Equity Returns and Inflation: The Puzzlingly Long Lags

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Abstract. This paper examines data for stock prices and price levels of 14 developed countries during the post-WWII era and compares their behavior in that sample with behavior over the past two centuries in the UK and the US. Contrary to much of the literature of the past several decades, we find that nominal equity prices do, in fact, keep pace with movements in the overall price level. Our results suggest, however, that this is only the case over long periods. The puzzle therefore is not that equities fail the test as inflation hedges, as had been quite widely believed, but that they take so long to pass. G10, E44. Keywords. Stock prices, inflation, Fisher effect, neutrality, cointegration.

In principle, equities ought to be an inflation hedge. In practice, however, evidence of such behavior has been difficult to come by. With few exceptions, studies show that nominal returns on equities do not keep pace with inflation and that real returns and inflation are, in fact, negatively correlated. Indeed the disparity between theory and data was so common place that it became a staple of textbook discussions of financial market behavior.

The purpose of this paper is to reexamine the evidence. To do so we use data for fourteen OECD countries over the post-World War II period and time series for the UK and the US over the longer period 1790 to 2000. What emerges is a rather different picture from the commonly accepted one. During the high inflation period in the 1970s and early 1980s in industrial countries real stock prices did fall. Real returns were therefore substantially negative. In subsequent years, however, those declines were reversed and more than offset by increases. The patterns of these later increases moreover were very similar to those that took place in the immediate postwar years in what arguably was a catch up from the previous wartime inflation. The puzzle therefore is not that equities fail the test as inflation hedges, but that they take so long to pass.

These results and those of related tests are reported in section II below. As an entree to that discussion, we briefly review the theory and the previous literature in section I. Section III presents some broader conclusion suggested by these findings and outlines additional work to be undertaken on this subject.

I. Theoretical Considerations

In the standard textbook version of the Fisher equation the nominal yield on a bond is decomposed into two components, the anticipated rate of inflation over the life of that instrument and the ex ante, or anticipated, real interest rate on the bond. The real variable, the ex ante real interests rate, thus is immune to changes in the nominal variable, inflation, provided that the latter are anticipated. Employing that same logic, researchers, have posited a direct counterpart of this bond-market Fisher equation in studying equity market behavior (see, for example, Fama and Schwert, 1977). The nominal return on equities R^e_t is assumed to consist of two components, the *ex ante* real return ρ^e_t and the rate of inflation that agents anticipate $E[\hat{p}_t \mid \Theta]$ given their information set Θ

The difference between equities and bonds, however, is that the income stream yielded by bonds is fixed before the fact, while that yielded by equities is not, since firms' nominal earnings and hence equity prices and dividends can be expected to rise in line with product prices. This difference in the characteristics of the two types of instruments means that the conventional Fisher equation should be rewritten to allow for both unanticipated as well as anticipated inflation:

$$R_t^e = \rho_t^e + E[\hat{p}_t \mid \Theta] + \{\hat{p}_t - E[\hat{p}_t \mid \Theta]\}, \tag{1}$$

or simply,

$$R_t^e = \rho_t^e + \hat{p}_t, \tag{2}$$

where $\hat{p}_t$ is the actual rate of inflation and the bracketed term on the far right side of (1) is therefore the unanticipated rate of inflation.

The reasoning underlying the relationship between returns and inflation could also be applied to the level of equity prices and the price level. In the standard model of equity pricing the nominal price of a common stock in the absence of earnings growth, P^e is simply the present discounted value of the infinitely lived nominal stream of nominal dividends, D:

$$P_t^e = D_t^e / R_t^e. \tag{3}$$

Assuming neutrality, we can write an analogous relation linking the nominal equity price, real dividends and the price level which in logarithmic form becomes:

$$p_t^e = \delta_t^e + p_t, \tag{4}$$

where lowercase variables are the natural logs of their uppercase counterparts and $\delta = \ln\{(D/P)/R^e\}$.

Tests of the relationship between equity returns and inflation have generally used regression analogues to equation (1) or (2) as their basis, equations taking the form either of

$$R_t^e = \alpha + \beta E[\hat{p}_t \mid \Theta] + \epsilon_t, \tag{1a}$$

or of

$$R_t^e = \alpha + \beta E[\hat{p}_t \mid \Theta] + \lambda\{\hat{p}_t - E[\hat{p}_t \mid \Theta]\} + \epsilon_t, \tag{2a}$$

where α, β and λ are coefficients to be estimated and ϵ is an error term. In the terminology of Fama and Schwert (1977), equities are hedges against anticipated inflation if β=1, and a complete inflation hedge if β=λ=1.

A considerable number of studies have rejected one, and generally both of these hypotheses, finding low and even negative coefficients for β and λ. In addition to Fama and Schwert, other studies reporting such results include Guletkin (1983) and Kaul (1987). Indeed a sizable literature has developed on this subject, the object of which has been to test various explanations for this failure.[1]

Overlooked in most of this recent work, however, are the details of Irving Fisher's analysis of these issues. In contrast to modern interpretations of the Fisher effect, which usually simply posit the Fisher equation and justify it on the grounds of market efficiency, Fisher took care to trace out the process by which interest rates adjusted to inflation. Perhaps influenced by the weak empirical results that he had obtained, Fisher viewed the adjustment as a circuitous and lengthy affair. As a practical matter he considered such adjustments more often than not to be incomplete and characterized by informational problems including asymmetries between the more sophisticated business sector borrowing funds and the less sophisticated consumers ultimately doing the lending. To our knowledge, he engaged in no formal analysis of equity-market behavior. Much of the current discussion of the Fisher effect is therefore historically inaccurate. More important, it ignores issues that Fisher himself quite evidently regarded as crucial, and which in fact may be a good deal more so than generally believed.

The recent literature on the bond-market Fisher equation is a case in point. Problems related to expectations' formation and differences between short-run and long-run behavior are highlighted in these studies. For example, Mishkin (1993) using bivariate cointegration tests finds a long-run Fisher effect, but no short-run effect. Crowder and Hoffman (1996) using more powerful multivariate analogues of such tests corroborate these findings. Evans and Lewis (1995) do provide evidence of a short-run effect, but they model inflation expectations using a Markov switching model to allow for the effects of changes in the monetary regime, rather than relying on the simple proxies used in

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[1] These studies include Feldstein (1980), Fama (1981), and Geske and Roll (1983).

most other studies.

Several recent studies have presented more positive evidence on the relation between equity returns and inflation. These include Boudokh and Richardson (1993) and Lothian and Simaan (1998).[2] The first uses long historical time series for the U.S. and the U.K.; the second, multi-country time series for the period since 1973. Both report results that are consistent with a long-run, but not a short-run, relationship. One question that can be raised with regard to Boudokh and Richardson's findings, however, is whether they are applicable to recent decades, since it is quite possible that the structure of the relationship has changed and that this structural change has gone undetected in their long data set, which by its very nature gives heavy weight to earlier periods. The evidence presented by Lothian and Simaan suggest that such a change has not occurred, but that paper is subject to a potential problem of another sort. It uses very simple techniques, presenting graphical evidence and bivariate regressions of period averaged data for the 18 OECD countries studied.[3] As the authors point out, the results could be purely artifacts of the period over which the data had been averaged.

II. Empirical Results

We use two bodies of data in the paper. The first is a panel data set consisting of annual series for the United States and 13 other OECD countries over the period 1949 to 1999. The price level is measured by the consumer price index or similar cost-of-living index; the equity-price index used in computing nominal equity returns is whatever index is reported by the International Monetary Fund in their International Financial Statistics, generally, though not always, an index of industrial

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[2] Two other studies that bear mention are Choudhry (2000) and Solnik and Solnik (1999). Both provide evidence of a positive relationship between nominal equity returns and various measures of anticipated inflation. Neither, however, shows anything close to complete adjustment to actual inflation.

[3] Their conclusions are in fact in line with those reached earlier by Cagan (1974) in one of the few earlier studies that have provided at least some support the hypothesis of invariance of real equity returns to inflation. Cagan used average rates of change of equity prices from the late 1930s (the starting dates differ slightly from one country to the next) to 1969 as his units of observation. He found a positive relation between the two but concluded that there were extremely long lags in the adjustment of equity prices to inflation. As Kaul (1987) pointed out, however, adjustment still appeared to be incomplete in the Cagan data, since for many of Cagan's countries viewed individually the average rate of inflation exceeded the average rate of change of nominal equity prices.

share prices.[4] Since we want data for a large number of countries over as lengthy a period as possible we use these data even though dividends are omitted.

Our second body of data is a long historical data set for the UK and the US. The UK data are for the period 1790 to 2000; the US data are for the somewhat shorter period 1800 to 2000. As will become apparent from the discussion immediately below, long historical data of this sort can be a useful complement to the panel data.

II.A. Overview of the Panel Data

Figures 1 and 2 and Table 1 provide an overview of behavior in the panel data. Shown in Figure 1 is the average of the de-meaned log of the real equity price indexes in the 14 counties and one-standard-deviation bands about those averages. Shown in Figure 2 are comparable figures for the first differences of this ratio, our proxy for real returns. Presented in Table 1 are means, standard deviations and first order autocorrelation coefficients for the changes in log real equity prices.

Several features of these charts deserve comment. The first is the upward drift in average real equity prices over the full time period. We see this quite clearly in Figure 1. The bands plotted in that chart suggest, moreover, that a similar pattern must have prevailed in most of the countries individually. This is confirmed further by a glance at the country means listed in Table 1. The changes in log real equity prices range from an average of 0.4 per cent per annum for New Zealand to an average of 6.0 per cent per annum for Germany. A second feature of the data brought out by both Figure 1 and Figure 2 is the series of long swings that characterize these data. For roughly the first two decades real equity prices increased. Over the next decade and a half or so, in contrast, they decreased, only to increase once again over the latest decade and a half.

The result is two protracted cycle-like movements. In the terminology used in the National Bureau of Economic Research's business cycle analysis, there is an "expansion phase" running from 1949 to 1969, the year in which the average series peaks, followed by a "contraction phase" ending with a trough in 1982. If we arbitrarily call the year 2000 the peak and assume that 1939 marked an

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[4] Figures for the cost-of-living indexes are yearly averages as listed in line 64 of the International Financial Statistics; the equity price indexes are yearly average series listed in line 62. Because the IMF equity data do not include dividends, we use measures of the average rates of growth of equity prices, as a proxy for equity returns.

earlier peak (as in Cagan, 1975), we see two approximately thirty-year, peak-to-peak cycles with expansions that in both instances are approximately twice as long as the contractions.

In making such calculations, we are not hinting at a structural explanation for the two sets of movements. The object is simply to provide a description of behavior over this period. What it shows – and this is the third important point brought out by the charts – is the paucity of truly independent movements in the data over these fifty years. This is *a fortiori* the case for the period from the early 1960s through the mid-1980s that has been most heavily studied. It is not surprising that negative findings have been reported for that sample. Real equity prices declined during these years in the midst of high and, in many countries secularly rising, inflation. While such behavior does not now appear representative of experience over the sample period as a whole, this was not at all clear fifteen or so years ago. Still dominating the data at that point was the downward swing in real equity prices that began in the mid-1970s. That being so, no amount of econometric ingenuity could have solved the problem and come up with positive findings using data for that sample alone.

This last feature of the data is a further illustration of the phenomenon that has been discussed very heavily in the literature on purchasing power parity: the importance of data span, as opposed to data frequency. Distinguishing between slow mean reversion and unit-root behavior has proven almost impossible using conventional tests and single-country time series; extremely long time series or panel data are required.[5] This should not be surprising. If the deviations from long-term equilibrium are persistent, the number of episodes in which such deviations occur rather than year-by-year or quarter-to-to-quarter behavior within those episodes will contain the more meaningful information about equilibrium behavior.

II.B. Unit Root Tests for Nominal Equity Prices and CPIs

Since it is clear that equity prices and the general level of prices at best are only very weakly connected over the short run, we focus on the long run. The specific issues that we address are whether over such time spans, the two do in fact bear the positive relationship suggested by theory and whether nominal equity prices have (at least) kept pace with inflation.

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[5] See Lothian and Taylor (1996, 1997) for discussions of this issue in the context of the PPP relationship. General discussions in the time-series econometric literature include Shiller and Perron (1985) and Hakkio and Rush (1991).

A standard way of addressing long-term behavior is in terms of cointegration analysis, since as Engel and Granger (1987) have demonstrated, there is a direct correspondence between the econometric concept of cointegration and the economic concept of long-run equilibrium. A necessary condition for two series to be cointegrated is that they both integrated of the same order. We therefore first conducted unit root tests for log equity prices, log price levels and their first differences. We used both augmented Dickey Fuller tests and Phillips Perron tests in each instance. Test results for the log levels of the CPI and the equity price index are reported in Tables 2a and 2b respectively. In no instance, could we reject the unit root null for either of the two series. We therefore conducted similar tests for the first differences of the logs of the two series. These are reported in Tables 3a and 3b. Here the results were mixed. For the difference in log nominal equity prices we could always reject the unit root null. For the CPI the results varied by country. Log equity prices therefore appear to be I(1); while for the difference in the log CPI, the evidence is less clear.

Given this indeterminancy we ran tests for both the levels and the first differences of real equity prices. In the first case, we allowed for a deterministic trend in real equity prices. The second presupposes the existence of a stochastic trend. Arguably, it is the more realistic of the two. Any upward drift in real equity prices is, we suspect, most likely due to factors like productivity shocks the effects of which are one-off but highly persistent. They will therefore be better approximated by a stochastic trend rather than a deterministic trend.

II.C. Unit Root Tests for Real Equity Prices

In analyzing real equity prices, the question that we address is whether the two series do in fact share a common trend – whether CPIs and equity price indexes are cointegrated and whether, as theory suggests, they bear a one-to-one long-run relationship to one another. The simplest way to test for such a relationship is to examine the behavior of log real equity prices. To see why this is so consider the following regression:

$$p^e_t = \alpha + \beta\, p_t + \mu_t, \qquad (5)$$

where α and β are coefficients and μ is the error term.

Suppose p^e and p are cointegrated and related one-to-one. Then μ will be stationary and β will be unity. One way to test whether these conditions hold is to impose the constraint β=1, in which case (5) becomes:

$$rp^e_t = \alpha + \mu_t, \tag{6}$$

where $rp^e = p^e - p$, the log of real equity prices, and then to apply a unit root test to this ratio. In the simplest case, this is a test of the hypothesis $\lambda = 1$ in the autoregression:

$$rp^e_t = \alpha + \lambda rp^e_{t-1} + \eta_t, \tag{7}$$

If we reject this hypothesis, the implication is that rp^e is mean reverting and does so at a speed of 1-λ per period.

Hamilton (1994) argues that this is more powerful than the conventional two-step Engle and Granger cointegration tests. This increased power comes about, however, as a result of imposing the constraint β= 1 in equation (5). This can be tested using the procedure described in Hamilton (1994, pp. 608-613).

We conducted these tests for each of the countries individually both with and without a deterministic trend in the regression. We report these results in the upper half of Table 4a. In no instance could we reject the unit root null. In all instances, however, we obtained point estimates of λ that were less than one, but in most instances not substantially so. This is consistent with very slow reversion to trend, but certainly offers little proof that such behavior is actually characteristic of these data. It could be that it is and that low test power is obscuring that fact; alternatively real equity prices could contain a permanent component.

To try to increase test power, we treat the data for the 14 countries as a panel data set and use the Im et al (1997) t-bar test. The t-bar test statistic is the average of the individual t statistics from augmented Dickey-Fuller (ADF) unit root tests for each of the 14 countries in our study standardized by subtracting the expected value of the distribution of the t statistics under the null hypothesis of a unit root and dividing by the standard deviation. Estimates for the expected values and variances are calculated by Im et al (1997) and reported in Table 2 of their paper for ADF tests with and without a trend for lag lengths from 0 to 8. This standardized t-bar statistic is shown to be distributed standard normal. The results for our panel are presented in the lower half of Table 4a. These tests reject the unit-root null in three of four instances when a time trend is included in the test regression, but in only one instance when the trend is excluded. There are, moreover, several potential econometric problems surrounding the use of these tests that could cause them to be misleading[6] For these and for other reasons we turn to the differenced data.

The results of these tests are presented in Table 4b. Here the evidence is unambiguous. In all of the countries viewed individually, we can reject the unit root null at very high levels of significance in the differenced data. Hence even though there may permanent shocks may affect the relationship between the level of nominal equity prices and the price level, their rates of change converge.

II.D. Evidence from the long-term time series

Fifty years worth of observations on the surface appears to be more than sufficient for valid statistical inference. As earlier discussion indicated, the data examined here are dominated by only two long-lived cycle-like movements. Separating trend from cycle and discerning the true long-term drift of real equity prices is, therefore, extremely difficult. The use of multi-country data solves the problem to some degree but not completely, since there is likely to be at least some cross-country correlation in the data, and since in any event such data sets are still too short,

The historical data for the UK and the US mentioned above are an additional way of dealing with this issue. Figures 4 and 5 show plots of the log real equity price indexes for the UK and the US and their two nominal components over the periods 1790 to 2000 and 1800 to 2000, respectively. Figure 3 provides a plot of one county's real equity price series against the other's. Several features of these charts deserve comment. The first is the substantial and accelerating upward trends in the nominal series for both countries particularly during the course of the last century. The other is the more moderate, but still upward, drift in both countries' real equity price series. As in the panel data, equities again prove to be an inflation hedge. Again, however, it takes an exceedingly long time for this to happen. At shorter, though still quite lengthy time horizons, there is a negative relationship

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[6] One such problem is cross-country correlation (see O'Connell, 1998); a second is the possibility that one country (or a few countries) are responsible for the rejection (see Taylor and Sarno, 1998).

between price levels and the real equity prices. Following accelerations in the price level we see long cyclical-type downswings in real equity prices lasting years. We see this in both countries during both of the two World Wars, the US Civil War, and the Napoleonic Wars, in the US a few decades later in the nineteenth century and, as earlier mentioned, in both countries during the two decades following the breakdown of Bretton Woods.

More formal evidence on the behavior of real equity prices in the two countries over this long period is presented in Table 5. Shown there are the results of unit root tests for both the levels and first differences of log real equity prices. The results for the levels are mixed – rejection for the UK but not the US. As with the panel data, the results for the first differences show strong rejections in all instances. The inference to be drawn is therefore similar: Over the long run we see inflation and nominal equity returns converge as well as a tendency for nominal returns to drift up relative to inflation.

III. Summary and conclusions

In this paper we examine data for stock prices and price levels of 14 developed countries during the post-WWII era and compare behavior in that sample with behavior over the past two centuries in the UK and the US. The evidence derived from both is virtually the same.

Contrary to much of the literature of the past several decades, we find that equities are in fact an inflation hedge. Our results suggest, however, that this is only the case over very long periods. They, therefore, corroborate the findings reported by Cagan (197 4) in a paper written as the decline in real equity prices of the 1970s had just gotten underway. Examining a multi-country data set similar to ours, he concluded that equity markets did adjust to inflation, but that the adjustment period lasted more than a decade.

The issue that needs to be addressed is why this adjustment process is so drawn out and, hence, why departures of real equity prices from neutrality are so long lived. Viewed from the perspective of the modern literature in both finance and in economics more broadly defined, such behavior is a puzzle of the first dimension. Market efficiency and the rational expectation's hypothesis are usually taken to imply quick adjustment of economic variables to shocks and neutral behavior of real variables even over rather short periods. This is not what either we or Cagan have found. On the contrary, both sets of results are much more in line with those reported very much earlier by Irving Fisher in his empirical analysis of what subsequently has been termed "the Fisher effect." Commenting on the extremely slow and incomplete adjustment of nominal interest rates to inflation, he wrote (1930, p. 416) "[I]f adjustment were perfect, unhindered by any failure to foresee future changes in the purchasing power of money or by custom or law or any other impediment, we should have found a very different set of facts."

The alternative explanation sees the negative relation between returns and inflation as behavioral phenomenon. Barnes, Boyd, and Smith (1999) argue that inflation is non-neutral, that increases in inflation give rise to capital market inefficiencies and hence adversely affect investment and real income. The appearance of long lags in adjustment in their view is really two separate movements – a decline in real equity prices brought about by the increase in inflation and resultant decline in real income, followed by a rise in real equity prices due to the subsequent decline in inflation and rise in real income.

Friedman, in contrast, sees the variability of inflation – rather than the rate *per se* – as the problem. More variable inflation, in his view, mean less predictable inflation. The result is a decrease in the efficiency of the price system, which in turn leads to lower real income and higher unemployment. Friedman makes this argument in connection with the Phillips Curve, but it is also applicable here.

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Fig. 1. 14-Country average real equity prices

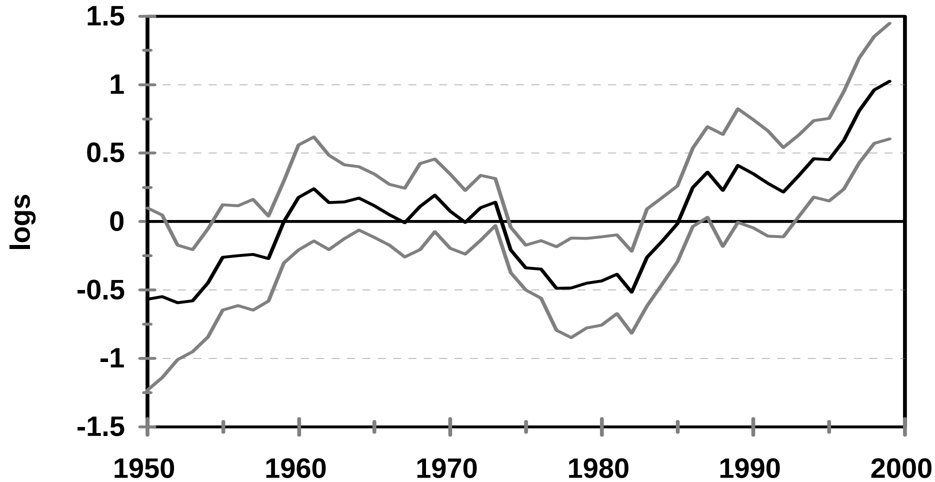

Fig. 2. 14-country average real equity returns

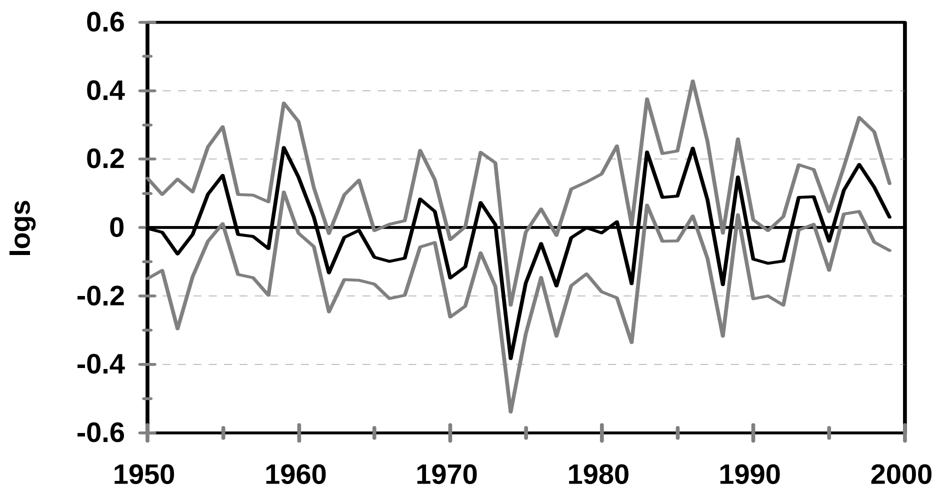

Fig. 3. US and UK real equity prices,1800-2000

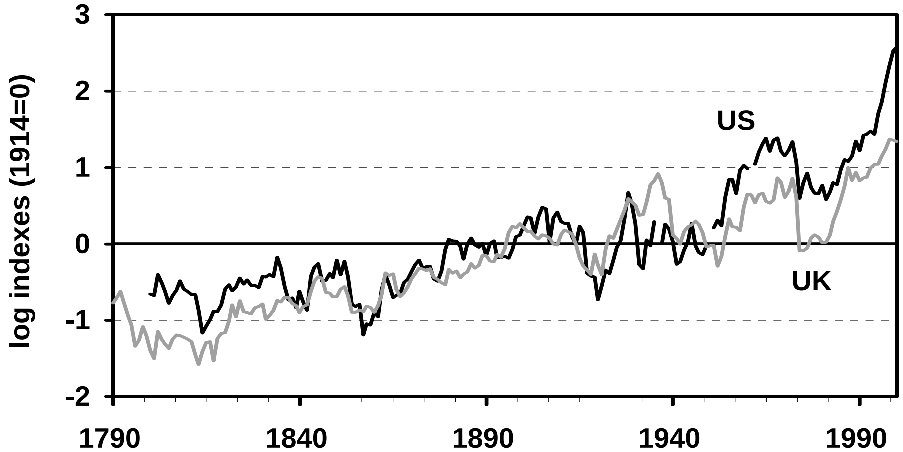

Fig. 4. UK consumer prices, nomina and real equity prices, 1790-2000

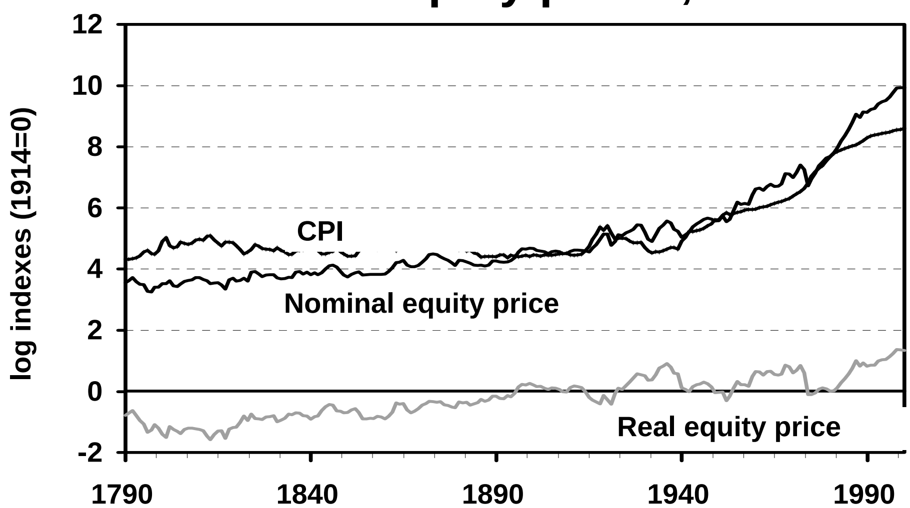

Fig. 5. US consumer prices, nominal and real equity prices1800-2000

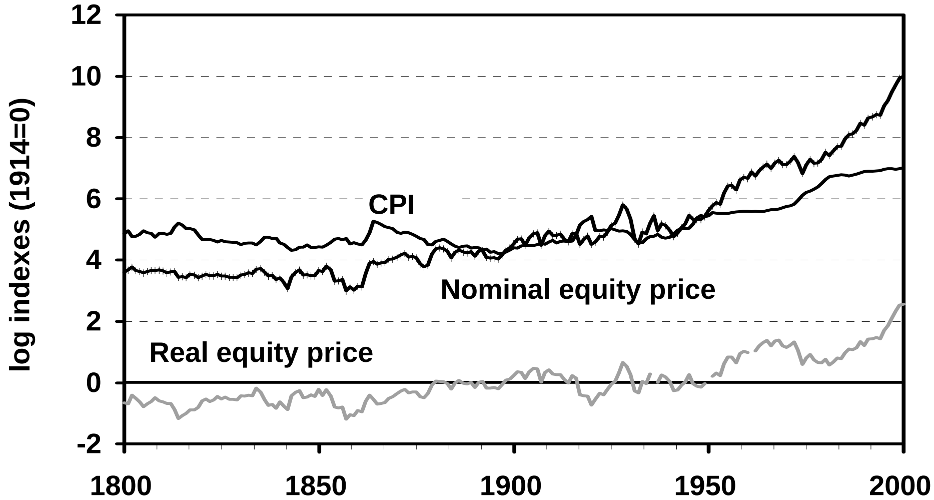

Table 1. Summary statistics for real equity returns, 1950-1999

	Mean	Std. Dev.	Rho
Australia	1.1	16.9	0.188
Canada	2.9	13.8	-0.003
Denmark	2.3	17.5	0.056
France	3.9	18.5	0.189
Germany	6.0	19.3	0.362
Ireland	2.7	20.0	0.338
Italy	2.0	24.5	0.333
Japan	4.8	19.2	0.226
Netherlands	3.9	16.9	0.456
New Zealand	0.4	17.6	0.186
Norway	1.4	18.9	0.154
Sweden	5.6	17.5	0.120
UK	3.0	15.8	0.196
US	5.4	13.5	0.190
14-Country average	3.2	17.9	0.214

Note: Means and standard deviations were computed for first differences of log real equity prices and have been multiplied by 100. Rho is the first order autocorrelation coefficient.

Table 2a. Unit root tests for log level of CPI; annual data, 1949 to 1999

$$\Delta p_t = a_0 + a_1 \; p_{t-1} + \Sigma\, b_i \; \Delta p_{t-i}$$

	ADF test	P-P test
Australia	-0.024	-0.635
Canada	-0.490	0.176
Denmark	-0.683	-0.554
France	-0.343	-1.020
Germany	-0.053	0.369
Ireland	-0.856	-0.229
Italy	-0.696	0.404
Japan	-2.158	-1.234
Netherlands	-1.161	-1.181
New Zealand	-0.622	-0.070
Norway	-0.165	-0.662
Sweden	-0.792	-0.196
UK	-0.753	0.074
US	-0.550	0.634

Critical values for Dickey-Fuller, Augmented Dickey-Fuller tests and the Phillips-Perron tests at 0.01 and 0.05 levels of significance are -3.58 and -2.93. Lag length on ADF chosen using Akaike Criterion.

Table 2b. Unit root tests for log level of nominal equity prices; annual data, 1949 to 1999

$$\Delta p^e_t = a_0 + a_1\ p^e_{t-1} + \Sigma\, b_i\ \Delta p^e_{t-i}$$

	ADF test	P-P test
Australia	0.150	0.238
Canada	-0.741	-0.699
Denmark	0.289	0.408
France	-0.192	-0.340
Germany	-1.148	-1.660
Ireland	1.043	1.158
Italy	-0.382	-0.270
Japan	-1.665	-1.640
Netherlands	0.495	0.817
New Zealand	0.369	0.241
Norway	1.136	1.313
Sweden	1.209	1.231
UK	0.688	0.738
US	0.651	0.662

Critical values for Dickey-Fuller, Augmented Dickey-Fuller and Phillips-Perron tests at 0.01 and 0.05 levels of significance are -3.58 and -2.93. Lag length on ADF chosen using Akaike Criterion.

Table 3a. Unit root tests for first difference of log CPI; annual data, 1949 to 1999

$$\Delta^2 p_t = a_0 + a_1 \ \Delta p_{t-1} \ + \Sigma \, b_i \ \Delta^2 p_{t-i}$$

	ADF test	P-P test
Australia	-2.203	-2.336
Canada	-2.347	-2.298
Denmark	-2.478	-2.307
France	-3.391	-3.284
Germany	-6.301	-6.275
Ireland	-2.309	-2.040
Italy	-2.503	-2.496
Japan	-4.886	-5.198
Netherlands	-3.646	-3.490
New Zealand	-1.922	-1.741
Norway	-3.246	-2.913
Sweden	-3.528	-3.592
UK	-2.349	-2.278
US	-3.100	-3.150

Critical values for Dickey-Fuller, Augmented Dickey-Fuller and Phillips-Perron tests at 0.01 and 0.05 levels of significance are -3.58 and -2.93.Lag length on ADF chosen using Akaike Criterion.

Table 3b. Unit root tests for first difference of log nominal equity price; annual data, 1949 to 1999

$$\Delta^2 p^e_t = a_0 + a_1 \ \Delta p^e_{t-1} \ + \Sigma\, b_i \ \Delta^2 p^e_{t-i}$$

	ADF	P-P
Australia	-5.939	-6.070
Canada	-7.415	-7.690
Denmark	-6.276	-6.229
France	-5.820	-5.897
Germany	-3.264	-4.767
Ireland	-5.405	-5.201
Italy	-5.146	-4.858
Japan	-6.007	-6.048
Netherlands	-4.949	-4.489
New Zealand	-3.698	-5.608
Norway	-6.013	-5.961
Sweden	-6.305	-6.304
UK	-6.161	-6.312
US	-6.306	-6.277

Critical values for Dickey-Fuller, Augmented Dickey-Fuller and Phillips-Perron tests at 0.01 and 0.05 levels of significance are -3.58 and -2.93. Lag length on ADF chosen using Akaike Criterion.

Table 4a. Unit root tests for log level of real equity price; annual data, 1949 to 1999

$$\Delta rp^e_t = a_0 + a_1\ rp^e_{t-1} + a_2 t + \Sigma\, b_i\ \Delta rp^e_{t-i}$$

	ADF test no trend	ADF test with trend	P-P test no trend	P-P test no trend
Australia	-2.279	-1.990	-1.881	-2.167
Canada	-2.406	-2.860	-2.380	-2.875
Denmark	-0.702	-1.835	-0.620	-1.813
France	-1.319	-1.456	-1.121	-1.356
Germany	-1.791	-2.270	-2.285	-2.397
Ireland	-0.735	-2.344	-0.720	-1.607
Italy	-2.256	-2.292	-1.784	-1.776
Japan	-1.781	-3.077	-1.687	-2.093
Netherlands	-1.047	-1.392	-0.099	-0.667
New Zealand	-2.836	-2.874	-2.150	-2.003
Norway	-0.567	-0.532	0.805	-0.451
Sweden	0.839	-0.614	-0.714	-0.743
UK	-0.735	-1.901	-0.750	-1.544
US	-0.379	-0.849	-0.610	-1.193

Critical values for Dickey-Fuller, Augmented Dickey-Fuller and Phillips-Perron tests without trend at 0.01 and 0.05 levels of significance are -3.58 and -2.93. Critical values for Dickey-Fuller, Augmented Dickey-Fuller and Phillips-Perron tests with trend at 0.01 and 0.05 levels of significance are -4.15 to -4.17 and -3.50 to 3.51, depending upon the exact form of the test regression. Lag length on ADF chosen using Akaike Criterion.

t-bar test

	0 lags	1 lag	2 lags	3 lags
ADF tests; with intercept, no trend	2.528	0.433	1.841	1.403
ADF tests; with intercept and trend	3.666	1.014	2.924	2.281

The t-bar statistic (Im et al, 1997) is distributed standard normal. Critical values for 2 tail test for 0.01 and 0.05 level of significance are 2.54 and 1.96.

Table 4b. Unit root tests for first difference of log real equity price; annual data, 1949 to 1999

$$\Delta^2 rp^e_t = a_0 + a_1 \, \Delta rp^e_{t-1} + \Sigma \, b_i \, \Delta rp^e_{t-i}$$

	ADF test	P-P test
Australia	-5.432	-5.573
Canada	-6.965	-7.043
Denmark	-5.748	-6.415
France	-5.652	-5.731
Germany	-6.014	-4.523
Ireland	-5.116	-4.702
Italy	-4.952	-4.730
Japan	-3.200	-5.700
Netherlands	-4.506	-4.145
New Zealand	-3.683	-5.645
Norway	-5.872	-5.808
Sweden	-6.079	-6.068
UK	-5.576	-5.493
US	-5.664	-5.641

Critical values for Dickey-Fuller, Augmented Dickey-Fuller and Phillips-Perron tests at 0.01 and 0.05 levels of significance are -3.58 and -2.93. Lag length on ADF chosen using Akaike Criterion

Table 5. Unit root tests for log levels and first differences of log levels of real equity prices; long-term UK and US times series.

	P-P test	ADF test
Levels		
UK	-3.954[b]	-4.266[a]
US	-2.269	-2.279
Differences		
UK	-12.418[a]	-10.682[a]
US	-12.986[a]	-8.062[a]

Note: t values are listed beneath the coefficients; the superscripts a and b indicate significance at the .01 and .05 levels.