based on similar beliefs of a large cross section of investors.

Shiller's argument is based on the mean reversion of the valuation ratios. Shiller (2000) and Campbell and Shiller (2001) show that deviations from these historical means have provided valuable forecasting information for future stock prices. When the valuation ratios have been below their historical mean stock prices tended to fall. This implies that the stock market is substantially overvalued, even after its recent correction.

Miller, Weller and Zhang (2002) point out that the risk premium may be reduced by one-sided intervention policy on the part of the Federal Reserve that leads investors into the erroneous belief that they are insured against downside risk. They show that this « insurance » – referred to as the Greenspan put – is consistent with the observation that implied volatility rises as the market falls. The overvaluation is then attributed to an exaggerated faith in the stabilizing power of Mr. Greenspan.

The baby boom and the excess demand for stock

Another explanation emphasizes the demographic effects of the baby boomers. In the 1990s, the baby-boomer cohort reached the wealth-building years of its life cycle, increasing the demand for stocks relative to past cohorts. Then, the aging of the "Baby Boom" cohort would be the key factor for the recent rise in stock prices[4]. The corollary is that stock prices will decline when this group reaches retirement age and begins to reduce its asset holdings.

Such theoretical studies present simulation or analytic results suggesting that demographic change can affect equilibrium returns (e.g. Yoo, 1994; Brooks, 2000; Abel, 1999 and 2001) while determining at the same time the amount of saving according to the life cycle theory

[4] A similar argument mentioned by Campbell and Shiller (2001) is that baby boomers may be more risk-tolerant than earlier generations because they do not remember the 1930s.

(Brumberg and Modigliani, 1954; Ando and Modigliani, 1963) and its allocation. Empirically, however, it has proved difficult to find any conclusive evidence of a systematic relationship between asset returns and age structure (Poterba, 2001)

A shrinking equity premium

Heaton and Lucas ask whether the recent stock market boom can be explained by changes in economic fundamentals in an overlapping generation model. Four candidates are considered: Changes in corporate earning growth, changes in consumer preferences, changes in stock market participation and changes in the amount of diversification among participants. They find that while no single variable explains the large change in stock prices, assuming simultaneous changes in all four variables does. They conclude that the changes in participation that have occurred over this decade[5] are unlikely to be a major part of the explanation. This conclusion is based both on the data, which suggest wealth is now[6], and always has been, controlled by wealthy people, and the model, which implies that participation changes have to be quite extreme to substantially affect expected returns. Increased portfolio diversification, however, is likely to have had a larger effect. The authors suggest that one fundamental reason for the recent stock price run-up may be the rapid growth of mutual funds and the accompanying large increase in diversification.

Finally, we would suggest an alternative explanation for the recent ascent in stock prices: the decline in the macroeconomic risk. This macroeconomic risk would represent the risk associated to the business cycle and would influence the systematic risk[7]. It would be composed of both activity volatility (measured by the GDP volatility) and price inflation. Then, the decline in inflation and GDP volatility seen since the mid 1980s could explain the shrinking equity premium and the run-up in stock prices since almost two decades.

[5] The percentage of U.S. households owning stock in various forms rose by 29.89 percent, from 19.00% in 1983 to 48.89% percent (*Survey of Consumer Finances*; Avery and al., 1984). More than thirty million individuals became stockholders from 1983 to 1998. As noted in several theoretical studies (e.g. Mankiw and Zeldes, 1991; Basak and Cuoco, 1998), an increase in the stock market participation decreases, in theory, the required equity risk premium because it spreads market risk over a broader population.

[6] The change from 42.40 to 79.28 million participants is a 87% increase, but when the numbers are wealth-weighted, the increase is much smaller. According to the SCF, in 1998, 95% of stocks is held by 20% of households. This suggests that stock holdings remain extremely concentrated.

[7] The risk of holding risky assets is composed of a "systematic" and a "specific" risk. Systematic risk is a non-diversifiable risk. It is a market-wide and pervasively influences virtually all security prices. Specific risk (or idiosyncratic risk) involves unexpected events peculiar to a single security or a limited number of securities.

3. Framework

The basic framework of our analysis is a present value model ("the discounted cash-flow model"). In its basic form, a stock's price, P_t, is determined by the present value of its expected future dividends, D_{t+i}, and of the expected terminal price for the holding period K, P_{t+K}:

$$(1) \quad P_t = E_t\left[\sum_{i=1}^{K}\left(\frac{1}{1+R_{t+i}}\right)^i D_{t+i}\right] + E_t\left[\left(\frac{1}{1+R_{t+K}}\right)^K P_{t+K}\right]$$

where R is the expected return (or discount rate) composed of the risk free rate and a risk premium. Under the assumptions that dividends grow at a constant rate, the discount rate is time invariant, and $R > g$, (1) simplifies to

$$(2) \quad P_t = \frac{(1+g)D_t}{R-g},$$

this relationship, traditionally called the Gordon growth model (Gordon, 1962) very compactly illustrates the connection between a stock's price, the current level of its dividends, and the discount rate.

Campbell and Shiller (1988a,b) propose a log-linear approximation of the present value framework that enables us to investigate stock prices behavior under model of a time-varying discount rate. Their formulation leads to:

$$(3) \quad p_t = \frac{k}{1-\rho} + E_t\left[\sum_{j=0}^{\infty}\rho^j\left[(1-\rho)d_{t+1+j} - r_{t+1+j}\right]\right],$$

where p_t denote the log of the stock price, d_t the log of the dividends and r_t the log of the time-varying discount rate. ρ and k are linearization parameters defined by $\rho = 1/(1+\exp(d-p)$ and $k = -\log(\rho) - (1-\rho)\log\left(\frac{1}{\rho}-1\right)$.

Rewriting equation (3) in terms of the log dividend-price ratio, and imposing the transversality condition, leads to:

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

Under the assumptions that the dividend growth and the discount rate in logarithms follow a stationary process, the log stock price and the log dividends are cointegrated, and the log dividend-price ratio follows a stationary process. But if we suppose a time varying-discount rate, the present value model does not generally imply the existence of a stationary relationship between dividends and stock price (Timmermann, 1995).

We suppose that the non-stationary of the valuation ratios could reflect the non-stationary of the risk premium required by investors. We estimate an ECM for which we estimate in the cointegration relation the implicit equity risk premium starting from variables likely to determine this premium. The choice of these variables is inspired by the analyses suggested and reviewed in the preceding section.

4. Data and stationary tests

We determine the order of integration of the series starting from the following procedure: we use initially the tests Augmented Dickey-Fuller (Said and Dickey, 1984) and Kwiatkowsky, Phillips, Schmidt and Shin (1992), to detect the presence of an unit root in the series. The ADF test asserts the variable is I(1) in the null hypothesis while the KPSS test formulates the stationary assumption as the null. However, it is not possible to use the standard approach to testing for unit root, given the low power of the test in the presence of structural breaks (Perron 1992, 1997; Zivot and Andrews 1992).

To test for structural change in the valuation ratios, we use procedures suggested by Perron (1997). The Perron test has an advantage over other unit root tests, which allow for structural breaks, by not requiring the end points of the sample to be trimmed. That is, we undertake estimation without assuming any prior knowledge of any potential break dates. The model is estimated over all possible break dates in the data set, and the break date is chosen to maximize the probability of rejection of the unit root hypothesis.

Variables tested are D/P and E/P ratios of S&P 500 index, respectively *dp* and *ep*; the growth rate of the GDP, *gy*; the growth rate of earnings per share of the index S&P 500, *ge*; the growth rate of dividends per share, *gd* and the real 3 months Treasury Bill rate, *rtb*. [8] All

[8] The appendix I describes the data sources and calculations.

variables in this study are expressed in logarithm. We use quarterly and annual data over the sample period 1952Q1-2001Q2. The use of annual data is justified by research of the "structural" determinants of the stock market valuation and by the introduction of demographic variables, which present a very strong inertia.

Table 1
Results of ADF and KPSS tests
for the E/P and D/P ratios

Variables	ADF t-test	ADF Z-test	lags (BIC)	KPSS eta(mu)	Conclusion
annual					
dp	-0.09	-0.25	0	1.44	UR non-rejected
ep	-1.38	-4.84	0	0.86	UR non-rejected
quarterly					
dp	-1.24	-4.34	1	3.13	UR non-rejected
ep	-2.29	-10.97	1	1.8	UR non-rejected

t-test: 1% critical value –3.46 (***); 5% critical value –2.88 (**) ; 10% critical value –2.57 (*).
Z-test: 1% critical value –20.3 (***); 5% critical value –14.0 (**) ; 10% critical value –11.2 (*).
KPSS: 5% critical value 0.463 (**); 2.5% critical value 0.574 (*).
Critical value from Hamilton (1994).

Table 2
Perron (1997) Tests

Variables	t(alpha=1)		Break date		Lags		Conclusion
	intercept	Slope and intercept	intercept	Slope and intercept	intercept	Slope and intercept	
annual							
dp	-2.12	-3.21	1972 :01	1976:01	4	11	UR non-rejected
ep	-3.77	-4.65	1971 :01	1972:01	0	0	UR non-rejected
Quarterly							
dp	-3.44	-3.65	1994:04	1986:03	1	1	UR non-rejected
ep	-4.05	-4.30	1972:03	1973:01	9	9	UR non-rejected

Intercept break model : 1% critical value –5.41; 5% critical value –4.80; 10% critical value –4.58.
Intercept and slope break model : 1% critical value –5.57; 5% critical value –5.08; 10% critical value –4.82.

Table 1 presents results of ADF and KPSS tests for the E/P and D/P ratios. For both these tests, we cannot reject the presence of a unit root. The tests of Perron (1997) that test the stationary of the series with break on the intercept and/or the slope, lead to the same results (table 2).

The interest rate, the growth rate of dividends, the growth rate of earnings and the growth rate of the GDP appear stationary (table 3). As the unit root is not rejected for the E/P and D/P ratio, the tests are also applied to the first difference to see whether two unit roots are present in their level. These tests conclude that E/P and D/P ratios are I(1).

These results suggest, under the assumption of the absence of rational bubble, that the not-observable component of the model – the equity risk premium – is integrated of order one.

Table 3
ADF and KPSS Tests

Variables	ADF t-test	ADF Z-test	Lags (BIC)	KPSS eta(mu)	Conclusion
annual					
Δdp	-6.24***	-43.61***	0	0.29**	I(0)
Δep	-6.19***	-43.63***	0	0.13**	I(0)
gy	-6.67***	-47.15***	0	0.07**	I(0)
gd	-4.26***	-44.23***	1	0.15**	I(0)
ge	-6.39***	-45.07***	0	0.06**	I(0)
rtb	-2.83*	-13.88**	0	0.39** (with lag = 1)	I(0)
quarterly					
Δdp	-9.61***	-126.60***	0	0.25**	I(0)
Δep	-8.15***	-99.91***	0	0.15**	I(0)
gy	-9.76***	-130.03***	0	0.12**	I(0)
gd	-6.05***	-78.68***	1	0.31**	I(0)
ge	-4.78***	-134.89***	4	0.04**	I(0)
rtb	-3.36**	-31.10***	5	0.48*	I(0)

We will use an error correction model for our estimate that makes it possible to use series with different orders of integration. We estimate, applying the Johansen procedure, the implicit equity risk premium of the model in the cointegration relation, from variables likely to determine this risk premium required by investors.

5. The determinants of the equity risk premium

Initially, we undertake the determination of the integration order of the series, likely to compose the cointegration vectors. We will determine, in the next section, the cointegration vectors from the variables with the same integration order as the D/P and E/P ratios.

The vector of cointegration estimated will represent to some extent a "proxy" for the equity risk premium required by investors. Different types of variables are likely to compose the cointegration vector:

- Demographic variables: *age1*, *age2* and *age3* which respectively represent the fraction of the population in the « asset accumulating years », 40-64[9]; the average age of the population over the age of 19; and the ratio population between 40 and 64 /population over the age of 65.

[9] This indicator is often cited as a key variable in discussions of asset demand and demographic structure.

- The share of the U.S. stocks held in indirect form[10], *inv*. Institutional investors gradually replaced the individual savers like "shareholders" of the companies in the United States. These investors manage a long-term and stable saving (Choi, Laibson and Metrick, 2000; Ameriks and Zeldes, 2000) in diversified portfolios. They are likely to have a lower risk aversion than households, which invest directly on the stock market. Siegel (1998) shows that for long-horizon investors, the risk of holding stocks is less than one would expect by just looking at the annual standard deviation of returns. Also, this variable is used as a proxy to test the diversification hypothesis suggested by Heaton and Lucas (1999).
- Price inflation, *cpi*, which seems to exert a negative influence over the real stock prices and the risk premium (e.g., Sharpe, 2001; Blanchard, 1993).
- GDP Volatility, *voly*, according to the methodology retained by Blanchard and Simon (2001), which represents a macroeconomic risk likely to influence the systematic risk.

We apply the same procedure as in the preceding section to determine the integration order of the series. The unit root and stationary tests appear in tables 4 and 5. The variable *age3* is I(2) and variables *age1* and *age2* are stationary. The variables *inv*, *voly*, *cpi* are integrated of order one. These variables are likely to compose the cointegration vectors in the error correction model.

[10] Whereas the households directly held 93% of the U.S. equities in 1945, they held more than 41% in 2000 according to the *Flow of Funds*. Over the same period, the institutional investors detention rose from 3.48% to 45.42% of the U.S. equities. Institutional investors comprise pension funds, mutual funds, banks, endowments, foundations, corporations and insurance companies. These institutions manage or make professionally manage the funds they have.

Table 4
ADF and KPSS Tests

Variables	ADF t-test	ADF Z-test	Lags (BIC)[+]	KPSS eta(mu)	Conclusion
Annual					
age1	-1.92	-13.27*	2	0.34**	I(0)
age2	-1.55	-13.82***	2	0.43**	I(0)
age3	-1.99	-9.50	1	2.22	UR non-rejected
inv	-1.93	-1.79	0	4.38	UR non-rejected
voly	-1.59	-5.70	1	1.50	UR non-rejected
cpi	-1.81	-6.92	2	0.55 (with 1 lag)	UR non-rejected
Δ*age3*	-0.13	-0.27	0	2.55	UR non-rejected
Δ*inv*	-7.24***	-51.08***	0	0.28**	I(0)
Δ*voly*	-4.94***	-33.32***	0	0.09**	I(0)
Δ*cpi*	-6.95***	-103.34***	1	0.06**	I(0)
Δ^2*age3*	-7.08***	-51.08***	0	0.24**	I(0)
Quarterly					
inv	-2.19	--2.17	0	17.77	UR non-rejected
voly	-1.39	-3.64	1	5.78	UR non-rejected
cpi	-2.40	-12.20*	4	0.81	UR non-rejected
Δ*inv*	-14.76***	-207.93***	0	0.40**	I(0)
Δ*voly*	-11.86***	-165.99***	0	0.09**	I(0)
Δ*cpi*	-6.93***	-192.38***	3	0.06***	I(0)

[+] 3 maximum lags for annual data and 12 maximum lags for quarterly data.

Tableau 5
Perron Tests (1997)

Variables	t(alpha=1)		Break date		Lags[+]		Conclusion
	intercept	Slope and intercept	intercept	Slope and intercept	intercept	Slope and intercept	
Annual							
age3	-2.83	-4.65	1992:01	1987:01	3	2	UR non-rejected
Δ*age3*	-3.38	-4.35	1968:01	1968:01	0	0	UR non-rejected

[+] 3 maximum lags for annual data and 12 maximum lags for quarterly data.

6. The Error Correction Model

Owing to the non-stationary of the E/P[11] ratio, we estimate an error correction model framework based on the procedure developed by Johansen (1988, 1991) as well as Johansen and Juselius (1992, 1994). This approach is suited to detect stationary linear combinations (i.e. cointegration relationships) between I(1) variables. These relationships are interpreted as the long-term economic equilibrium relationships.

Variables likely to compose the vector of cointegration are thus *in fine* the E/P ratio, *ep*, inflation, *cpi*, the GDP volatility, *voly*, the share of U.S. equities held in indirect form, *inv*, the

[11] We estimate the ECM with the E/P ratio since D/P ratio can be affected by corporate financial policy.

growth rate of the ratio population of 40-64 /population over the age 65, *dage3*. The Johansen procedure makes it possible to estimate the relations of cointegration in a multivariate model since the number of cointegration relation is unknown.

Annual data

The number of lags was selected in order to accept the assumption that residuals of the model are white noise and to obtain interpretable long-term relations. We introduced, in accordance with the theoretical model (4), the growth rate of earnings and real interest rates as short-term variables before identifying the cointegration vectors. The short-term specification selected is: ge_{t-1}, rtb, rtb_{t-1}, rtb_{t-2}, rtb_{t-3}.[12] These series can be considered stationary according to tests' practiced previously. The estimate relates, for the moment, on annual data (demographic variables available) to the sample 1952-2000.

The number of lags is two in the VAR-model and one in the error correction form. The determination of the cointegration rank is based on the Trace test. This Trace test indicates the existence of only one relation of cointegration (table 6). We did not include linear trends in the model but we have constrained intercepts to appear only in cointegration relations.

Estimates of the cointegration vector appear on table 7 (model 1). The vector is normalized on the E/P ratio (*ep*) and multiplied by –1.

This vector depicts a positive long-run relationship between the E/P ratio, the GDP volatility and inflation; a negative relation between the E/P ratio and the financial intermediation variable; and a surprisingly positive relation between the E/P ratio and the demographic variable.

The error correction term coefficient has the right sign (-0.153) and is statistically significant (t=-2.100). Long-run exclusion and stationary tests appears on table 8. Stationary tests confirm that the variables are I(1). However, the hypothesis that the demographic variable, Δ*age3*, the financial intermediation variable, *inv*, and the intercept are not in the cointegration space cannot be rejected at a 5% significance level.

[12] Others lags do not appear significant in the different estimates.

Table 6
Determination of Cointegration Rank (1952:01-2000:01)
VAR(2), intercepts in the cointegration space
model 1 : [*ep cpi voly inv* Δ*age3*]

Eigen-value	H0 : r =	Trace test		
		Statistic	Critical value	
			5%	1%
0.5565	0	88.45	75.737	83.930
0.4102	**1**	**52.67**	**53.423**	**60.422**
0.3031	2	29.44	34.795	40.837
0.2002	3	13.55	19.993	24.735
0.0812	4	3.72	9.133	12.731

Table 7
Cointegration relation
model 1

ep	*cpi*	*voly*	*inv*	Δ*age3*	intercept
-1	12.711	0.679	-0.009	2.125	-1.173
Error correction term : α = -0.153 (-2.100)					

Table 8
Long-run exclusion and Stationary Tests (1952:01-2000:01)
VAR(2), intercept in the cointegration space, one cointegration relation
model 1

	Critical value (5%)	*ep*	*cpi*	*voly*	*inv*	Δ*age3*	*intercept*
Exclusion	3.84	8.09	6.75	4.81	**0.00**	**0.02**	**0.63**
Stationary	11.07	21.86	28.83	21.30	21.74	27.76	

We begin again the estimate procedure of the cointegration rank by excluding the demographic variable from the cointegration space[13]. The short-term specification is the same as previously. The variables likely to compose the cointegration vector are now: *ep*, *cpi*, *voly*, *inv*. The number of lags is four in the VAR-model and three in the error correction form. The intercept is restricted to the cointegration space. At a 1% significance level, the Trace test indicates the existence of only one relation of cointegration (table 9).

Table 9
Determination of Cointegration Rank (1952:01-2000:01)
VAR(4), intercepts in the cointegration space
model 2 : [*ep cpi voly inv*]

Eigen-value	H0 : r =	Trace test		
		Statistic	Critical value	
			5%	1%
0.6543	0	79.72	53.423	60.422
0.4242	**1**	**35.12**	**34.795**	**40.837**
0.1546	2	11.93	19.993	24.735
0.1096	3	4.88	9.133	12.731

Estimates of the new cointegration vector appear on table 10 (model 2). Coefficients have the expected signs. The error correction term is statistically significant (t=-8.136) and the associated coefficient has the right sign (-0.638). Long-run exclusion tests show, at a 5% significance level, that all the variables are in the cointegration space (table 11). At a 2.5% significance level (critical value =5.02), the variable *inv* is not in the cointegration space.

The LM tests reject the first and fourth order autocorrelation of residuals. Furthermore, residuals are homoscedastics with regard to the ARCH test. R^2 =0.93; LM(4) = 13.57 (p = 0.63); LM(1) = 8.85 (p = 0.92); ARCH(4) = 4.529 (p = 0.34).

[13] We also carried out the estimate of the cointegration vectors with the demographic variable in level, *age3*, (model 1bis) although this variable is integrated of order two. The results of this estimate appear in appendix. The exclusion tests indicate that the variables *age3*, *inv* and the intercept are not in the cointegration space. The error correction term does not appear statistically significant.

Table 10
Cointegration relation
model 2

ep	*cpi*	*voly*	*inv*	*intercept*
-1	16.114	0.309	-0.197	-2.648
Error correction term : α = -0.638 (-8.136)				

Tableau 11
Long-run exclusion and Stationary Tests (1952:01-2000:01)
VAR(4), intercept in the cointegration space, one cointegration relation
model 2

	Critical value (5%)	*ep*	*cpi*	*voly*	*inv*	intercept
Exclusion	3.84	20.23	19.44	4.17	3.92	12.14
Stationary	9.49	37.52	37.98	37.92	37.64	

We also estimate the error correction model by retaining in the cointegration space only the three variables *ep*, *cpi*, *voly*, in order to consider the influence of the two macroeconomic variables (model 3). The detail of the results of this estimate appears in appendix. The Trace test indicates only one relation of cointegration. The error correction term is statistically significant (t=-5.797). The associated coefficient has the right sign and is high (-0.648).

We conduct a last estimate on annual data with the same specification as in model 2 (*ep*, *cpi*, *voly*, *inv*) but by integrating the demographic variable in the short-term specification, *age2*,[14] which is stationary (model 4). The detail of the results appears in appendix. The Trace test indicates a single vector of cointegration. The coefficient of the demographic variable is negative (-8.395) as expected and is strongly significant (t=-10.379). In this specification, the error correction term is statistically significant (t=-10.362) and the coefficient is larger than in the other specifications (-0.917). We will use these different models during dynamic simulations.

[14] We also carried out the estimates with age1 and d^2age3. Coefficients appeared always negative. The variable, *age2*, retained appeared more significant.

Quarterly data

On quarterly data, the number of lags retained is 6 in an autoregressive writing. The short-term specification is: ge_{t-1}, rtb, rtb_{t-1}, rtb_{t-2}. The Trace test indicates the existence of only one relation of cointegration at the 5% significance level (table 12). The intercept is restricted to the cointegration space. Estimates of the cointegration vector appear in table 13. The signs of the coefficients are identical to those estimated on annual data (model 2). The error correction term is statistically significant (t=-3.818) and the associated coefficient has the right sign (-0.058). R^2 is 0.532. The LM tests reject the autocorrelation of residuals. The ARCH test indicates that residuals are homoscedastics. LM(4) = 23.51 (p =0.10) ; LM(1) = 14.32 (p=0.57) ; ARCH(6) = 4.563 (p = 0.601).

Table 12
Determination of Cointegration Rank (1952:01-2001:02)
VAR(6), intercepts in the cointegration space
Quarterly model : [*ep cpi voly inv*]

Eigen-value	H0 : r =	Trace test		
		Statistic	Critical value	
			5%	1%
0.1522	0	57.86	53.423	60.422
0.0807	**1**	**26.49**	**34.795**	**40.837**
0.0481	2	10.49	19.993	24.735
0.0059	3	1.12	9.133	12.731

Table 13
Cointegration relation
Quarterly model

ep	*cpi*	*voly*	*inv*	*intercept*
-1	20.758	0.094	-0.400	-3.920
Error correction term : α = -0.058 (-3.818)				

7. Dynamic simulations

We finally conduct dynamic simulations of the error correction models. These simulations enable us to apprehend the fit E/P ratio, *ep*, in level. Simulations were performed on models 2,3 and 4 using annual data on the entire sample (figures 1, 2, 3). Simulation in quarterly data was performed over the period 1971:01-2001:02 (figure 4). The estimation results produce reliable simulation results. The over-estimate of the E/P ratio at the end of the sample (1996-2000) is compatible with an overvaluation of the U.S. stock market during this period. However, the extent of this over-estimate is quite different depending on the models selected. The model 4 that integrates the demographic variable, *age2*, in the short-term specification, seems closer to the E/P ratio observed, *ep*, at the end of the period.

Figure 1
Model 2 : [*ep cpi voly inv*]

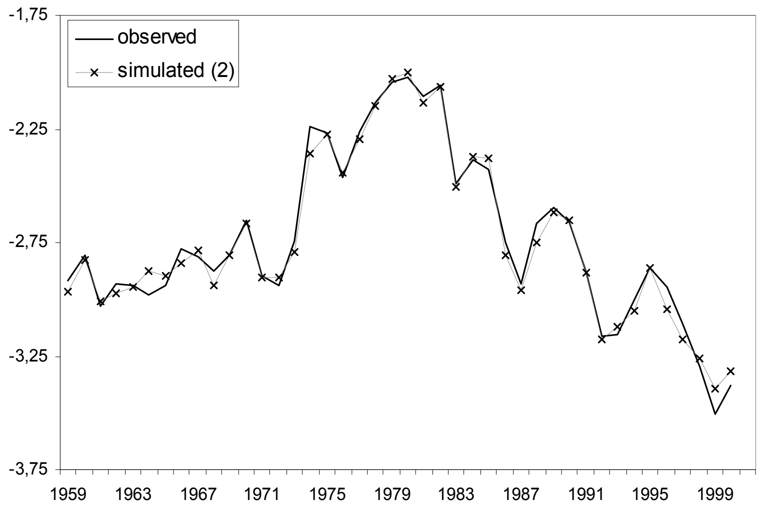


Figure 2
Model 3 : [*ep cpi voly*]

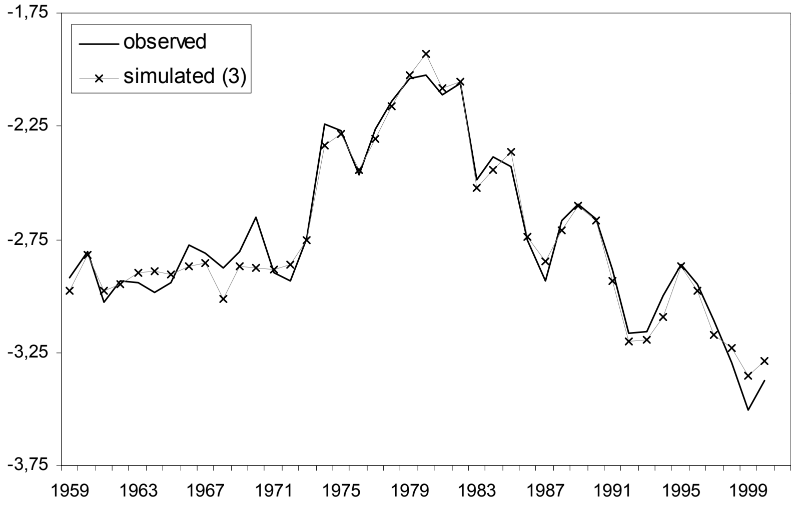


Figure 3
Model 4 : [*ep cpi voly inv*]
with *age2* in the short-term specification

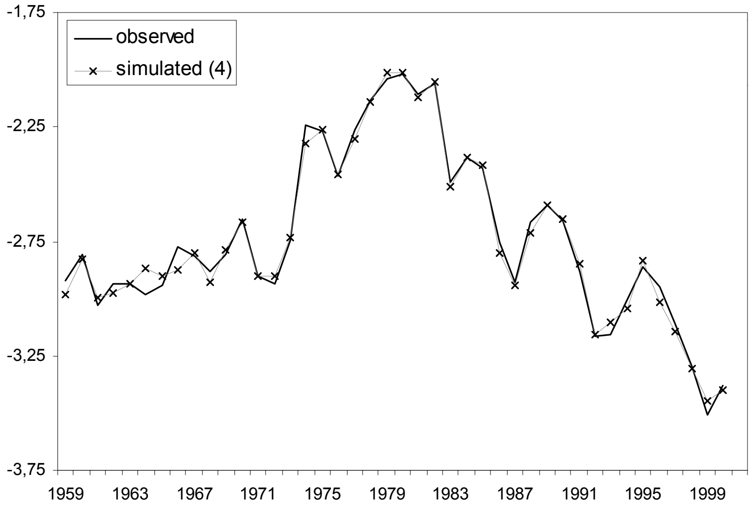
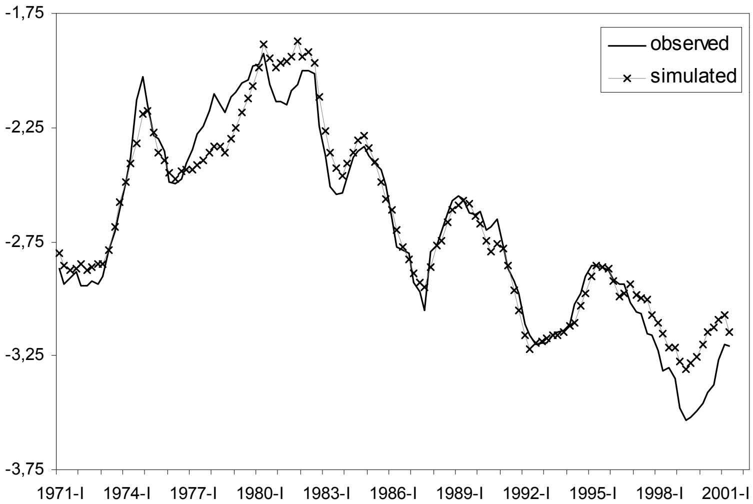


Figure 4
Model 5 : [*ep cpi voly inv*]
Quarterly data



8. Conclusion

Stationary and unit root tests practiced conclude that E/P and D/P ratios are integrated of order one over the sample period 1952-2001. The non-stationary of these ratios implies that they do not have a constant mean. This result contradicts the analysis of Campbell and Shiller (2001) suggesting these valuation ratios revert to their historical means. According to the present value model, the valuation ratios are determined by the real interest rate, the growth rate of earnings or dividends and by the equity risk premium required by investors, which is not observable directly.

However series of real interest rates, earnings and dividend growth appear stationary. Consequently, under the assumption of the absence of rational bubble, the equity risk premium should present an unit root. Thus, we implicitly estimated the equity risk premium to explain the E/P ratio with cointegration test in an error correction model. The decline in the E/P ratio since the 1980s is primarily explained by the decline in inflation and GDP volatility.

Recent studies uncover a sensitive decline in the volatility of the US economy beginning in the 1980s. The volatility of GDP growth since 1984 has been 50 percent lower than it was in the post-war period before 1984. The phenomenon also appears to extend beyond US borders. Blanchard and Simon (2001) show that all G-7 countries except Japan have experienced a decline in volatility in recent periods. Only the UK and Canada show the sharp drop in the mid-1980s, though, with the other countries exhibiting more gradual declines or later falls. Three types of assumptions are advanced to explain this phenomenon for the U .S. economy: (1) Good Luck; (2) Good Policy; (3) Structural Change. Advocates of the "Good Luck" hypothesis argue that the decline in volatility results of a fortuitous decline in the volatility of shocks hitting the economy (e.g. Ahmed, Levin, and Wilson (2000)). This explanation implies that volatility could easily rise again. Advocates of the "Good Policy" hypothesis argue that improved monetary policy to stabilize the economy is the key source of the decline in volatility of the U.S. economy (e.g. Clarida, Gali, and Gertler (2000). Finally, advocates of the "Structural Change" hypothesis argue that structural changes[15] in the private economy are the key to understanding the decline in volatility (e.g. Kim, Nelson and Piger, 2001). These last two hypotheses suggest a permanent decline in volatility[16].

Also, one of the most surprising features of the last business cycle has been the decline in price inflation through the late 1990s. There are several explanations on this low level of inflation. A first attributes the low inflation to a series of "positive supply shocks". These shocks include periodic declines in commodity and energy prices and improvement of the terms of trade with a strong dollar (e.g., Gordon, 1998). A second attributes this decline in inflation to the policy credibility of the Federal Reserve against inflation. It would reduce

[15] Kahn, McConnell and Perez-Quiros (2000) have argued that improvements in information technology and inventory management are the principal source of the decline in output volatility since 1984.

[16] The recent GDP evolution suggests a weak recession since the fall cumulated of the GDP over the first three quarters 2001 appears weaker than that observed on average at the time of the preceding recessions (Note de conjoncture internationale de l'INSEE, Direction de la Prévision, October 2002). After these three quarters of recession, the U.S. economy recorded a strong growth in the last quarter 2001 (2.7% at annual rates) and over the four quarter of 2002 (5.0%; 1.3%; 4.0% ; 1.4%) according to the BEA.

anticipated inflation and would limit the rise of producers prices. The low inflation of decade 1990 would be also explained by a decline in the NAIRU (Brayton, Roberts and Williams, 1999). Katz and Krueger (1999) point to changes in labor markets as the prime reason for the decline in the NAIRU over the past decade. They highlight the demographic trend toward an older, more experienced workforce, increased incarcerations rates, and greater use of temporary help services as factors leading to a secular decline in the unemployment rate since the mid-1980s.

While the negative relation between real stock prices and inflation is largely documented in the empirical literature, the theoretical explanation is controversial. First, this relation reflects an inflation-related risk premium on stocks relative to bonds. However, Modigliani and Cohn (1979) argued that the magnitude of the effect was difficult to rationalize. The more plausible explanation, they argued, is that a form of «money illusion» plagues investors. Investors, on the one hand, would use a nominal rate to discount real dividends and, on the other hand, would underestimate the future dividends while omitting to correct the fall of the actual value of the companies' debt. Another channel, proposed by Fama (1981), suggests that higher inflation induces lower expected real economic activity and/or uncertainty over the conduct of future monetary and fiscal policies, leading investors to demand higher risk premium. We also suggest that low inflation indicates future directions of the monetary policy and induces a lower expected risk free rate.[17]

Our results also show that the evolution of the demographic structure influences stock prices but only as a short-run phenomenon in the error correction model. The influence of the demographic variables appears modest compared to the influence of the identified macroeconomic risk (GDP volatility and inflation). This result relativizes the "meltdown" hypothesis (Poterba, 2001) that is the "Baby Boom" cohort would be a key factor in explaining the recent rise in asset values, and by predictions that asset prices will decline when this group reaches retirement age and begins to reduce its asset holdings.

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[17] The negative relation between real returns and inflation may also be related to Lucas (1973) signal extraction problem associated with inflation. In Lucas model, agents have problems assessing whether a price increase is the result of higher demand in a specific market or the result of an overall price inflation. This may result in less efficient resource allocation that translates into higher required real return as inflation goes up, possibly through a higher equity premium. Feldstein (1980) develops a market equilibrium model of share valuation and shows that an increase in steady-state inflation lowers share prices because of the way depreciation costs and capital gains are treated in tax codes. Danthine and Donaldson (1986) develop a monetary model with rational expectations in which the stock market does provide full insurance against monetary shocks but not against temporary inflation induced by real shocks. In their model, real shocks have an effect on consumption goods prices without changing the expected stream of dividends (because of dividends stickiness) and thus have an impact on stock market real returns.

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Appendix 1 : alternative models

Model 1bis : [*ep cpi voly inv age3*]

Determination of Cointegration Rank (1952:01-2000:01)
VAR(2), intercepts in the cointegration space

Eigenvalue	H0 : r =	Trace test		
		Statistic	Critical value	
			5%	1%
0.5321	0	85.62	75.737	83.930
0.3861	**1**	**52.20**	**53.423**	**60.422**
0.3123	2	30.73	34.795	40.837
0.1753	3	14.26	19.993	24.735
0.1232	4	5.79	9.133	12.731

Cointegration relation

ep	*cpi*	*voly*	*inv*	*age3*	intercept
-1	14.943	1.238	-0.381	-1.917	2.050
Error correction term : α = -0.052 (-0.821)					

Long-run exclusion and Stationary Tests (1952:01-2000:01)
VAR(2), intercept in the cointegration space, one cointegration relation

	Critical value (5%)	*ep*	*cpi*	*voly*	*inv*	*age3*	intercept
Exclusion	3.84	3.91	5.26	10.80	**1.61**	**1.75**	**1.87**
Stationarity	11.07	24.23	28.68	23.72	21.88	22.91	

Model 3: [*ep*, *cpi*, *voly*]

Determination of Cointegration Rank (1952:01-2000:01)
VAR(4), intercepts in the cointegration space

Eigen-value	H0 : r =	Trace test		
		Statistic	Critical value	
			5%	1%
0.4772	0	42.37	34.795	40.837
0.1838	**1**	**15.13**	**19.993**	**24.735**
0.1454	2	6.60	9.133	12.731

Cointegration relation

ep	*cpi*	*voly*	intercept
-1	12.325	0.492	-1.624
Error correction term : α = -0.648 (-5.797)			

Long-run exclusion and Stationary Tests (1952:01-2000:01)
VAR(4), intercept in the cointegration space, one cointegration relation

	Critical value (5%)	*ep*	*cpi*	*voly*	*intercept*
Exclusion	3.84	17.11	18.16	18.09	10.53
Stationarity	7.81	19.32	21.97	20.03	

R^2 = 0.871
LM(4) = 6.59 (p-value =0.68)
LM(1) = 2.43 (p-value = 0.98)
ARCH(4) = 3.867 (p-value = 0.42)

Model 4 : [*ep cpi voly inv*] + *age2* in the short-term specification

Determination of Cointegration Rank (1952:01-2000:01)
VAR(4), intercepts in the cointegration space

Eigen-value	H0 : r =	Trace test		
		Statistic	Critical value	
			5%	1%
0.7437	0	87.92	53.423	60.422
0.4195	**1**	**30.75**	**34.795**	**40.837**
0.1310	2	7.91	19.993	24.735
0.0468	3	2.01	9.133	12.731

Cointegration relation

ep	*cpi*	*voly*	*inv*	intercept
-1	9.781	0.171	-0.104	-9.953
Error correction term : α = -0.917 (-10.362)				

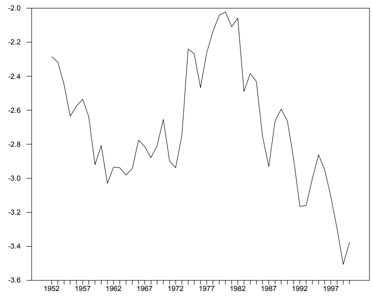
Long-run exclusion and Stationary Tests (1952:01-2000:01)
VAR(4), intercept in the cointegration space, one cointegration relation

	Critical value (5%)	*ep*	*cpi*	*voly*	*inv*	intercept
Exclusion	3.84+	33.11	27.59	4.27	3.59	17.33
Stationary	9.49	52.28	47.10	45.91	52.01	

+ 10% critical value 2.71.

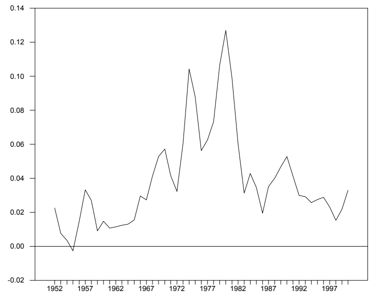
Coefficient *age2* = -8.395 (-10.379)
R^2 = 0.95
LM(4)=15.28 (p-value =0.50)
LM(1) = 10.13 (p-value = 0.86)
ARCH(4) = 8.633 (p-value = 0.07)

Appendix 2 : data



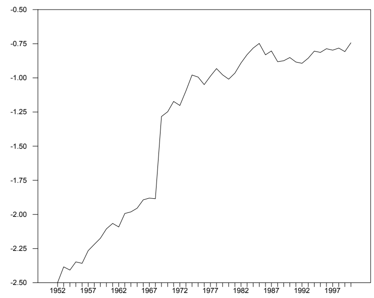
ep

dp



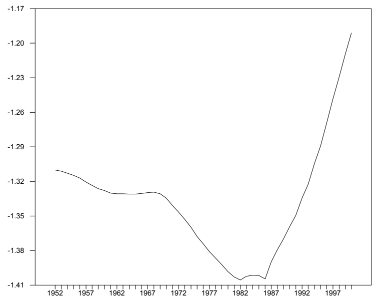
cpi

voly

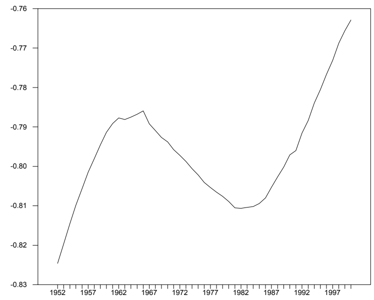


inv

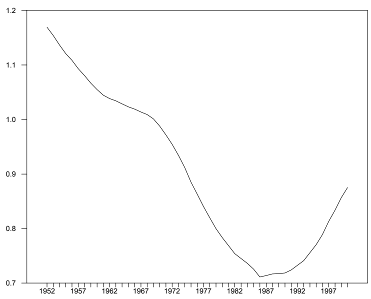
rtb



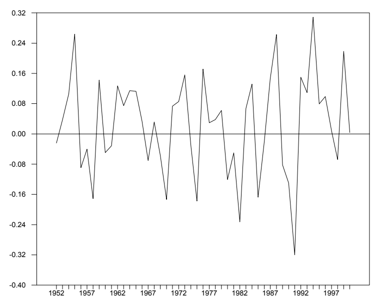
age1



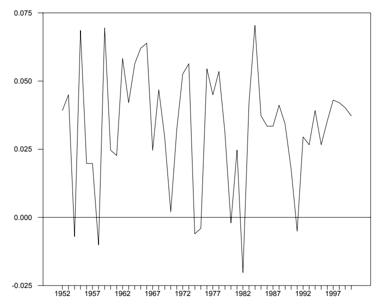
age2



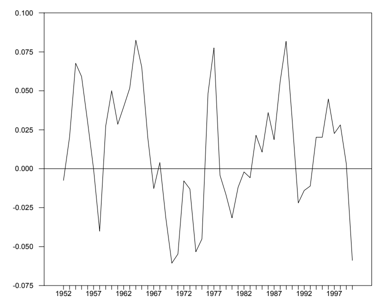
age3



ge



gy



gd

Variables in logarithm	Description	Source
dp ; ep	D/P and E/P ratios	series developed by Professor Robert J. Shiller and available on his web site http://www.econ.yale.edu/~shiller/
gy	GDP growth at annual rates	BEA
ge, gd	Earning per share growth rate and dividend per share growth rate at annual rates (S&P 500)	Robert J. Shiller
rtb	3 months treasury bills rate deflated by the CPI inflation	Board of Governors of the Federal Reserve System
inv	Share of equities held by institutional investors	Flow of Funds Table L.213 (lines 8 to 18)
voly	Rolling standard deviation of quarterly real GDP growth (measured at a annual rate) with a 20 quarters window	BEA Methodology from Blanchard and Simon (2001)
cpi	CPI inflation at annual rates (CPI all urban Consumers)	BLS
age1 ; age2 ; age3	40-64 population/total population; average age of the population over the age of 19; 40-64 population /population over the age of 65	CPS Reports P25-1130 of the U.S. Census Bureau in Poterba (2001).