THE PRICE OF GOLD AND STOCK PRICE INDICES FOR THE UNITED STATES

by

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Abstract

This paper provides empirical evidence on the relationship between the price of gold and stock price indices for the United States over the period beginning in January 1991 and ending in October 2001. Four gold prices and six stock price indices are used. The short-run correlation between returns on gold and returns on US stock price indices is small and negative and for some series and time periods insignificantly different from zero. All of the gold prices and US stock price indices are I(1). Over the period examined, gold prices and US stock price indices are not cointegrated. Granger causality tests find evidence of unidirectional causality from US stock returns to returns on the gold price set in the London morning fixing and the closing price. For the price set in the afternoon fixing, there is clear evidence of feedback between the markets for gold and US stocks.

Keywords: Gold price, correlation coefficients, stock price indices, cointegration, Granger causality.

I. INTRODUCTION

In the immediate aftermath of the catastrophic events of September 11, 2001 the gold price set in the London afternoon fixing increased by 5.6% and, by the close of the day's trading, the US$-adjusted FTSE Gold Mines Index had risen by 6.4%. Stock markets around the world were affected to different degrees. The Hong Kong stock market was down 9.2%. In Japan, the Nikkei 225 decreased by 6.9%. The Australian All Ordinaries was down 4.2%—the same as the Johannesburg Stock Exchange All Share Index. However, in Germany the DAX increased by 1.4% and in France the CAC 40 rose 1.3%. New York markets were closed for four days. In the first week of trading after they reopened, the Dow Jones Industrial Average fell by 15.4% and the Nasdaq by 17.5%. The US$-adjusted FTSE All Share Index decreased by 9.0% and the Swiss Franc, viewed as a safe currency in times of crisis, appreciated by 4.4% against the US$. On September 21st, the gold price set in the London afternoon fixing was US$ 292.5 per troy ounce compared with US$ 271.5 per troy ounce on September 10th, an increase of 7.45%.

There is clear evidence that in a time of crisis, as equity prices fall the price of gold rises. With the more uncertain economic environment, attention focused on gold as a safe haven:

> The metal is reassuringly tangible; it tends to be a good hedge against inflation; it tends to move in the opposite direction to shares and bonds; and, unlike most financial assets, it does not represent anyone else's liability.
> Moreover, several analysts have upgraded their predictions of where prices are headed.
>
> Adrienne Roberts *FT Personal Finance*, October 27th 2001, p 14.

This paper examines empirical evidence on the prices of gold and US stocks over the last decade. Four gold prices are used. Two of them, the US$ prices determined in the 10.30 am and 3 pm fixings held at the offices of N. M. Rothschild & Sons in New Court in the City of London, are accepted worldwide. Since the London closing price is determined by additional information, this price is also included. Gold prices set in New York are also used. Since many US stock price indices are widely quoted, although to differing degrees, and because it is possible that any one index can be unrepresentative of others, six different US stock price indices are employed.

In a study of this type, it is important to focus on appropriate relationships. For example, suppose *G* is the price of gold and *S* is a stock price index (both in US$). If neither series is trended and both have a constant variance then one might focus on a linear relationship of the form

$$G = \alpha_0 + \alpha_1 S + u_t \qquad (1)$$

or

$$\ln G = \beta_0 + \beta_1 \ln S + v_t \qquad (2)$$

in which u_t, v_t are disturbance terms which satisfy the usual classical assumptions. However, the price of gold and stock price indices for the US are trended over the period under consideration. In these circumstances, relationships of the form given by (1) and (2) can generate spurious results. More specifically, a regression between two independent nonstationary variables can result in an apparently significant relationship with apparently high correlation between variables even though there is no linear relationship between *G* and *S*. Unless there is a long-run equilibrium relationship between the variables—that is, unless they are cointegrated—a linear relationship between *G* and *S* is spurious, nonsense and the results are of no use. This paper avoids problems of spurious regressions in two ways. First, correlation coefficients between returns on gold and returns on stocks are examined and not correlations between the prices of gold and stocks. Secondly, the time series properties of the data are examined and appropriate tests of long-run equilibrium presented.

The rest of this paper is organised as follows. Section II discusses the empirical methodology employed. Section III describes the data and their characteristics. In section IV the results are presented. Section V provides a brief conclusion.

II. METHODOLOGY

There can be both short-run and long-run relationships between financial time series. Correlation coefficients are widely used for examining short-run co-movements between stock price indices (see, for example, Dwyer and Hafer, 1993, Erb *et al*, 1994, and Peiró *et al*, 1998). The population correlation coefficient, ρ, $(-1 \le \rho \le 1)$ measures the degree of linear association between two random variables and is the ratio of the covariance between them and the product of their standard deviations. When sample information is used, we have the coefficient of linear correlation, 'the correlation coefficient', Pearson's *r*. With financial markets, correlation coefficients are usually calculated between returns. If *G* is the price of gold and *S* is a stock price index (both in US$)

$$r = \frac{\mathrm{cov}(R_g, R_s)}{\sigma_{R_g}\sigma_{R_s}} \qquad (3)$$

where

$$R_g = \Delta g = \Delta \ln G$$

and σ_{R_g} is the standard deviation of gold returns, etc.

Two hypotheses are tested. The first of these is $H_0\colon \rho = 0$ against $H_1\colon \rho \neq 0$. Under the null hypothesis, the random variable $z = \frac{1}{2}\ln\left(\frac{1 + r}{1 - r}\right)$ is asymptotically distributed as $N\left(0, \frac{1}{T-3}\right)$ where T is the sample size. Hence, the test statistic

$$\frac{\ln\left(\frac{1 + r}{1 - r}\right)}{2/\sqrt{T - 3}} \qquad (4)$$

has an asymptotic standard normal distribution if the null hypothesis is true.

Kendall and Stewart (1961) show that for two independently sampled populations, $H_0\colon \rho_1 = \rho_2$ can be tested against $H_1\colon \rho_1 \neq \rho_2$. In large samples, $z_1 - z_2 = \frac{1}{2}\ln\frac{1 + r_1}{1 - r_1} - \frac{1}{2}\ln\frac{1 + r_2}{1 - r_2}$ is normally distributed with zero mean and variance $1/(T_1 - 3) + 1/(T_2 - 3)$ where T_1 and T_2 are the sample sizes. The test statistic

$$\frac{\ln\left[\frac{(1 + r_1)}{(1 - r_1)}\frac{(1 - r_2)}{(1 + r_2)}\right]}{2\sqrt{\frac{1}{T_1 - 3} + \frac{1}{T_2 - 3}}} \qquad (5)$$

has an asymptotic standard normal distribution under the null hypothesis of equal correlation coefficients.

When there is long-run equilibrium among financial time series, they share common stochastic trends. This possibility arises with two or more time series which are I(d), $d \geq 1$, in which case a linear combination can be cointegrated. Before testing for cointegration, it is necessary to establish the order of integration of the series. Tests of orders of integration are carried out with Phillips and Perron (1988) unit root tests for the logarithms of the series, g and s. These tests are implemented sequentially from the general model which includes both an intercept and time trend

$$y_t = \tilde{\mu} + \tilde{\beta}\left(t - \frac{T}{2}\right) + \tilde{\alpha} y_{t-1} + \tilde{\epsilon}_t \tag{6}$$

to the more specific model which has an intercept but no time trend

$$y_t = \mu^* + \alpha^* y_{t-1} + \epsilon_t^* \tag{7}$$

and the model with neither intercept nor trend

$$y_t = \hat{\alpha} y_{t-1} + \hat{\epsilon}_t. \tag{8}$$

Phillips-Perron tests involve non-parametric corrections to test statistics and allow for possible autocorrelation and heteroscedasticity in the residuals of the regression on which the test is based—characteristics frequently found with financial time series.

The Engle-Granger approach is used to test for the existence of a long-run equilibrium relationship between I(1) gold prices and stock price indices based on the relationship

$$g_t = \alpha_0 + \beta_0 s_t + w_t \tag{9}$$

where g_t and s_t are the logarithms of the price of gold and a stock market price index at time t and w_t is the disequilibrium error, that is, the deviation from long-run equilibrium. Conventional, cointegrating regression augmented Dickey-Fuller tests are used.

Following Engle and Granger (1987), if both g_t and s_t are cointegrated then they are generated by Error-Correction Models (ECMs) of the form

$$R_{g,t} = \alpha_1 + \sum_{j=1}^{n} \beta_{1j} R_{g,t-j} + \sum_{j=1}^{n} \gamma_{1j} R_{s,t-j} + \lambda_1 w_{g,t-1} + v_{g,t} \tag{10}$$

and

$$R_{s,t} = \alpha_2 + \sum_{j=1}^{n} \beta_{2j} R_{s,t-j} + \sum_{j=1}^{n} \gamma_{2j} R_{g,t-j} + \lambda_2 w_{s,t-1} + v_{s,t} \tag{11}$$

in which $R_{g,t} = g_t - g_{t-1}$, $R_{s,t} = s_t - s_{t-1}$, the $v_{i,t}$ are stationary disturbances and the $w_{i,t-1}$ error-correction terms. The ECMs are useful because short- and long-run effects are separate and both can be estimated. The coefficients on lagged returns in equations (10) and (11), γ_{1j} and γ_{2j}, represent the short-run elasticities of $R_{g,t}$ and $R_{s,t}$ with respect to $R_{s,t}$ and $R_{g,t}$ respectively. The respective long-run elasticities are obtained from cointegrating regressions. The coefficients on the disequilibrium errors, λ_1 and λ_2, measure the speed of adjustment of g_t and s_t respectively to the error in the previous period. With cointegration, at least one of the $\lambda_i \neq 0$.

If the series are both I(1) and not cointegrated then there is no long-run equilibrium relationship between them. The regression in levels of g_t on s_t is spurious and $\lambda_1 = \lambda_2 = 0$. However, the first differences $R_{g,t}$ and $R_{s,t}$ are stationary and so can be used to examine short-run relationships.

Using (10) and (11), either with or without the error-correction terms as appropriate, Granger (1969) causality tests between gold and stock markets are implemented. Such tests are based on the idea that the future cannot cause the present or the past. If a change in stock returns occurs before a change in gold returns, that is, if changes in stock returns precede changes in gold returns then the former 'Granger-cause' the latter. These tests reveal, for example, whether lagged equity returns improve the accuracy of predictions of gold returns beyond that provided by lagged gold returns alone. There are two hypotheses based on equations (10) and (11):

H_{0A}: $R_{s,t}$ does not Granger-cause $R_{g,t}$ (that is, $\gamma_{1j} = 0$ for all j);

H_{0B}: $R_{g,t}$ does not Granger-cause $R_{s,t}$ (that is, $\gamma_{2j} = 0$ for all j). (12)

Tests of both hypotheses are implemented through F tests of

$\gamma_{11} = \gamma_{12} = \ldots \gamma_{1n} = 0$ and $\gamma_{21} = \gamma_{22} = \ldots \gamma_{2n} = 0$ against the alternatives that at least one γ_{1j} (or γ_{2j}) is not zero.

III. THE DATA AND THEIR PROPERTIES

Four gold prices and six US stock price indices are used. There are three London gold prices: those set at the 10.30 am and 3 pm fixings held at the offices of N. M. Rothschild & Sons in the City of London and the closing price. The Handy & Harmon series is determined in the US market. All of these gold prices are expressed in US$ per troy ounce.

A wide variety of US stock price indices is used. Some are widely quoted but track the price changes of a relatively small number of stocks; others are less well-known, at least in Europe, but cover much larger numbers of stocks. The Dow Jones Industrial Average tracks the prices of 30 large, widely-traded blue-chip stocks on the New York Stock Exchange (NYSE). The index is price-weighted and so a change in the price of a stock with a relatively high price changes the average more than the same change in the price of a relatively low-priced stock. This is probably the most widely-known stock market index in the world but it is not representative of the market as a whole. The NASDAQ Composite is a broadly-based index which tracks the performance of all stocks, including American Depositary Receipts (ADRs), traded on the Nasdaq National Market System and Nasdaq SmallCap Market. The New York Stock Exchange (NYSE) Composite tracks the prices of all NYSE-listed common stocks accounting for 85% of US market capitalization. Standard and Poor's 500 (S&P 500) Composite Stock Price Index covers the performance of 500 leading large capitalization stocks listed on the NYSE, the American Stock Exchange (AMEX) and the Nasdaq National Market System. The Russell 3000 Index tracks the performance of 3000 stocks of US domiciled corporations, consisting of the common stocks which are constituents of the Russell 1000 and Russell 2000 indices. The

Wilshire 5000 Equity Index is the most broadly-based of these indices and covers the performance of all actively traded US issues. Except for the Dow Jones, all indices have capitalization weights. For each stock, the price is multiplied by the number of shares outstanding. Hence, the larger the capitalization of a company, the greater the weight the stock has in the index regardless of whether the share price alone is high or low. The data begin in January 1991 and end in October 2001 and, where appropriate, daily, weekly and monthly frequencies are used. The sources are *EcoWin* and *Datastream*.

Table 1 reports descriptive statistics for daily returns. For the gold prices, average returns are negative, for the stock price indices they are positive. Under the hypothesis that returns are normally distributed, the coefficient of skewness is asymptotically distributed as $\mathrm{N}(0, 6/T)$ where T is the sample size. The distributions of returns on gold are skewed to the right. For the six US stock market indices, the distributions are negatively skewed. Under normality, the coefficient of excess kurtosis asymptotically follows a $\mathrm{N}(0, 24/T)$ distribution. For all series, the distributions of daily returns are leptokurtic, that is, they have higher peaks about the mean and thicker tails than the normal distribution. The evidence clearly rejects the hypothesis that daily returns on gold and US stock price indices are normally distributed.[1] This is consistent with earlier evidence for stock returns which found similar departures from normality (Peiró, 1994, and Aparicio and Estrada, 2001).

[1] Descriptive statistics for weekly and monthly returns, not reported here, show similar nonnormal characteristics.

IV. RESULTS

Table 2D reports contemporaneous correlation coefficients for daily returns for the entire period. None of the correlation coefficients involving returns on the gold price set in the London morning fixing or the Nasdaq Composite is significantly different from zero. All of the other correlation coefficients are negative, small in magnitude and significantly different zero. Slightly different results are obtained with weekly and monthly returns. Table 2W reports contemporaneous correlation coefficients for weekly returns over the same sample period. All of the point estimates are negative. Again, none of the correlation coefficients involving the Nasdaq Composite are significantly different from zero. However correlations with returns on the London morning gold price are now significant and those with the afternoon gold price insignificantly different from zero. None of the correlation coefficients for monthly returns, reported in Table 2M, are significantly different from zero.

One of the differences between holding gold and stocks is that only price fluctuations are relevant for gold whereas stocks provide a stream of dividend payments in addition to price fluctuations. This can be captured by the use of total return indices which cover both equity prices and dividends paid and reinvested. Indices of this type are frequently used in measuring performance. The percentage change in a total return index measures the total return in terms of the change in the capital value of the index and the reinvestment of gross dividend income in additional units of the index. However, allowing for dividend payments has negligible effect on the results. With monthly data, the correlation coefficient between returns on the gold price set in the afternoon fixing and the Dow Jones total return index is –0.1027, compared with –0.1013 for the corresponding stock price index. Using the S&P 500 total return index, the correlation coefficient is –0.1175 (compared with –0.1180). All of these four correlation coefficients are insignificantly different from zero.

The trading times of the New York and European markets overlap. Consider times in GMT and the 24 hour period ending with the close of New York markets at 21:00, New York's day *t* determining period. This period includes the 10:30 London gold price. New York markets open at 14:30, soon followed by the 15:00 London gold price and the 17:00 closing price in London. Normally, gold prices take between 5 and 10 minutes to set in the fixings. The information signalled by the price set in the 10:30 fixing is therefore soon superceded by the price set at 15:00 and so the latter is more relevant for New York markets. New York's determining period covers all of the gold prices and so there may be contemporaneous correlations between daily returns on these, and returns calculated from the closing values of New York stock market indices.

Now consider the 24 hour period ending with the London closing price of gold at 17:00, London's day *t* determining period. This includes, of course, all of the London gold prices *and* the previous day's closing stock market indices in New York. Yesterday's returns in New York may influence today's London gold prices.

Table 3D reports correlation coefficients between London gold returns in day *t* and returns on US stocks in day *t*-1. Comparing these with the correlation coefficients obtained with current US closing prices reported in Table 2D, the introduction of the one-day lag results in correlations with returns on the gold price set in the London morning fixing becoming more negative and significant. The previous day's closing prices in New York are the most recent 'news' from the US market at the time of the morning fixing. This result, therefore, confirms other evidence in Table 2D of a small significant negative correlation between returns on gold and returns on stock price indices in the United States. Correlations with the gold price set in the London afternoon fixing being largely unchanged. Correlation coefficients between returns on the London closing price and those on the previous day's US stock closing prices are closer to zero and insignificant—not surprising since there has now been 2½ hours current day's trading in the US.

Table 4D reports contemporaneous correlation coefficients for daily returns with the sample split into two equal subperiods, periods 1 and 2. The principal findings are that correlation coefficients involving returns on the London afternoon fixing and closing prices, and also the Handy and Harmon price, are negative and significant in period 1. All correlation coefficients are insignificantly different from zero in period 2. The hypothesis of equal correlation coefficients for the two subperiods is rejected in all except one case for returns on the London afternoon fixing and closing prices. However, this evidence of significantly different correlation coefficients for the two subperiods does not carry over to weekly and monthly returns reported in Tables 4W and 4M respectively. In all of these cases, the hypothesis of equal correlation coefficients is not rejected. Also, for monthly returns almost all of the correlations are insignificantly different from zero, providing further support for the evidence for the entire period which is reported in Table 2M.

Figure 1 illustrates results for correlation coefficients for daily returns on the gold price set in the London afternoon fixing and the Dow Jones Industrial Average. Four results are illustrated. The set of eleven correlation coefficients calculated for each successive year from 1991 to 2001—'Annual'—is included together with their

95% confidence interval. Only two of these contemporaneous correlation coefficients—those for 1991 and 1995—are significantly different from zero at the 0.05 level. The correlation coefficients for the Entire Period, –0.0525 from Table 2D, Period 1, –0.1177 from Table 4D, and Period 2, –0.0244 from Table 4D, are also shown.

In summary, the weight of the evidence for the last decade is that the short-run contemporaneous correlation between returns on gold and US stock price indices is small and negative and sometimes insignificantly different from zero. This result holds irrespective of the particular gold price or US stock price index employed.[2] A typical result is illustrated in Figure 2 which shows the scatter of observations of daily gold returns (based on the London afternoon fixing) and returns on the Dow Jones Industrial Average . The sample correlation coefficient, reported in Table 2D, is –0.0525. Clearly, any short-run association between daily returns on gold and a stock price index is weak.

If the logarithms of gold prices and a stock price indices are generated by difference stationary processes then they may be cointegrated. Phillips-Perron unit root tests are used to examine orders of integration. Preliminary results, not reported here, test the hypothesis that the logarithm of each series is I(2) against the alternative that it is I(1) and the null hypothesis is always rejected. Tests that each series is I(1) against the alternative of I(0) are reported for daily data in Table 5D. Tests are carried out sequentially beginning with the general model

$$g_t = \tilde{\mu} + \tilde{\beta}\left(t - \frac{T}{2}\right) + \tilde{\alpha} g_{t-1} + \tilde{\epsilon}_t \tag{13}$$

[2] At least when blue-chip, composite/all-share indices are used. Given the myriad of subsector indices generated for the US, there remains the possibility that some obscure index might generate a contrary result. However, such index-specific evidence is of little interest.

for which $Z(\tilde{\alpha})$, $Z(t_{\tilde{\alpha}})$ are the normalized coefficient and t-tests for $H_0: \tilde{\alpha} = 1$, $Z(t_{\tilde{\beta}})$ is the t–statistic for the test $H_0: \tilde{\beta} = 0$, $Z(t_{\tilde{\mu}})$ is the t-test for $H_0: \tilde{\mu} = 0$ and $Z(\Phi_3)$, $Z(\Phi_2)$ are the regression F-tests for $H_0: \tilde{\alpha} = 1 , \tilde{\beta} = 0$ and $H_0: \tilde{\alpha} = 1 , \tilde{\beta} = 0 , \tilde{\mu} = 0$ respectively. When $\tilde{\beta} = 0$, we have the regression equation

$$g_t = \mu^* + \alpha^* g_{t-1} + \epsilon_t^* \tag{14}$$

Here, $Z(\alpha^*)$, $Z(t_{\alpha^*})$ test $H_0: \alpha^* = 1$, $Z(t_{\mu^*})$ is the t-test for $H_0: \mu^* = 0$ and $Z(\Phi_1)$ is the regression F-test of $H_0: \alpha^* = 1 , \mu^* = 0$. Perron (1988) notes that the test statistics based on this model are not invariant with respect to the drift parameter and are appropriate only if $Z(\Phi_2)$ does not reject $H_0: \tilde{\alpha} = 1 , \tilde{\beta} = 0 , \tilde{\mu} = 0$. If a series has zero mean then the following restricted regression may be more appropriate

$$g_t = \hat{\alpha} g_{t-1} + \hat{\epsilon}_t \tag{15}$$

for which $Z(\hat{\alpha})$, $Z(t_{\hat{\alpha}})$ test $H_0: \hat{\alpha} = 1$. Not one of the test statistics in Table 5D is significant and so the evidence from unit root tests can be summarised concisely: all of the series are I(1).[3]

Is there long-run equilibrium between gold prices and US stock price indices? Cointegration tests address this question by focusing on whether a linear combination of nonstationary series is stationary. A variety of tests might be employed but since pairwise tests of cointegration between a single gold price and a single US stock price index a relevant, Engle-Granger tests are most suitable for present purposes and are reported in Table 6D. These statistics are obtained from sets of cointegrating regression augmented Dickey-Fuller tests generated by including from zero to twelve

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[3] The same inferences are drawn from tests on weekly and monthly data which are not reported here.

lagged dependent variables. Individual test statistics were selected using the Schwarz Bayesian Criterion. Not one test rejects the non-cointegration null hypothesis and so there is no long-run equilibrium between gold prices and US stock price indices.[4] This result is not surprising. Figure 3 illustrates the logarithms of the Dow Jones Industrial Average and the gold price set in the London afternoon fixing, both re-based.[5] The difference between the two is clearly nonstationary. By way of contrast, when series are cointegrated, they move together in the long-run. Figure 4 illustrates such a case for the S&P 500 and Wilshire 5000 stock price indices (both re-based).

Since no gold price and US stock price index are cointegrated, regressions (10) and (11) but without the lagged disequilibrium errors, are used to implement Granger causality tests. Results using daily data are reported in Table 7D. This table is partitioned into two. Panel A provides results for tests of the hypothesis that lagged stock returns do not Granger-cause gold returns. Panel B reports results of tests that lagged gold returns do not Granger-cause returns on US stocks. The tests are implemented by calculating the *F* statistic for the test of the null hypothesis that all of the coefficients on lagged stock (gold) returns in an equation explaining returns on gold (US stocks) are zero. These results were obtained with the number of lagged variables under consideration, $n = 10$. Similar results were obtained for $n = 5$. Table 7D reports the *F* statistic and the sum of the coefficients on lagged stock (gold) returns in a regression explaining gold (stock) returns. Consider the first column of Panel A. The hypothesis tested is: lagged US stock returns do not Granger-cause

4 Further tests using data of different frequencies are not reported since it is data span, rather than frequency of observation, which is important in determining the power of cointegration tests (Otero and Smith, 2000).

5 Each series is multiplied by a constant so that the initial observation has a value of 100.00 and then natural logarithms are taken.

returns on the gold price set at the London morning fixing. This hypothesis is rejected irrespective of the particular US equity index used and in every case the sum of the coefficients on lagged variables is negative. The first row in Panel B reports results for the reverse causality: the hypothesis under consideration is lagged returns on the gold price set in the London morning fixing do not Granger-cause US stock returns. For every index except the Dow Jones Industrial Average, this hypothesis is not rejected. That is, for the remaining set of five indices there is unidirectional causality from US stock returns to returns on the London morning gold price. Similar results are found for the results obtained from the Handy and Harmon series and the London closing price. For the gold price set in the London afternoon fixing—half-an-hour after the New York market opens—the results are different. Here there is clear evidence of bidirectional causality, feedback between gold and stock markets. This feature is also apparent with returns on the Dow Jones Industrial Average and returns on the gold price.

V. CONCLUSIONS

For the period beginning in January 1991 and ending in October 2001 empirical evidence is examined on the relationship between the price of gold and stock price indices for the United States. Three gold prices set in London and one set in New York are used, together with six stock price indices of varying coverage. The short-run correlation between returns on gold and returns on US stock price indices is small and negative and for some series and time periods insignificantly different from zero. All of the gold price and stock price indices are I(1). Over the period examined, there is no cointegration involving a gold price and US stock price index. That is, there is no long-run equilibrium and the series do not share a common stochastic trend. Only short-run relationships are evident. Granger causality tests find evidence of unidirectional causality from US stock returns to returns on the gold price set in the London morning fixing and the closing price. For the price set in the afternoon fixing, there is clear evidence of feedback between the markets for gold and US stocks.

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Figure 1

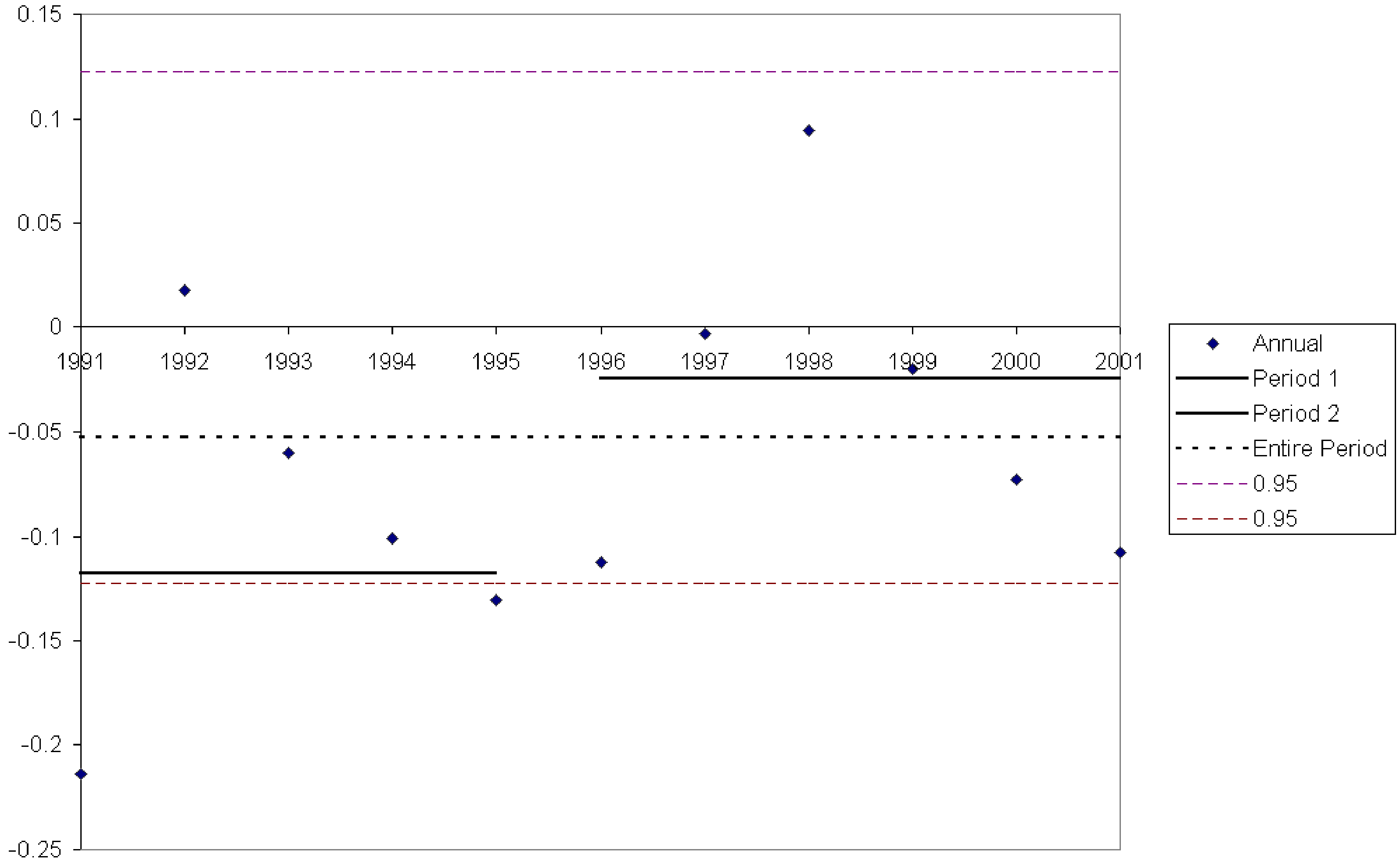

0.15
0.1
0.05
0
-0.05
-0.1
-0.15
-0.2
-0.25
1991
1992
1993
1994
1995
1996
1997
1998
1999
2000
2001
Annual
Period 1
Period 2
Entire Period
0.95
0.95

Figure 2 Scatter Plot of Daily Gold Returns on Dow Jones Returns

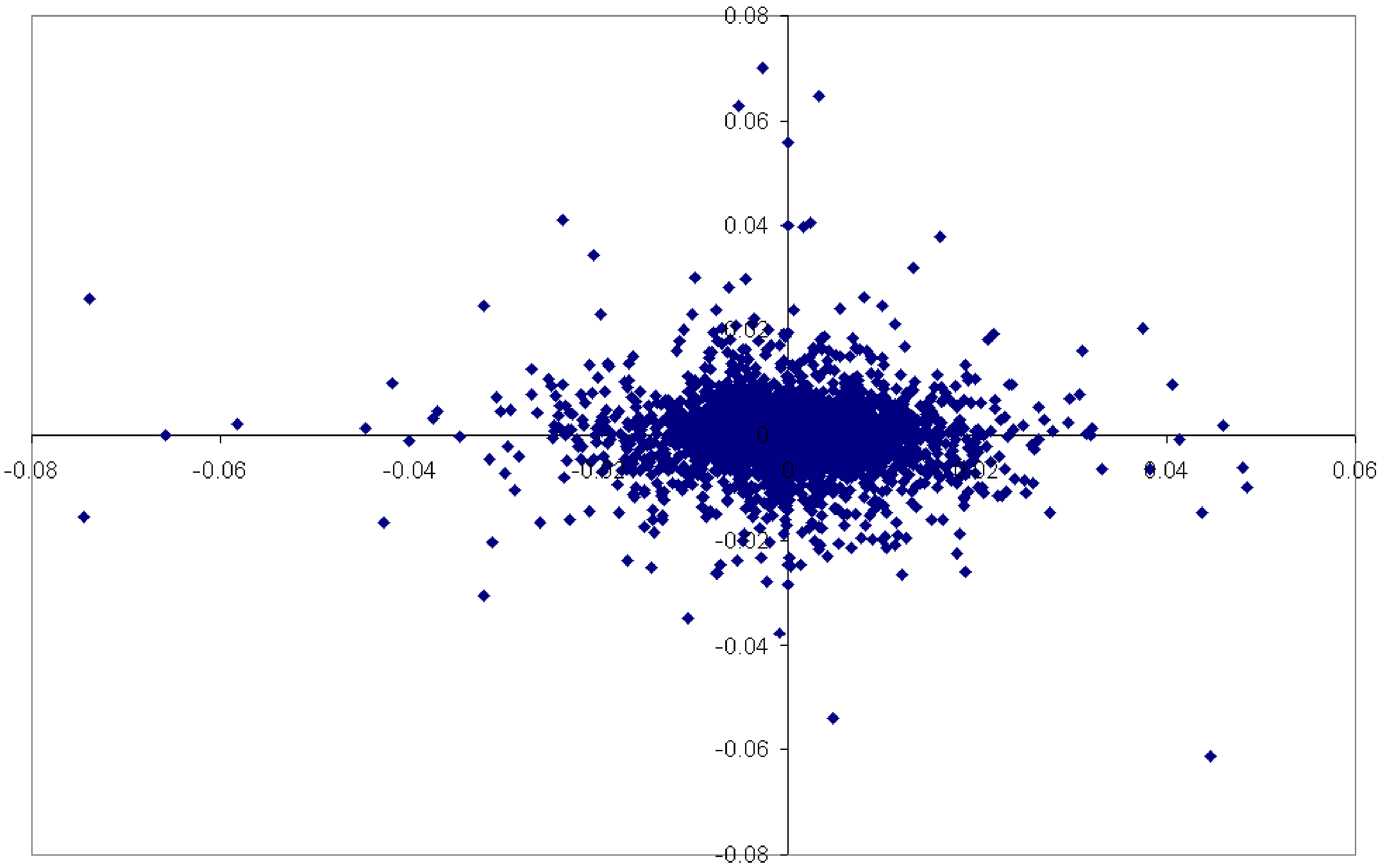


Figure 3 London Gold Price and Dow Jones Index

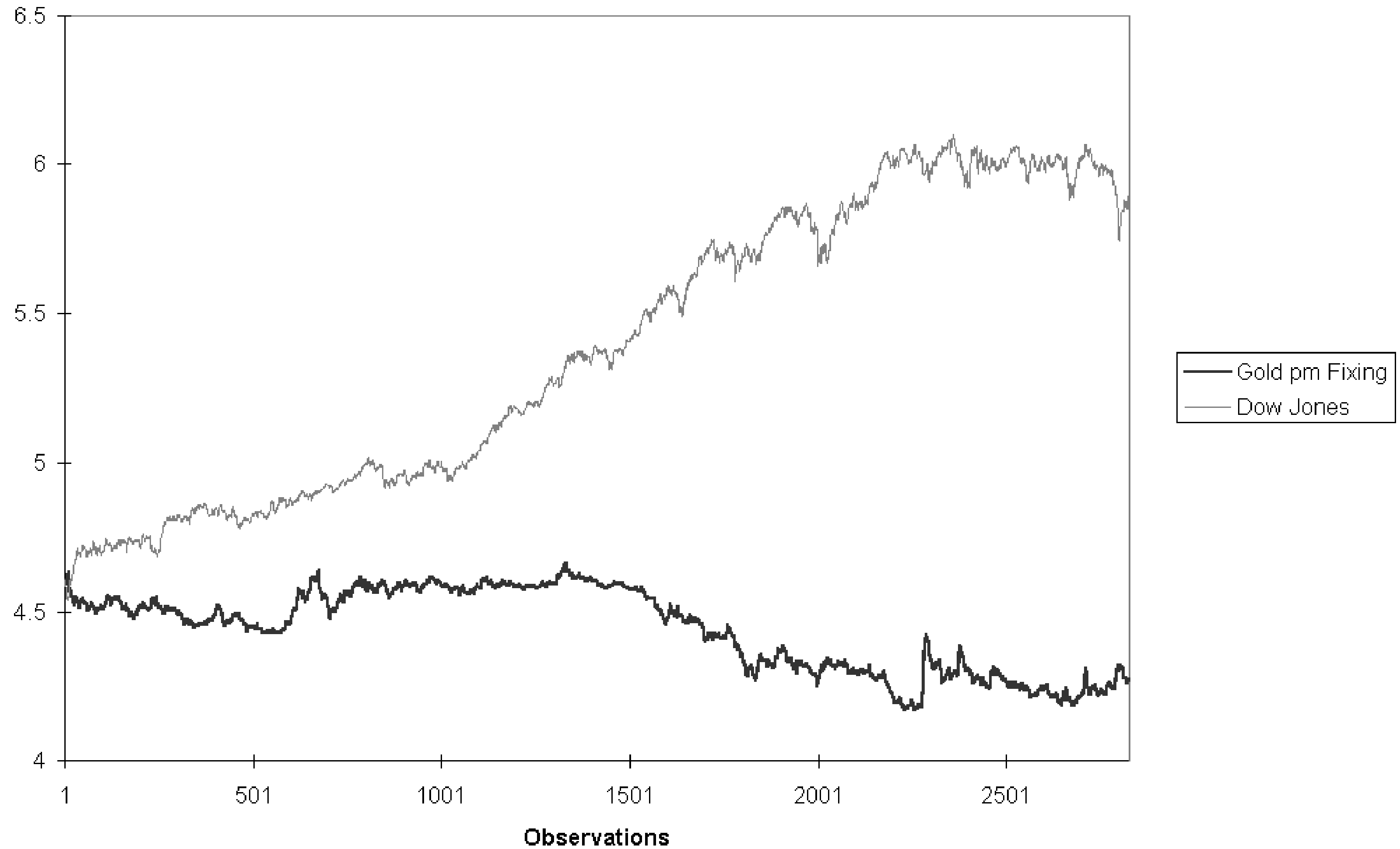


Figure 4 Cointegrated Series

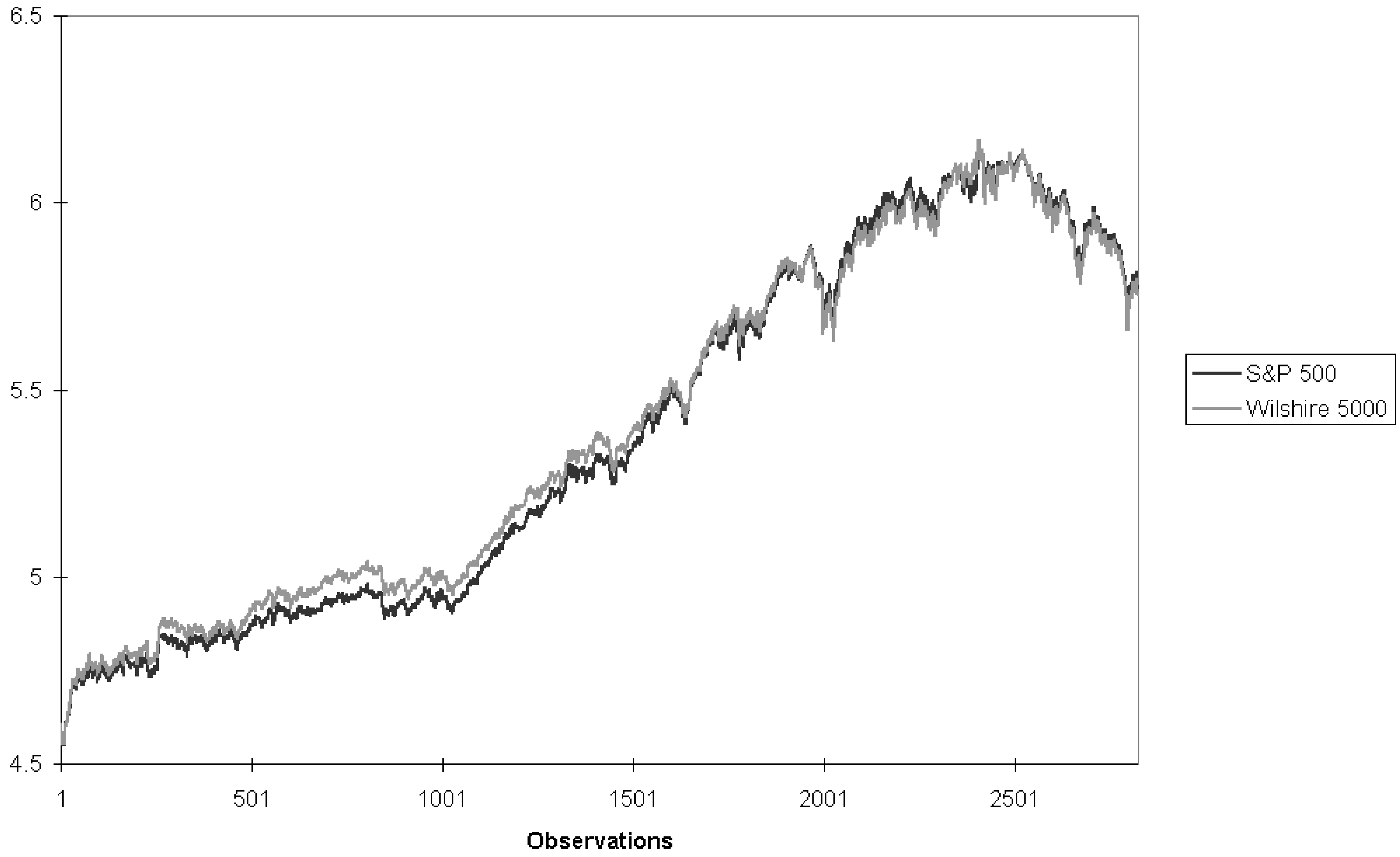


Table 1D Summary Statistics for Daily Returns

	AM Fixing	PM Fixing	Closing	Handy & Harmon	Dow Jones	Nasdaq	NYSE Composite	S&P 500 Composite	Russell 3000	Wilshire 5000
Mean	–0.0001	–0.0001	–0.0001	–0.0001	0.0004	0.0005	0.0004	0.0004	0.0004	0.0004
Median	0.0000	0.0000	0.0000	0.0000	0.0003	0.0009	0.0002	0.0001	0.0002	0.0003
Standard Deviation	0.0078	0.0074	0.0071	0.0077	0.0096	0.0158	0.0084	0.0097	0.0095	0.0095
Skewness	1.6190	0.7150	0.5535	0.3259	–0.5018	–0.0927	–0.4373	–0.2668	–0.3844	–0.3945
Excess Kurtosis	21.9916	13.1918	16.3580	28.5520	5.9211	6.9856	5.7341	4.987	5.1998	5.5293
Jarque-Bera	58,161.6	20,724.8	31,641.1	96,007.7	4,245.3	5,478.0	3,960.3	2,961.3	3,252.1	3,672.0

Notes: The standard errors of the coefficients of skewness and excess kurtosis under the hypothesis of normality are 0.046 and 0.092 respectively. The .05 critical value for the Jarque-Bera test is 5.99.

Table 2D Correlation Coefficients for Daily Returns, January 1991 – October 2001

	Dow Jones	Nasdaq	NYSE Composite	S&P 500	Russell 3000	Wilshire 5000
AM Fixing	–0.0129	0.0057	–0.0125	–0.0070	–0.0053	–0.0028
PM Fixing	–0.0525*	–0.0125	–0.0459*	–0.0411*	–0.0414*	–0.0369*
Closing	–0.0542*	–0.0195	–0.0539*	–0.0517*	–0.0516*	–0.0458*
Handy & Harmon	–0.0502*	–0.0215	–0.0457*	–0.0425*	–0.0413*	–0.0410*

Notes: The asymptotic standard errors are 0.0188.

*Significantly different from zero at the .05 level.

Table 2W Correlation Coefficients for Weekly Returns, January 1991 – October 2001

	Dow Jones	Nasdaq	NYSE Composite	S&P 500	Russell 3000	Wilshire 5000
AM Fixing	–0.1041*	–0.0471	–0.1032*	–0.1001*	–0.0903*	–0.0873*
PM Fixing	–0.0998*	–0.0492	–0.0932*	–0.0939*	–0.0856*	–0.0838*
Closing	–0.0772	–0.0353	–0.0756	–0.0764	–0.0685	–0.0666
Handy & Harmon	–0.1111*	–0.0595	–0.1040*	–0.1046*	–0.0967*	–0.0950*

Notes: The asymptotic standard errors are 0.0422.
*Significantly different from zero at the .05 level.

Table 2M Correlation Coefficients for Monthly Returns, January 1991 – October 2001

	Dow Jones	Nasdaq	NYSE Composite	S&P 500	Russell 3000	Wilshire 5000
AM Fixing	–0.0771	0.0220	–0.0962	–0.0949	–0.0692	–0.0491
PM Fixing	–0.1013	–0.0031	–0.1154	–0.1180	–0.0922	–0.0706
Closing	–0.0979	0.0045	–0.1118	–0.1084	–0.0851	–0.0648
Handy & Harmon	–0.1051	–0.0016	–0.1242	–0.1176	–0.0962	–0.0759

Notes: The asymptotic standard errors are 0.0887.
*Significantly different from zero at the .05 level.

Table 3D Correlation Coefficients for Daily Returns Using Previous Day's US Closing Prices, January 1991 – October 2001

	Dow Jones	Nasdaq	NYSE Composite	S&P 500	Russell 3000	Wilshire 5000
AM Fixing	–0.0851*	–0.0477*	–0.0807*	–0.0848*	–0.0825*	–0.0758*
PM Fixing	–0.0419*	–0.0218	–0.0386*	–0.0440*	–0.0420*	–0.0352
Closing	–0.0228	–0.0025	–0.0136	–0.0148	–0.0137	–0.0084

Notes: The asymptotic standard errors are 0.0188.
*Significantly different from zero at the .05 level.

Table 4D Correlation Coefficients for Daily Returns

		Dow Jones	Nasdaq	NYSE Composite	S&P 500	Russell 3000	Wilshire 5000
AM Fixing	Period 1	–0.0460	–0.0322	–0.0499	–0.0494	–0.0531*	–0.0573*
	Period 2	0.0002	0.0159	0.0019	0.0085	0.0117	0.0159
PM Fixing	Period 1	–0.1177*§	–0.0668*	–0.1246*§	–0.1229*§	–0.1260*§	–0.1279*§
	Period 2	–0.0244§	0.0027	–0.0127§	–0.0088§	–0.0090§	–0.0029§
Closing	Period 1	–0.1519*§	–0.0812*§	–0.1513*§	–0.1556*§	–0.1573*§	–0.1550*§
	Period 2	–0.0112§	–0.0021§	–0.0120§	–0.0099§	–0.0104§	–0.0047§
Handy & Harmon	Period 1	–0.0882*	–0.0532*	–0.0920*	–0.0902*	–0.0942*§	–0.1012*§
	Period 2	–0.0321	–0.0126	–0.0239	–0.0216	–0.0187§	–0.0162§

Notes: Period 1 begins on 2 January 1991 and ends on 31 May 1996 (1413 observations); period 2 starts on 3 June 1996 and ends on 31 October 2001 (1412 observations).
The asymptotic standard errors are 0.0266.
*Significantly different from zero at the .05 level.
§Correlation coefficients for periods 1 and 2 are significantly different at the .05 level.

Table 4W Correlation Coefficients for Weekly Returns

		Dow Jones	Nasdaq	NYSE Composite	S&P 500	Russell 3000	Wilshire 5000
AM Fixing	Period 1	–0.1121	–0.0790	–0.1335*	–0.1342*	–0.1327*	–0.1273*
	Period 2	–0.1019	–0.0392	–0.0907	–0.0880	–0.0739	–0.0725
PM Fixing	Period 1	–0.0938	–0.0634	–0.1149	–0.1174*	–0.1165*	–0.1090
	Period 2	–0.1049	–0.0475	–0.0841	–0.0847	–0.0733	–0.0747
Closing	Period 1	–0.0831	–0.0529	–0.1128	–0.1157	–0.1129	–0.1041
	Period 2	–0.0760	–0.0320	–0.0592	–0.0603	–0.0504	–0.0522
Handy & Harmon	Period 1	–0.1123	–0.0790	–0.1302*	–0.1325*	–0.1316*	–0.1247*
	Period 2	–0.1126	–0.0562	–0.0925	–0.0935	–0.0826	–0.0839

Notes: Period 1 begins on 9 January 1991 and ends on 29 May 1996 (282 observations); period 2 starts on 5 June 1996 and ends on 31 October 2001 (283 observations).
The asymptotic standard errors are 0.0598.
*Significantly different from zero at the .05 level.
The hypothesis of equal correlation coefficients is not rejected.

Table 4M Correlation Coefficients for Monthly Returns

		Dow Jones	Nasdaq	NYSE Composite	S&P 500	Russell 3000	Wilshire 5000
AM Fixing	Period 1	0.0026	–0.1910	–0.1476	–0.1609	–0.1762	–0.1687
	Period 2	–0.1202	0.0781	–0.0803	–0.0744	–0.0333	–0.0089
PM Fixing	Period 1	–0.0613	–0.2991*	–0.2083	–0.2226	–0.2344	–0.2225
	Period 2	–0.1260	0.0520	–0.0825	–0.0834	–0.0447	–0.0204
Closing	Period 1	–0.0623	–0.2321	–0.2088	–0.2233	–0.2354	–0.2268
	Period 2	–0.1220	0.0666	–0.0754	–0.0677	–0.0316	–0.0077
Handy & Harmon	Period 1	–0.0775	–0.2439	–0.2182	–0.2313	–0.2454*	–0.2364
	Period 2	–0.1264	0.0621	–0.0891	–0.0774	–0.0431	–0.0193

Notes: Period 1 begins in January 1991 and ends in May 1996 (65 observations); period 2 starts in June 1996 and ends in October 2001 (65 observations).
The asymptotic standard errors are 0.1270.
*Significantly different from zero at the .05 level.
The hypothesis of equal correlation coefficients is not rejected.

Table 5D Phillips-Perron Unit Root Tests

Series	$Z(\tilde{\alpha})$	$Z(t_{\tilde{\alpha}})$	$Z(t_{\tilde{\beta}})$	$Z(t_{\tilde{\mu}})$	$Z(\Phi_3)$	$Z(\Phi_2)$	$Z(\alpha^*)$	$Z(t_{\alpha^*})$	$Z(t_{\mu^*})$	$Z(\Phi_1)$	$Z(\hat{\alpha})$	$Z(t_{\hat{\alpha}})$	
AM Fixing	−10.55	−2.37	−1.70	2.21	2.47	1.87	−4.11	−1.43	1.41	1.36	−0.06	−0.86	I(1)
PM Fixing	−9.54	−2.20	−1.69	2.17	2.37	1.83	−3.86	−1.37	1.35	1.32	−0.06	−0.90	I(1)
Closing	−8.73	−2.02	−1.68	2.15	2.34	1.81	−3.82	−1.37	1.34	1.31	−0.06	−0.91	I(1)
Handy & Harmon	−10.21	−2.33	−1.67	2.18	2.39	1.84	−4.02	−1.42	1.40	1.37	−0.06	−0.89	I(1)
Zurich	−19.55	−4.06	−1.86	2.43	1.59	1.23	−4.58	−1.47	1.46	1.34	−0.03	−0.75	I(1)
Dow Jones	−5.81	−1.24	0.90	1.20	1.29	2.89	−1.38	−1.33	1.47	3.95	0.14	2.40	I(1)
NASDAQ	−2.64	−0.65	0.07	0.73	1.72	2.22	−2.47	−1.86	2.01	3.33	0.19	1.62	I(1)
NYSE Composite	−4.76	−1.07	0.80	1.17	1.40	2.92	−1.57	−1.47	1.64	4.08	0.18	2.34	I(1)
S&P 500	−3.16	−0.81	0.31	0.71	1.02	2.51	−1.44	−1.40	1.57	3.74	0.17	2.24	I(1)
Russell 3000	−3.60	−0.83	0.44	0.86	1.31	2.69	−1.68	−1.56	1.74	3.95	0.18	2.21	I(1)
Wilshire 5000	−3.16	−0.73	0.38	0.80	1.33	2.54	−1.74	−1.59	1.70	3.83	0.13	2.18	I(1)
.05 crit. val	–21.80	–3.41	±3.11	±3.38	6.25	4.68	–14.10	–2.86	±2.83	4.59	–8.10	–1.95	

Table 6D Engle-Granger Cointegration Tests, January 1991 – October 2001

	Dow Jones	Nasdaq	NYSE Composite	S&P 500	Russell 3000	Wilshire 5000
AM Fixing	–2.2373	–2.1021	–2.3883	–2.3901	–2.3454	–2.3171
PM Fixing	–2.2242	–2.2138	–2.3833	–2.3769	–2.3378	–2.4465
Closing	–2.2627	–2.1114	–2.4154	–2.4188	–2.3678	–2.3375
Handy & Harmon	–2.3182	–2.1687	–2.4725	–2.4841	–2.4303	–2.3966

Note: The .95 critical value is –3.3398.

Table 7D Granger Causality Tests

Panel A

	Dependent Variable			
	AM	PM	CL	HH
DJ	4.6117*	2.2690*	2.0621*	2.1458*
$\sum\gamma_{1j}$	−0.0950	–0.0509	−0.0756	−0.0641
NS	2.0303*	2.1066*	1.5094	2.0952*
$\sum\gamma_{1j}$	−0.0236	−0.0006	−0.0125	−0.0149
NY	4.4966*	2.4608*	1.9748*	2.4609*
$\sum\gamma_{1j}$	−0.1006	−0.0646	−0.0724	−0.0646
SP	4.5106*	2.7392*	2.2630*	2.5912*
$\sum\gamma_{1j}$	−0.0871	−0.0577	−0.0591	−0.0566
R3	4.2079*	2.6292*	1.8814*	2.5707*
$\sum\gamma_{1j}$	−0.0809	−0.0523	−0.0525	−0.0522
W5	4.0427*	2.6844*	2.1745*	2.5725*
$\sum\gamma_{1j}$	−0.0764	−0.0492	−0.0492	−0.0484

Panel B

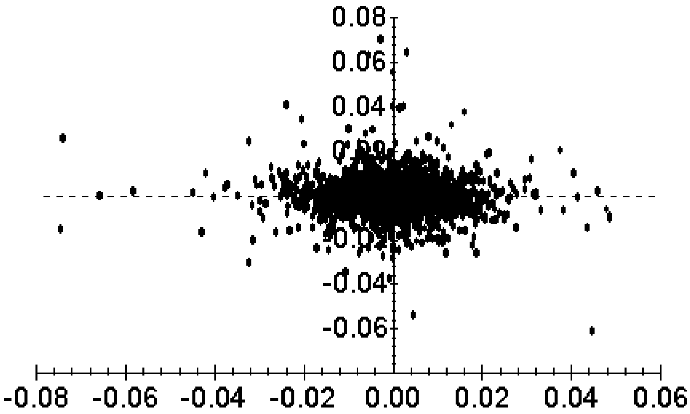
	Dependent Variable					
	DJ	NS	NY	SP	R3	W5
AM	1.9092*	0.6899	1.3378	1.0530	1.0052	0.8214
$\sum\gamma_{2j}$	−0.0487	−0.1109	−0.0760	−0.0821	−0.0832	−0.0821
PM	3.2735*	1.8586*	2.1141*	2.1629*	2.1650*	2.0151*
$\sum\gamma_{2j}$	−0.0326	−0.0947	−0.0656	−0.0674	−0.0685	−0.1080
CL	2.3670*	0.8499	1.4678	1.3019	1.0753	0.9573
$\sum\gamma_{2j}$	−0.0210	−0.0900	−0.0523	−0.0550	−0.0551	−0.0561
HH	1.9177*	1.3446	1.4039	1.3005	1.3981	1.2434
$\sum\gamma_{2j}$	−0.0009	−0.0640	−0.0408	−0.0393	−0.0448	−0.0402

Note: * Significant at the .05 level; the critical value is 1.83.

APPENDIX

Figure 5 illustrates again the scatter of observations of daily gold returns (based on the London afternoon fixing) and returns on the Dow Jones Industrial Average.

FIGURE 5 **SCATTER PLOT OF DAILY GOLD RETURNS ON DOW JONES RETURNS**



Any association between daily returns on gold and a stock market price index is weak. The sample correlation coefficient is –0.0525 (Table 2D, p 20). However, this can be viewed within a linear regression framework. Kendall and Stewart (1961, p 296) note that a test of the hypothesis that the population correlation coefficient is zero is equivalent to a test of the hypothesis that the slope coefficient in a simple linear regression is zero. Consider the regression of gold returns on an intercept and returns on the Dow Jones stock price index and the reverse regression of Dow Jones returns on gold returns. The following results are obtained.

$$\Delta\hat{g}_t = -0.0001 - 0.0407\Delta s_t$$
$$(0.7317) \qquad (2.7946) \qquad\qquad |\,t\,|$$

$$\Delta\hat{s}_t = 0.0004 - 0.0679\Delta g_t$$
$$(2.3893) \qquad (2.7946)$$

In both regressions: the t statistic for the test of the hypothesis that the slope coefficient is zero is –2.7946, the F statistic is 7.8099 and R^2 is 0.0027589.[6] The following relationships between test statistics are evident

$$r = -0.0525 \text{ and } \pm\sqrt{R^2} = \pm 0.0525$$

$$F_{(1,2823)} = 7.8099 = (-2.7946)^2 = t_{2823}.$$

However, both regressions have non-spherical disturbances.

Estimation of the basic GARCH(0, 1) model

$$\Delta g_t = \beta_0 + \beta_1 \Delta s_t + u_t$$

where $u_t | \Psi_{t-1} \sim \mathrm{N}(0, h_t)$

and variance of the error term is positively related to the size of previous errors, that is, it depends on past volatilities,

$$h_t = \alpha_0 + \alpha_1 u_{t-1}^2$$

gives

$$\underset{}{\Delta \hat{g}_t} = \underset{(0.0001)}{-0.0001} \; - \underset{(0.0139)}{0.0285} \Delta s_t \qquad (\textit{Asymp SE})$$

with

$$h_t = \underset{(0.0000)}{-0.0000} \; - \underset{(0.0312)}{0.3023} u_{t-1}^2 ,$$

$F_{(2,2821)} = 9.9603$ and $R^2 = 0.007012$.

The (negative) $\sqrt{R^2} = -0.0501$, which is similar to previous estimates.

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6 $t_{T-2} \equiv \sqrt{F_{(1,T-2)}}$