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Putting the Dividend-Price Ratio Under the Microscope: Evidence from Developed Countries

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Putting the Dividend-Price Ratio Under the Microscope: Evidence from Developed Countries

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Abstract

We analyze the time-series properties of the dividend-price ratios (DPRs) of 11 developed countries since the early 70s. Despite its frequent use in research as a valuation method for stock prices and a determinant of stock returns, previous studies suggest that there is mixed evidence of the time-series properties of DPRs predicted by economic theory. We argue that this mixed result is attributable to the sample size used in previous studies. Here, we have opted to implement the panel data approach (i.e., $N > 2$) to increase the total size of observations rather than relying on the traditional method (i.e., increasing the size of T). In this way, we can increase the total number of observations without increasing the likelihood of structural breaks. For this purpose, we implement the panel unit root test taking account of cross sectional dependence, and obtain clear evidence of stationary DPRs. Thus, we conclude that a significant one-to-one long-run relationship exists between stock prices and dividends for the countries in our samples.

JEL classification: G100, E440, C120, C220

Keywords: Persistence, dividends, stock prices, panel unit root tests, structural breaks

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1 Introduction

Stock dividends are an important factor for investors making financial portfolio decisions. Furthermore, the dividend-price ratio (DPR), which can be obtained by dividing dividends by stock prices, has been frequently considered as one factor containing useful information for predicting stock returns. For example, in the framework of the present value model, Campbell and Shiller (1989) developed the theoretical relationship between the log stock return and the log DPR. Their theoretical model predicts a positive relationship between them; a high DPR must be accompanied by a high stock return and/or low dividend growth.

Campbell and Shiller, furthermore, discuss the time-series properties of the DPR in their model. Their theory indicates that dividends and stock prices are integrated of order one (i.e., $I(1)$) and are co-integrated. Thus, the DPR is stationary (i.e., $I(0)$), which implies that a long-run co-movement exists between dividends and prices. In other words, it is consistent with the behavior of firms increasing dividends along with rises in stock prices.[1]

However, empirical findings with respect to a stationary DPR are quite mixed. Using the conventional unit root tests, Campbell and Shiller (1989) raise evidence in favor of stationarity for the US DPR. In contrast, using the Augmented Dicky-Fuller test (ADF), Balk and Wohar (2002) and Bohl and Siklos (2004) provide evidence of a non-stationary dividend-price ratio. However, Bohl and Siklos (2004) and Madsen and Milas (2005) report some supportive evidence of stationarity in the US ratio by introducing possible nonlinearity in the specification. For the UK data, the ADF test seems sufficient to reject the null (Madsen and Milas 2005).

Such weak support for a stationary DPR could result from two factors. A non-stationary DPR may reflect the lack of power of standard unit root tests to reject the null hypothesis, which may arise from the small number of observations. Alternatively, this result is attributable to changes in corporate financial policy. For example, in the 1990s, US firms started to reduce a practice of paying dividends to shareholders and increased reliance on repurchasing shares (Liang and Sharp, 1999). This trend seems to reflect the companies' tax burden; repurchasing shares is the least costly form of payout. In this connection, Boudoukh, Michaely, Richardson, and Roberts (2004) consider the total payout ratio (dividends plus repurchases) as an explanatory variable for stock returns, and prove its usefulness compared with the DPR.

Against this background, this paper analyzes once again the time-series properties of the DPR and its components, prices and dividends, in an attempt to elucidate these mixed results.

[1] We also acknowledge that even if the DPR is non-stationary, a test for return predictability can be meaningful as long as the proper econometric method is used (e.g., Campbell and Yogo 2006).

In particular, we are concerned with the first reason for the failure to reject the null hypothesis. This paper deals with the problem of small sample bias by pooling the data of developed countries; namely, Australia, Canada, France, Germany, Italy, Japan, the Netherlands, Sweden, Switzerland, the UK, and the US, and attempts to draw a general conclusion for this set of countries. It enables us to increase the number of observations without increasing the likelihood of structural breaks. This approach has been widely used in research in the area of international finance (e.g., MacDonald 1996, O'Connell 1998), but has been used rarely in financial research. Therefore, the statistical methodology employed in this paper contrasts sharply with ones in previous research focusing solely on country specific data, and thus our study should provide more reliable results than employing the traditional approach given the sample period (i.e., fixed T).[2] Using the recently developed panel unit root test which considers cross sectional dependence in the data, we provide very convincing evidence that generally the DPR is stationary while stock prices and dividends are nonstationary. This study may appear to be a simple econometric exercise, and yet the results here have considerable implications for academia which often considers the DPR, which is assumed to be stationary, as one instrument by which to value stock prices.

2 Empirical Studies

2.1 Data

Our data are quarterly, and include the stock prices (P), dividends (D) and the dividend-price ratios (DPRs) for Australia, Canada, France, Germany, Italy, Japan, the Netherlands, Sweden, Switzerland, the UK, and the US. The sample period is from 1973Q1 to 2004Q2, and therefore we have the total observations of 1,386. All the data are obtained from DataStream except the US data. The latter are obtained from Professor Robert Shiller's homepage.[3] The definitions of the data are consistent with Campbell (2002) , and all data are converted into log form when investigating their time-series properties.

[2] Obviously, it would be best to analyze the panel data with possible structural breaks, but due to technical difficulties, we leave this for future investigation.

[3] http://www.econ.yale.edu/~shiller/. Campbell (2002) obtained all these data from the DataStream except the US data that are from the CRSP. Our decision to use Professor Shiller's data is because we do not have access to the CRSP and, unlike data for other countries, the DataStream does not contain S&P500 from the early 1970s. Finally, our sample period ends with 2004Q2 since US dividend data are available through this period from Professor Shiller's homepage. The beginning of the sample period is determined by the availability of the data in the DataStream.

2.2 Individual Unit Root Tests

As in previous studies summarized in the Introduction, we shall first of all examine whether the DPR is stationary using country specific data. This can be ascertained by the size of the differencing term, d. When $d = 0$, the data is said to follow a stationary process, $(I(0))$. As summarized in the previous section, the DPR is expected to be $I(0)$ while prices and dividends are non-stationary.

In order to examine the time-series properties of our data, we first implement unit root tests: the Augmented Dicky-Fuller-Generalized Least Square ($ADF-GLS$) (Elliott, Rothenberg, and Stock, 1996) and the Saikkonen-Lütkepohl (SL) tests. The latter test is implemented in order to account for the possible effects of unknown structural regimes in the data. Implementation of the SL test is important because in the presence of regime shifts, the standard unit root tests have insufficient power to reject the null hypothesis (Perron 1997).

More specifically, the SL unit root test examines the null hypothesis of the unit root ($d = 1$) based on the following general specification.

$$y_t = \mu_0 + \mu_1 t + f_t(\theta)'\gamma + x_t \tag{1}$$

where t is a time trend, and the residual, x_t, is assumed to follow a finite order autoregressive form, $b(L)(1-\rho L)x_t = \varepsilon_t$ where L is a lag operator, $-1 < \rho \leq 1$, and $\varepsilon_t \sim iid(0, \sigma^2)$. The shift function, $f_t(\theta)$, is dependent on θ and the regime shift date, T_B. We consider the following three forms of shift functions.

$$f_t^1\gamma = d_{1,t}\gamma = \begin{cases} 0 & t < T_B \\ \gamma & t \geq T_B \end{cases} \tag{2}$$

$$f_t^2(\theta)'\gamma = \begin{cases} 0 & t < T_B \\ (1 - \exp[-\theta(t - T_B + 1)])\gamma & t \geq T_B \end{cases} \tag{3}$$

$$f_t^3(\theta)'\gamma = \begin{cases} 0 & t < T_B \\ \gamma_1 & t = T_B \\ \gamma_1 + \sum_{j=1}^{t-T_B} \theta^{j-1}(\theta\gamma_1 + \gamma_2) & t > T_B \end{cases} \tag{4}$$

The shift function, f_t^1, represents a shift dummy indicating a permanent and swift level shift at time T_B, and is not dependent on θ. The second shift function, f_t^2, allows for a nonlinear gradual shift starting at T_B based on the exponential distribution. The rational function, f_t^3, is more flexible and can generate a smooth or abrupt shift from one regime to another.[4] The

[4] A smooth transition can be obtained for a smaller value of θ.

parameters, γ_1 and γ_2, correspond to those in the lag operator: $[\gamma_1(1-\theta L)^{-1}+\gamma_2(1-\theta L)^{-1}L]d_{1,t}$, which is another expression of $f_t^3(\theta)'\gamma$.

These tests are based on a two-step procedure. The statistics are obtained by first estimating parameters for the deterministic terms by the generalized least squares (GLS). The timing of the shift, T_B, corresponds to the date when the GLS objective function is minimized (Lanne *et al* 2002). The second step involves obtaining the t value *a la* the ADF test using the adjusted data. This adjusted data is equivalent to the difference between the original data and the deterministic components, and the OLS is used to calculate the t value. Lanne *et al* (2002) show that the performance of this procedure is reliable even in the case of small samples.

The results of the unit root tests are summarized in Tables 1 and 2. Table 1 highlights those of the $ADF-GLS$, and like previous studies shows that prices, dividends, and DPRs are non-stationary for all countries. Our result for the US is consistent with the ADF result from Balk and Wohar (2002) and Bohl and Siklos (2004). However, our result for the UK differs from Madsen and Milas (2005), which may be attributable to the different sample period under examination.[5] Therefore, based on our results from the analysis of the individual data, stock prices and dividends for most countries do not seem to move on a one-to-one basis even in the long-run.

Next, we shall consider the possibility of a structural break in the data. Table 2 reports results from the SL tests and shows the same outcomes as those from the $ADF-GLS$ test. It is very clear that stock prices and dividends are non-stationary in our samples. Furthermore, consistent with Table 1, DPRs are non-stationary for all countries. Interestingly, the fact that our results from the $ADF-GLS$ test are consistent with those of the SL test indicates that regime shifts in our data are not significant for our countries.[6] However, it is interesting to note that the t statistics for the US DPR from the SL tests are much higher (in absolute terms) than those of other countries, although not high enough to exceed its critical value of -2.880 (5 percent level). Furthermore, the fact that the SL statistics for the US DPR are closer to the rejection area compared with those from the $ADF-GLS$ is consistent with previous analyses (e.g., Liang and Sharp, 1999) and likely reflects the change in the payout policy of US firms opting to distribute profits in the form of repurchasing shares.

As our failure to reject the null hypothesis for these data may be attributable to the small

[5] Madsen and Milas (2005) use a dataset which covers the period from 1871 to 2002.

[6] We have also implemented Perron's unit root test (1997) that takes into account regime shifts. While the shift dates detected by Perron's test differ somewhat from those of the SL, the result as to whether or not the data are stationary is consistent with those from the SL test.

sample observations, we shall next examine the same problem in the panel data context.

2.3 Panel Data Studies

In order to improve the deficiencies of cross section (individual) unit root tests which may arise from the small number of observations, we shall first of all use the panel unit root tests developed by Im, Pesaran, and Shin [IPS] (2003), Fisher (1932), and Levin-Lin-Chu [LLC] (2002). To our knowledge, this is the first time to evaluate the DPR using panel unit root tests.[7]

The general statistical specification of these tests can be expressed as:

$$\Delta q_{i,t} = \alpha_i + \beta_i q_{i,t-1} + \sum_{j=1}^{k_i} \psi_{i,j} \Delta q_{i,t-j} + \varepsilon_{i,t}, \tag{5}$$

where $i = 1, ..., N$, $t = 1, ..., T$, and $\Delta q_{i,t} = q_{i,t} - q_{i,t-1.}$ The term, $q_{i,t}$, is a natural logarithmic form of the stock price, dividend, or DPR for country i at time t. Country specific events are captured by α_i, and θ_t captures the common time effect, and the term, $\varepsilon_{i,t}$, is a white noise residual.

As in the univariate ADF test, $q_{i,t}$ is covariance stationary (i.e., $I(0)$) when $|\rho_i| < 1$, where $\beta_i = \rho_i - 1$. If $\rho_i = 1$, then $q_{i,t}$ is said to follow the unit root process (i.e., $I(1)$). Statistical hypotheses differ slightly depending on the test employed. The LLC assumes a common unit root process across cross-sections (i.e., $\rho_1 = \rho_2 = \cdots = \rho_N = \rho$) and thus β_i becomes β in equation (1). Therefore, it tests the null hypothesis of $H_0 : \beta = 0$ against the alternative of $H_1 : \beta < 0$. In contrast, the IPS and Fisher tests relax the assumption of a common unit root process and analyze the null of $H_0 : \beta_i = 0$ for all i against the alternative of $H_1 : \beta_i < 0$ for $i = 1, 2, \cdots, N_1$ and $\beta_i = 0$ for $i = N_1 + 1, N_1 + 2, \cdots, N$.[8] One difference between these two tests is that while the IPS computes a panel unit statistic based on a modified t statistic,[9] the Fisher test relies on p-values for individual ADF tests. This test statistic is asymptotically χ^2 distributed, i.e., $-2\sum_{i=1}^{N} \log(p_i) \sim \chi^2_{2N}$, where p_i is p-values for individual, i. It is important to note that all these tests have been developed under the assumption of no cross-sectional dependence.

However, we all know that economic and financial data are likely correlated. For example, a significant and positive correlation is reported between stock returns in Japan and the US

[7] For details of these tests, see Maddala and Kim (1998). Among these tests, the Fisher, followed by the IPS, seems to be most reliable in terms of its ability to distinguish statistical hypotheses.

[8] Both the IPS and Fisher tests are based on N cross-section ADF unit root tests, and thus their testable specification is identical to equation (5).

[9] The IPS test is based on the average of t statistics of cross-section ADF tests. Based on the Monte Carlo experiments, the IPS provides an appropriate size of mean and variance in order to adjust this average t statistic to the context of the panel data.

(Nagayasu 2006). In the presence of cross sectional dependence, O'Connell (1998) argues that the standard test will suffer from significant size distortion, and demonstrates using real exchange rates the importance of controlling cross sectional dependence when conducting panel unit root tests by reporting the contrasting results from the conventional panel unit root tests. Since then, much research has been conducted in order to address this problem. For example, Pesaran (2006) advocated the individual cross sectionally augmented ADF ($CADF$) test. His idea can be summarized using the following specification.

$$\Delta q_{i,t} = \alpha_i + \beta_i q_{i,t-1} + \gamma_i f_t + \varepsilon_{i,t}, \tag{6}$$

where $f_t$is the unobserved common effect which is absent in equation (5). This procedure assumes one common effect. Like the standard ADF test, the null hypothesis of the unit root can be evaluated by $H_0 : \beta_i = 0$ for all i against the alternative $H_0 : \beta_i < 0$ for some i. Pesaran (2006) argues that this common factor can be proxied by the cross sectional mean of $q_{i,t}$ ($\overline{q}_t = N^{-1} \sum_{j=1}^{N} q_{i,t}$) and its lagged variables

$$\Delta q_{i,t} = \alpha_i + \beta_i q_{i,t-1} + \gamma_i \overline{q}_{t-1} + \phi_{i,j} \Delta \overline{q}_t + \varepsilon_{i,t}, \tag{7}$$

The CADF statistic, $t_i(N,T)$, can be calculated by pooling the t statistics obtained from the OLS regression for parameter β_i.

$$CIPS(N,T) = N^{-1} \sum_{i=1}^{N} t_i(N,T) \tag{8}$$

Furthermore, in order to avoid obtaining an extreme statistic due to the small (finite) sample observations (e.g., small T with $i = 10$ to 20), Pesaran proposes a truncated version of the $CIPS$. For a model with a constant term, this modified statistic can be expressed as:

$$CIPS^*(N,T) = N^{-1} \sum_{i=1}^{N} t_i^*(N,T) \tag{9}$$

Based on the simulation, he suggests the following threshold values.

$$t_i^*(N,T) = t_i(N,T) \text{ if } -6.19 < t_i(N,T) < 2.61$$
$$t_i^*(N,T) = -6.19 \text{ if } t_i(N,T) \leq -6.19$$
$$t_i^*(N,T) = \text{if } t_i(N,T) \geq 2.61$$

Furthermore, Pesaran extends this basic specification in order to take into account serial correlation, which can be written as:

$$\Delta q_{i,t} = \alpha_i + \beta_i q_{i,t-1} + \gamma_i \bar{q}_{t-1} + \sum_{j=0}^{k} \phi_{ij} \Delta \bar{q}_{t-j} + \sum_{j=1}^{k} \psi_{i,j} \Delta q_{i,t-j} + \varepsilon_{i,t}, \tag{10}$$

Since these statistics do not follow the conventional distribution under the null, Pesaran (2006) proposes several forms to evaluate the null hypothesis based on the $CIPS$ and $CIPS^*$, including the Fisher-type test following Maddala and Wu (1999). Furthermore, the critical values are provided in Pesaran (2006) according to the size of N and T and the composition of the deterministic terms in the specification. This paper calculates the $CIPS^*$ and employs the critical values reported in Pesaran (2006) which seems more appropriate for our finite data sample.

The results are reported in Table 3 where appropriate lag lengths are determined by the Schwarz information criterion (SIC). This table provides strong evidence of a stationary DPR and non-stationary prices and dividends when the $CADF$ is employed. We also conducted a sensitivity analysis by excluding the US DPR from our data set since this data was found to be non-stationary in some previous studies. However, our result remains unchanged. Our conclusion is consistent with the theoretical view (Campbell and Shiller 1989) that dividends tend to move equi-proportionally along with stock prices in the long-run. The fact that other tests, the LLC, IPS, and Fisher, failed to reject the null for the DPRs underlines the existence of significant cross-sectional dependence in our data, and suggests the importance of properly addressing it.[10]

Finally, we checked whether the assumption of one common effect, which the $CADF$ hinges on, is supported by the data. In this connection, the modified SIC developed to analyze the number of common effects in the panel data context (Bai and Ng 2002) is implemented. Amongst several criteria discussed in Bai and Ng (2002), this paper uses the modified SIC which is found to perform better in cases of finite sample analysis. Like the standard information criteria, a smaller SIC indicates a better fit of model to data. Here, the SIC is calculated with the maximum of two common trends, and our result suggests one common trend for all data. Therefore, it seems appropriate to use the $CADF$.

[10] The size of ρ ($\beta_i = \rho_i - 1$) ranges from 0.829 to 0.981 based on the $CADF$ procedure. Some uncertainty exists regarding the size of this parameter due to the statistical power of the unit root test, but the $CADF$ is a useful test since critical values are provided for a number of finite sample cases.

3 Summary and Discussion

This paper examines the time-series properties of stock prices, dividends, and dividend-price ratios. Using the panel unit root test developed by Pesaran (2006) which takes account of cross-sectional dependency in the data, we provide very clear evidence that while stock prices and dividends are non-stationary, the dividend-price ratios are stationary. This clear-cut result stands in sharp contrast to previous studies that evaluate country specific data on the DPR. We argue that our finding results from the implementation of a more advanced (panel unit root) test, which takes account of cross-sectional correlations in the data and which given the sample period (i.e., fixed T) increases the sample observations but maintains the likelihood of structural breaks. Overall, our findings support a one-to-one effect of price changes on dividends, and confirm the dividends as an important form of payout in these countries.

It should be noted that our finding is not inconsistent with previous studies that used a longer sample period for a specific country and found a stationary dividend-price ratio when taking into account structural shifts in the data. In this case, the failure to reject the stationary dividend-price ratio by the standard test (e.g., $ADF - GLS$) may be attributable to regime shifts since the longer the sample, the more likely the existence of regime shifts in the data. We did not find strong evidence of structural breaks here, and this may result from our relatively short sample period, required to create a balanced panel data set. Our findings are valid as long as the dividend-price ratios are considered for a group of countries for our sample period.

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Table 1: Individual Unit Root Test (ADF-GLS)

	P (lag)	D (lag)	D (lag)
Australia	-2.082 (1)	-2.076 (1)	-2.116 (1)
Canada	-2.503 (1)	-2.699 (1)	-2.957 (2)
France	-1.764 (1)	-1.971 (1)	-1.736 (1)
Germany	-1.978 (1)	-2.217 (1)	-1.932 (1)
Italy	-1.886 (1)	-1.898 (1)	-1.870 (1)
Japan	-1.016 (1)	-1.373 (1)	-0.998 (1)
Netherlands	-1.439 (1)	-1.882 (3)	-1.386 (1)
Sweden	-1.916 (1)	-2.041 (1)	-1.892 (1)
Switzerland	-1.243 (1)	-1.288 (1)	-1.242 (1)
UK	-1.897 (1)	-1.819 (1)	-1.882 (1)
US	-1.595 (1)	-0.865 (1)	-1.817 (2)

Note. The data are log stock prices (P), dividends (D), and the dividend-price ratio (DPR). This test is based on the Augmented Dickey-Fuller (ADF)-Generalized Least Square (GLS). The sample period is from 1973Q1 to 2004Q2. The appropriate lag length is determined by the Schwarz information criterion with a maximum of four. The critical value at the five percent significance level is -3.030 for $T = 100$ (Elliott, Rothenberg and Stock, 1996).

Table 2: Individual Unit Root Tests with Possible Structural Breaks

	P (lag)	D (lag)	DPR (lag)
Shift type: Shift dummy			
Australia	-0.469 (0)	-0.505 (0)	-0.493 (0)
Canada	-0.438 (0)	-0.130 (0)	-0.590 (1)
France	-0.280 (0)	-0.264 (0)	-0.433 (0)
Germany	-0.703 (0)	-0.713 (0)	-0.572 (0)
Italy	-0.474 (0)	-0.465 (0)	-0.590 (0)
Japan	-1.376 (0)	-1.375 (0)	-1.034 (0)
Netherlands	-0.452 (0)	-0.466 (0)	-0.066 (0)
Sweden	-0.824 (0)	-0.855 (0)	-0.786 (0)
Switzerland	0.206 (0)	0.200 (0)	0.157 (0)
UK	-0.594 (2)	-0.598 (2)	-0.452 (0)
US	-0.021 (0)	-1.021 (2)	-2.586 (0)
Shift type: Exponential			
Australia	-0.475 (0)	-0.496 (0)	-1.301 (0)
Canada	-0.431 (0)	-0.099 (0)	-0.652 (1)
France	-0.280 (0)	-0.263 (0)	-0.439 (0)
Germany	-0.698 (0)	-0.706 (0)	-0.571 (0)
Italy	-0.464 (0)	-0.315 (0)	-0.593 (0)
Japan	-1.337 (0)	-1.341 (0)	-1.041 (0)
Netherlands	-0.446 (0)	-0.454 (0)	0.423 (0)
Sweden	-0.832 (0)	-0.863 (0)	-0.869 (0)
Switzerland	0.210 (0)	0.205 (0)	0.157 (0)
UK	-0.513 (2)	1.310 (3)	-0.431 (0)
US	-0.271 (0)	-0.826 (2)	-2.588 (0)
Shift type: Rational			
Australia	-0.673 (0)	-0.803 (0)	-0.691 (0)
Canada	-0.418 (0)	-0.495 (0)	-0.779 (1)
France	-0.536 (0)	-0.527 (0)	-0.303 (0)
Germany	-1.089 (0)	-1.102 (0)	-0.978 (0)
Italy	-0.588 (0)	-0.429 (0)	-0.361 (0)
Japan	-1.522 (0)	-1.535 (0)	-0.959 (0)
Netherlands	-0.892 (0)	-0.906 (0)	0.283 (0)
Sweden	0.822 (0)	0.746 (0)	-0.919 (0)
Switzerland	-0.058 (0)	-0.062 (0)	0.001 (0)
UK	-0.837 (0)	-0.840 (0)	-0.541 (0)
US	-0.276 (0)	-2.009 (2)	-2.710 (2)

Note: The data are log stock prices (P), dividends (D), and the dividend-price ratio (DPR). These three types of unit root test were developed by Saikkonen and Lütkepohl (2002) and are based on shift functions (2), (3) and (4) in the main text respectively. The constant term is included in the unit root test. The critical value (the five percent significance level) for these tests is -2.880 (Lanne *et al* 2002). The appropriate lag orders are determined by the Schwarz information criterion.

Table 3: Panel Unit Root Tests

	All countries			Without US
	P (p-value)	D (p-value)	DPR (p-value)	DPR (p-value)
LLC	-0.273 (0.392)	0.037 (0.515)	-0.665 (0.253)	-0.990 (0.161)
IPS	-1.609 (0.381)	-1.486 (0.568)	-1.637 (0.340)	-1.077 (0.141)
Fisher	3.896 (0.999)	6.528 (0.994)	9.575 (0.990)	3.208 (0.999)
CADF	-2.176	-2.227	-2.528	-2.375
SIC				
f=0	0.0102	0.0099	0.0381	0.0386
f=1	0.0067	0.0066	0.0231	0.0207
f=2	0.0081	0.0078	0.0254	0.0244

The panel unit root tests are conducted for a group of log stock prices (P), dividends (D), and the dividend-price ratio (DPR). The statistical tests are LLC (Levin-Lin-Chu, 2002), IPS (Im, Pesaran, and Shin, 2003), and Fisher (1932). The CADF test is the individually cross section dependence test (Pesaran 2006). Of four tests, only the CADF considers cross sectional dependency in the data when calculating the statistics. The critical value (5 percent) is -2.250. Detailed explanations of these tests are provided in Section 2.3. The *SIC* is the Schwarz information criterion modified by Bai and Ng (2002) for panel data analysis.