# Are stock prices and economic activity cointegrated? Evidence from the United States, 1950-2005

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**Abstract**

The potential cointegrating relationship between stock prices and economic activity suggested by financial and economic theory is examined. It is found that the commonly employed tests of Engle and Granger (1987) and Johansen (1988) fail to detect cointegration between stock prices and industrial production for a long span of U.S. data. In recognition of factors which may result in a failure to detect a genuine cointegrating relationship, the analysis is extended to consider higher-powered cointegration tests, tests which allow for structural change in the cointegrating relationship and tests of asymmetric cointegration. However, despite considering a range of tests, no evidence of cointegration is detected. The results therefore do not support the predictions of financial and economic theory.

**Implications For Practice**

Potential cointegration between stock prices and economic activity in the U.S. is examined over a longer span of data than considered previously in the literature. The noted absence of cointegration has a number of implications for practitioners. First, the failure to detect a long-run relationship should be recognised and incorporated by modellers who wish to formulate models of stock price behaviour. In particular modellers should be careful to avoid potential problems of avoid spurious regression between the series. Second, the failure to detect cointegration may be due to the greater volatility exhibited by the stock price data. This is a further issue to consider when examining stock price behaviour.

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# 1 Introduction

In a recent years, a large literature has emerged analysing the relationship between financial and macroeconomic variables. While some authors have examined the link between stock market returns and differenced macroeconomic time series (see, *inter alia*, Pesaran and Timmerman 1995, 2000; Henry *et al.* 2004), others have examined potential long-run relationships between stock market and macroeconomic variables (see, *inter alia*, Nasseh and Strauss 2000; McMillan 2005). As noted by McMillan (2005), the latter approach is of particular interest as cointegration between stock prices and economic activity is consistent with theoretical research in economics and finance. In this paper the hypothesis of a cointegrating relationship between economic activity and stock prices is examined using monthly U.S. data on industrial production and stock prices over the period 1950 to 2005. The empirical analysis undertaken complements the recent research of McMillan (2005) in two ways. First, in comparison to the 1970-2000 sample considered by McMillan (2005), a longer span of U.S. data is employed. This issue is of importance as following Hakkio and Rush (1991) it has become recognised that the span of the data considered is of crucial importance for the power of cointegration tests. Second, a range of tests is applied to detect potential cointegration between industrial production and stock prices. Therefore, in addition to the familiar Engle-Granger (1987) and Johansen (1988) tests, further tests of cointegration are applied to extend the analysis in three directions to allow for the possibility that these tests may fail to uncover an underlying cointegrating relationship for alternative reasons. Initially the higher-powered cointegration tests of Rodriguez and Perron (2001) and Kanioura and Turner (2005) are employed to allow for the possibility that commonly considered tests fail to detect cointegration due to a lack of power. Second, the cointegration tests of Gregory and Hansen (1996) which explicitly extend the Engle-Granger test to allow for potential structural change in the cointegrating relationship are applied. These tests permit examination of the hypothesis that failure to detect cointegration using standard tests may result from time variation in an underlying cointegrating relationship. Finally, the threshold autoregressive and momentum-threshold autoregressive based cointegration tests of

Enders and Siklos (2001) are applied to consider the hypothesis that the cointegrating relationship between industrial production and stock prices is of an asymmetric form, with differing speeds of adjustment back to an underlying equilibrium relationship present. The results of Enders and Siklos (2001) show that standard tests of cointegration exhibit low power in the presence of asymmetric cointegration and consequently failure to detect cointegration using standard tests may be due to presence of asymmetric behaviour. The results of the present analysis show that the null of no cointegration is not rejected by any of the tests employed. The absence of cointegration between stock prices and economic activity over the longer span of data considered herein has a number of implications for practitioners. First, the failure to detect a long-run relationship should be recognised and incorporated by modellers who wish to formulate models of stock price behaviour. In particular modellers should be careful to avoid potential problems of avoid spurious regression between the series. Second, the failure to detect cointegration may be due to the greater volatility exhibited by the stock price data. This is a further issue to consider when examining stock price behaviour.

This paper proceeds as follows. In the following section the alternative cointegration tests to be employed are presented. Section [3] discusses the data examined, with the empirical results obtained from the application of the differing tests of cointegration provided also. Section [4] concludes.

# 2 Alternative tests for cointegration

## 2.1 The Engle-Granger test

The Engle-Granger (1987) (EG) test is the most commonly employed (single equation) approach to the analysis of cointegration in the econometrics literature. Given two variables of interest $\{y_t, x_t\}$, the first stage of this two-step procedure involves the estimation of the following static cointegrating regression:

$$y_t = d_t + \beta x_t + \epsilon_t \qquad t = 1, ..., T \tag{1}$$

where $d_t$ denotes a deterministic term which may be either an intercept ($d_t = \alpha$) or an intercept and linear trend ($d_t = \alpha + \beta t$). In the second step, potential cointegration between $\{y_t, x_t\}$ is examined via analysis of the order of integration of the residuals $\{\widehat{\epsilon}_t\}$ from (1) using a Dickey-Fuller (1979) test as below:

$$\Delta\widehat{\epsilon}_t = (\rho - 1)\,\widehat{\epsilon}_{t-1} + \nu_t \tag{2}$$

The null hypothesis of no cointegration is examined via the $t$-like statistic for $(\rho - 1)$. In empirical analysis, equation (2) is augmented as necessary via the inclusion of lagged values of the dependent variable.

## 2.2 The Johansen test

The Johansen (1988) approach to the analysis of cointegration is based upon the vector error correction (VECM) model below:

$$\Delta\mathbf{Z}_t = \sum_{i=1}^{k-1} \mathbf{\Gamma}_i \Delta\mathbf{Z}_{t-i} + \mathbf{\Pi}\mathbf{Z}_{t-k} + \mathbf{v}_t \tag{3}$$

where $\mathbf{Z}_t$ is a vector containing the I(1) variables of interest and $\mathbf{\Gamma}_i$ and $\mathbf{\Pi}$ are coefficient matrices. The extent of cointegration between the variables is examined via the eigenvalues of the long-run coefficient matrix $\mathbf{\Pi}$ using maximal and trace eigenvalue test statistics. In empirical applications of the Johansen procedure, decisions are required concerning which deterministic terms to include in the VAR and the cointegrating term of (3).

## 2.3 The GLS-based cointegration test

In response to the observed low power of the EG test, Perron and Rodriguez (2001) have suggested the use of local-to-unity GLS detrending as a means of increasing test power. Given two series of interest $\{y_t, x_t\}$, GLS-transformed data are created as follows:

$$\begin{aligned} y_{\overline{\alpha}} &= [y_1, y_2 - \overline{\alpha}y_1, ..., y_T - \overline{\alpha}y_{T-1}]' \\ x_{\overline{\alpha}} &= [x_1, x_2 - \overline{\alpha}x_1, ..., x_T - \overline{\alpha}x_{T-1}]' \\ z_{\overline{\alpha}} &= [z_1, z_2 - \overline{\alpha}z_1, ..., z_T - \overline{\alpha}z_{T-1}]' \end{aligned}$$

where $z_t$ denotes appropriately defined deterministic terms, $\overline{\alpha} = 1 + \overline{c}T^{-1}$ and $\overline{c}$ is a constant determining the extent of local-to-unity detrending. The deterministic term $z_t$ is given as an intercept in the case of variables not displaying a trend, $z_t = 1$, and an intercept and trend for trending variables, $z_t = (1, t)'$. With the deterministic term defined, the GLS detrended series $y_t^\alpha$ is then derived as $y_t^\alpha = y_t - \widehat{\beta}_0$ when $z_t = 1$, and $y_t^\alpha = y_t - \widehat{\beta}_0 - \widehat{\beta}_1 t$ when $z_t = (1, t)'$, with $\left\{\widehat{\beta}_i\right\}$ obtained from the regression of $y_{\overline{\alpha}}$ upon $z_{\overline{\alpha}}$. The GLS detrended version of $x_t$, denoted as $x_t^\alpha$, is derived using an identical approach. With regard to the choice of the differing parameter $\overline{c}$, Perron and Rodriguez (2001) follow Elliott *et al.* (1996) and advocate the use of $\overline{c} = -7$ when $z_t = 1$ and $\overline{c} = -13.5$ when $z_t = (1, t)'$. With the detrended series obtained, the following static regression is performed:

$$y_t^\alpha = \gamma x_t^\alpha + \eta_t \tag{4}$$

Cointegration between $\{y_t, x_t\}$ is then examined via the residuals $\widehat{\eta}_t$ from (4) using a Dickey-Fuller test of the form of (2) above.

### 2.4 The error correction model based F-test for cointegration

The recent research of Kanioura and Turner (2005) provides a higher-powered alternative to the EG test. Given the two series $\{y_t, x_t\}$ , the F-test of Kanioura and Turner (2005) is simply based upon the significance of the (lagged) levels terms in the following error correction model (ECM):

$$\Delta y_t = \alpha_0 + \alpha_1 \Delta x_t + \alpha_2 y_{t-1} + \alpha_3 x_{t-1} + \xi_t \tag{5}$$

The null hypothesis of no cointegration is given by $H_0$: $\alpha_2 = \alpha_3 = 0$. Kanioura and Turner note that unlike the closely related ECM-based $t$-test of Kremers *et al.* (1992), the critical values of their resulting F-test statistic are invariant to the parameters of the specific problem examined. As with previous tests, empirical application of this test may require inclusion of lagged values of the dependent variable to avoid problems of serial correlation.

### 2.5 Cointegration tests with regime shifts

The Monte Carlo analysis of Gregory *et al.* (1996) shows that the power of the EG test is substantially reduced when applied to cointegrated series which experience a change or break in their cointegrating relationship. In response to this, Gregory and Hansen (1996) extend the EG test to explicitly allow for breaks in either the intercept or the intercept and cointegrating coefficient at an unknown time. The above equation (1) for the first stage of the EG testing procedure is therefore revised as below to provide the following three models:

$$\begin{array}{ll} \text{Model C: Level shift} & y_t = \mu_0 + \mu_1 \varphi_t + \alpha x_t + \omega_t \\ \text{Model C/T: Level shift with trend} & y_t = \mu_0 + \mu_1 \varphi_t + \beta t + \alpha x_t + \omega_t \\ \text{Model C/S: Regime shift} & y_t = \mu_0 + \mu_1 \varphi_t + \alpha_1 x_t + \alpha_2 \varphi_t x_t + \omega_t \end{array} \tag{6}$$

Each of the above models therefore permits structural change via the dummy variable $\varphi_t$ which is defined as:

$$\varphi_t = \begin{cases} 1 & \text{if } t > \tau \\ 0 & \text{otherwise} \end{cases}$$

where $\tau$ denotes the point in the sample at which a break occurs. To determine $\tau$, Gregory and Hansen (1996) suggest the use of a grid search procedure, with all values in the central 70% of the sample being considered. For each value of $\tau$, the above models are estimated with the resulting residuals $\{\widehat{\omega}_t\}$ saved and employed in the following Dickey-Fuller testing equation:

$$\Delta\widehat{\omega}_t = (\rho - 1)\,\widehat{\omega}_t + \nu_t \tag{7}$$

which may be augmented as required by the addition of lagged values of $\Delta\widehat{\omega}_t$. The resulting test statistic for each model is then given as the minimum value obtained for the $t$-ratio of $(\widehat{\rho - 1})$.

### 2.6 Cointegration with asymmetric adjustment

Inspection of equation (2) shows that the EG test has an implicit underlying assumption of symmetry, with adjustment to equilibrium given by the single coefficient $(\rho - 1)$. More recently Enders and Siklos (2001) have extended this approach to permit differing speeds of adjustment to occur.[1] Based upon the threshold autoregressive methods of Tong (1983,1990), equation (2) is extended via the use of a Heaviside indicator function $I_t$ as below:

$$\Delta\widehat{\epsilon}_t = I_t\rho_1\widehat{\epsilon}_{t-1} + (1 - I_t)\,\rho_2\widehat{\epsilon}_{t-1} + \xi_t \tag{8}$$

[1] The extension of the EG test to allow for asymmetric behaviour is to be welcomed as a number of theoretical studies have suggested that macroeconomic series may exhibit asymmetry (see, *inter alia*, Ball and Mankiw 1994; Dixit 1992; Gale 1996; Krane 1994).

with appropriate augmentation included as required. Enders and Siklos propose two specifications of the Heaviside indicator function based upon $\{\widehat{\epsilon}_t\}$ and $\{\Delta\widehat{\epsilon}_t\}$ which lead to threshold autoregressive (TAR) and momentum threshold autoregressive (MTAR) cointegration tests. The Heaviside indicator functions proposed by Enders and Siklos (2001) are given as:

$$I_t = \begin{cases} 1 & \text{if } \widehat{\epsilon}_{t-1} \geq \tau \\ 0 & \text{if } \widehat{\epsilon}_{t-1} < \tau \end{cases} \tag{9}$$

$$I_t = \begin{cases} 1 & \text{if } \Delta\widehat{\epsilon}_{t-1} \geq \tau \\ 0 & \text{if } \Delta\widehat{\epsilon}_{t-1} < \tau \end{cases} \tag{10}$$

where $\tau$ is a consistent estimate of the threshold obtained via a grid search procedure. Under TAR adjustment, the residual series $\{\widehat{\epsilon}_t\}$ is reordered as $\widehat{\epsilon}_1^o < \widehat{\epsilon}_2^o < ... < \widehat{\epsilon}_T^o$. The central 70% of observations ($\widehat{\epsilon}_i^o, i = 0.15T, ...., 0.85T$) within this range of values are then considered as potential thresholds. The threshold value delivering the minimum residual sum of squares for equation (8) using the indicator function of (9) is then denoted as the consistent estimate of the threshold ($\tau$). Under MTAR adjustment, a similar approach is followed with the central 70% of observations from the reordered sequence $\{\Delta\widehat{\epsilon}_1^o, \Delta\widehat{\epsilon}_2^o, ..., \Delta\widehat{\epsilon}_T^o\}$ considered as potential thresholds. Again the selected threshold is that value minimising the residual sum of squares of (8) with (10) now employed as the appropriate indicator function. Under both TAR and MTAR adjustment, the null hypothesis of no cointegration is examined via the joint hypothesis $\mathsf{H}_0 : \rho_1 = \rho_2 = 0$. A statistically significance difference in the values of the asymmetric adjustment coefficients $(\rho_1, \rho_2)$ indicates that a long-run relationship exists but that reversion to it occurs at differing speeds depending upon the partitioning resulting from the indicator function employed.

## 3 Empirical results

The stock price index examined in this paper is the Standard and Poors 500 index while the measure of economic activity considered is U.S. industrial production. Both series are monthly observations over the period January 1950 to June 2005. Both series are considered in their natural logarithmic forms. Application of unit root tests to these series resulted in the inference that both are I(1) processes and consequently potential cointegration between them can be examined.[2] The results obtained from application of the cointegration tests outlined above are presented in Tables One and Two. Considering the results in Table One, it is apparent that the EG test fails to reject the null of no cointegration at the 5% level, irrespective of whether an intercept or intercept and trend are included at the first stage of the procedure. In these circumstances it can be seen that the calculated test statistics of $-1.610$ and $-1.894$, for the intercept and trend models respectively, are not sufficiently large in absolute terms to reject the null against 5% critical values of $-2.866$ and $-3.419$.[3] The results for the Johansen procedure using a VAR(12) lead to a similar inference of no cointegration, with the maximal and trace test statistics of 3.220 and 3.615 dwarfed by their 5% critical values of 14.265 and 15.495. The results of these commonly employed cointegration tests therefore lead to the conclusion that stock prices and industrial production are not cointegrated. However, rather than end the analysis at this stage, additional higher-powered cointegration tests can be considered to explore whether this inference can be reversed. Turning to the GLS-based test results, it can be seen that again the null of no cointegration cannot be rejected, the relevant calculated and critical values being $-1.807$ and $-3.35$ using $z_t = (1, t)'$. A similar inference is drawn using the other higher-powered cointegration test, with the F-test of Kanioura and Turner (2005) having calculated and critical values of 0.229 and 5.83 respectively. It therefore appears

[2] In the interests of brevity, the results of unit root testing are not reported here. However, they are available upon request.

[3] The degree of augmentation of the testing equation in the second stage of the EG procedure was determined using Akaike's Information Criterion subject to the absence of serial correlation using an LM test. A similar approach is followed for the other residual-based tests employed and the F-test of Kanioura and Turner (2005). In all instances, a lag length of zero resulted.

that failure of the EG and Johansen tests to detect cointegration between stock prices and industrial production is not due to a lack of power, as higher-powered tests lead to the same inference. To examine whether no cointegration is due to the use of an inappropriate alternative hypothesis, the analysis can be extended to permit the possibilities of structural change in the cointegrating relationship and asymmetric adjustment. However, considering the results of the Gregory-Hansen tests in Table One it can be seen that the no cointegration null hypothesis cannot be rejected under any of the Models (C, C/T, C/S) considered. Turning to the results from application of asymmetric cointegration tests presented in Table Two, the TAR model shows the adjustment parameters $\{\rho_1, \rho_2\}$ to take very different values. Although these values might be thought to be indicative of asymmetric cointegration, comparison of the calculated and critical values for the test again results in non-rejection of the null. Similar findings are obtained for the MTAR-based cointegration test with the calculated statistic substantially smaller than its corresponding critical value.

## 4 Conclusion

In this paper possible cointegration between stock prices and economic activity has been examined in the U.S. over the period 1950 to 2005. Rather than rely upon the results obtained from the commonly employed Engle-Granger and Johansen tests which failed to uncover significant evidence of cointegration, a variety of alternative tests of cointegration was employed. To explore the possibility that failure to detect cointegration arose as a result of a lack of test power, the higher-powered tests of Rodriguez and Perron (2001) and Kanioura and Turner (2005) were applied. Similarly, to examine whether non-rejection of the no cointegration null was due to failure to account for structural change in the cointegrating relationship or asymmetric adjustment, the tests of Gregory and Hansen (1996) and Enders and Siklos (2001) were utilised. However, in no instance did any of the 11 tests applied detect evidence of cointegration. It can therefore be concluded that for the US over the sample period considered, stock prices and economic activity are not cointegrated.

**Table One: Cointegration tests results I**

| Test | Calculated test statistic | 5% critical value |
|---|---|---|
| Engle-Granger | | |
| $d_t = \alpha$ | $-1.610$ | $-2.866$ |
| $d_t = \alpha + \beta t$ | $-1.894$ | $-3.419$ |
| Johansen | | |
| Maximal | 3.220 | 14.265 |
| Trace | 3.615 | 15.495 |
| GLS Engle-Granger | $-1.807$ | $-3.35$ |
| Cointegration F-test | 0.229 | 5.83 |
| Gregory-Hansen | | |
| Model C | $-3.731$ | $-4.61$ |
| Model C/T | $-3.641$ | $-4.99$ |
| Model C/S | $-3.913$ | $-4.95$ |

*Notes*:
The above tabulated figures are calculated test statistics and critical values for the alternative cointegration tests discussed in section [2]. The critical values for the EG, Johansen, EG-GLS, F- and Gregory-Hansen tests are drawn from MacKinnon (1991), MacKinnon *et al.* (1999), Rodriguez and Perron (2001), Kanioura and Turner (2005), and Gregory and Hansen (1996) respectively.

**Table Two: Cointegration tests results II**

| Asymmetric cointegration tests | |
|---|---|
| **TAR** | |
| $\rho_1, \rho_2$ | $-0.015, -0.005$ |
| $\rho_1 = \rho_2 = 0$ | 1.719 |
| 5% critical value | 6.93 |
| **MTAR** | |
| $\rho_1, \rho_2$ | $-0.024, -0.002$ |
| $\rho_1 = \rho_2 = 0$ | 3.203 |
| 5% critical value | 6.62 |

*Notes*:
The above tabulated figures are calculated test statistics and critical values for the TAR and MTAR cointegration tests discussed in section [2]. The critical values are drawn from Enders and Siklos (2001).

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