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The random walk hypothesis revisited: evidence from the 16 OECD stock prices

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Abstract

Using 16 OECD stock price indices data, this paper revisits the random walk hypothesis by inspecting the degree of persistence of stock prices. We adopt two recently developed econometric procedures, due to Hansen (1999) and Romano and Wolf (2001), in order to estimate 95% confidence intervals for the sum of the AR coefficients in AR representations of international stock prices. Confidence intervals provide much more information than knowing whether the null hypothesis of a unit root can be rejected or not. They serve as a measure of sampling uncertainty and describe the range of models that are consistent with the observed data. The results convincingly support the view that the stock price indices in the OECD countries are highly persistent. The high persistence in the OECD stock price indices provides strong evidence for the random walk hypothesis.

1 Introduction

Testing for a unit root or the random walk hypothesis (hereafter RWH) in stock price has been attracted substantial interest in the empirical finance literature ever since the studies conducted by Fama and French (1988a, 1988b), Lo and MacKinlay (1988) and Poterba and Summers (1988). This is because if there is a unit root in stock price, then this implies that stock market returns cannot be predicted from previous prices changes. Therefore, given only past price and return data, the current price is the best predictor of the future price, and price change or return is expected to be zero. This is the essence of the weak-form efficient market hypothesis (hereafter WEMH). It also implies that shocks have permanent effects and volatility in stock markets will increase in the long run without bound. However, if stock prices follow a mean reverting process, then there exists a tendency for the price level to return to its trend path over time and investors may be able to forecast future returns by using information on past returns.

A wealth of researches has been devoted their efforts to this issue. For example, to name a few, McQueen (1992), Urrutia (1995), Zhu (1998), Grieb and Reyes (1999), Chaudhuri and Wu (2003), Narayan (2005, 2006, 2008), Narayan and Smyth (2004). The findings are mixed, if not contradictory, which means there is no corroborative conclusion vis-à-vis the stationarity property for stock prices. Moreover, the majority applies the traditional method in testing for the null hypothesis of a unit root of stock prices.¹ Following the works of Lo and MacKinlay (1988, 1989), many researchers attempted to use the variance ratio test (see Chow and Denning, 1993; Richardson, 1993) for a random walk hypothesis and challenged the earlier findings. For recent applications, readers are referred to Belaire-Franch and Opong (2005a, 2005b), Hoque et al. (2007) and Kim and Shamsuddin

¹It is well-known that the traditional unit root test is powerless if the true data generating process of a series exhibits structural breaks (Perron, 1989). Therefore, a few of studies, e.g. Narayan and Smyth (2005, 2006, 2007), adopt new developed unit root test with structural breaks (Zivot and Andrews, 1992; Lumsdaine and Papell, 1997; Lee and Strazicich, 2003) to investigate the stationary property of stock prices.

(2008), and papers therein for details.²

The central aim of this paper is to revisit the RWH for 16 OECD stock price indices. However, the methodology used in this paper is different from the previous studies. Previous researches, employing conventional unit root tests, provide limited information on the degree of persistence in stock prices, as unit root tests concentrate solely on testing the null hypothesis that the sum of the autoregressive (AR) coefficients is unity in an AR representation of a series against the alternative hypothesis that the sum of the AR coefficients is less than unity. In contrast, a confidence interval for the sum of the AR coefficients provides us with a more informative statistical description of a variable's persistence, and, is helpful to testing the RWH. In this paper, we employ two recently developed econometric procedures, due to Hansen (1999) and Romano and Wolf (2001), in order to estimate 95% confidence intervals for the sum of the AR coefficients in AR representations of international stock prices. To the best of our knowledge, this paper is the first one to study the RWH of the stock prices by using the estimates of confidence interval.

As pointed out by Rapach and Wohar (2004), unlike conventional or bootstrapped confidence intervals, the Hansen (1999) and Romano and Wolf (2001) procedures generate confidence intervals for nearly integrated variables with correct first-order asymptotic coverage for the sum of the AR coefficients. Using Monte Carlo simulations, Hansen (1999) and Romano and Wolf (2001) find that their respective procedures for constructing asymptotically valid confidence intervals also provide good coverage in finite samples. In addition to the sum of the AR coefficients, we measure persistence through the half-life, or the number of years required for a shock to a variable to dissipate by one-half, as this is a popular measure of persistence outlined in Gospodinov (2004).

²There are some studies, however, find that stock price indices could be characterized as non-linear data generating processes. For example, Abhyankar et al. (1995, 1997), Atchison and White (1996), Kohers et al. (1997), Schaller and van Norden (1997), Qi (1999), Kanas (2001), Sarantis (2001), Shively (2003) Narayan (2005, 2006) and Chen (2008).

The remainder of this paper is organized as follows. Section 2 introduces the econometric methodology that we employ, and Section 3 describes the data and the empirical test results. Section 4 presents the conclusions that we draw from this research.

2 Methodology

This section provides a brief description of the Hansen (1999) and Romano and Wolf (2001) procedures.³ Let y_t denote the logarithm of the stock price index. Consider the following $AR(p)$ process for the variable y_t :

$$y_t = \mu + \alpha_1 y_{t-1} + \alpha_2 y_{t-2} + \cdots + \alpha_p y_{t-p} + e_t \quad (1)$$

for $t = 1, 2, \dots, T$. Andrews and Chen (1994) argue that an informative scalar measure of persistence in the AR process is the sum of the AR coefficients, $\alpha = \sum_{i=1}^p \alpha_i$, as the cumulative impulse response (CIR, the sum of the impulse response function over all time horizons) is related to α via $CIR = 1/(1 - \alpha)$. Andrews and Chen (1994) view α as more informative than the largest root of the AR model, since two $AR(p)$ models with identical largest roots can have very different persistence properties.

We can straightforwardly obtain a point estimate of α by rearranging equation (1) and using OLS to estimate the familiar augmented Dickey and Fuller (1979, ADF) regression model:

$$y_t = \mu' + \alpha y_{t-1} + \sum_{j=1}^k \beta_j \Delta y_{t-j} + e_t \quad (2)$$

where $\Delta y_t = y_t - y_{t-1}$. The construction of confidence intervals for α is problematic because the asymptotic distribution of the OLS estimator (as well as its rate of convergence) is different in the stationary and unit-root cases.

³The expositions of this section draw heavily from Rapach and Wohar (2004), who employ the same approach to investigate the persistence in international real interest rates for 13 industrialized countries. Interested readers can refer to Hansen (1999) and Romano and Wolf (2001) for a full theoretical derivation.

Hansen (1999) developed a procedure for constructing confidence intervals for α with correct first-order asymptotic coverage.⁴ More specifically, consider a grid of values for α , $\alpha_i (i = 1, \dots, B)$, covering $\hat{\alpha}$. In order to estimate the data-generating process for each α_i , we estimate equation (2), with α restricted to α_i , using restricted *OLS* for each α_i . The restricted *OLS* parameter estimates, together with resampled restricted *OLS* residuals, are used to build up a large number of pseudosamples (say, 5000) for each α_i . For each of the 5000 pseudo-samples for each α_i , we calculate the *t*-statistic, $t_i^* = (\hat{\alpha}_i^* - \alpha_i) / s(\hat{\alpha}_i^*)$, where $\hat{\alpha}_i^*$ is the *OLS* estimate of α in Equation (2) for a given pseudo-sample and α_i grid value. We sort the *t*-statistics, giving us an empirical distribution of *t*-statistics for each α_i from which we can calculate the 0.025 and 0.975 quantiles of *t*-statistics for each α_i . The upper bound for the 95% confidence interval for α is the α_i grid value such that $(\hat{\alpha} - \alpha_i) / s(\hat{\alpha}) = t_{i,0.925}^*$. The lower bound is the α_i grid value such that $(\hat{\alpha} - \alpha_i) / s(\hat{\alpha}) = t_{i,0.975}^*$.

Romano and Wolf (2001) developed a subsampling procedure for constructing confidence intervals for α that also provides correct first-order asymptotic coverage.⁵ This approach recomputes the *OLS* estimator on smaller blocks, or subsamples, of the observed series. More specifically, we begin with a block of size b and calculate the *t*-statistic, $\tau_b(\hat{\alpha}_{b,t} - \hat{\alpha}) / \hat{\sigma}_{b,t}$, for each subsample of size b for $t = 1, \dots, T - b + 1$, where $\hat{\alpha}_{b,t}$ is the *OLS* estimate of α for the t th block of size b , $\hat{\sigma}_{b,t} = b^{1/2} s(\hat{\alpha}_{b,t})$, and $\tau_b = b^{1/2}$. We generate the empirical approximating distribution for the subsample *t*-statistics:

$$L_b(x) = \frac{1}{T - b + 1} \sum_{t=1}^{T-b+1} 1\{\tau_b(\hat{\alpha}_{b,t} - \hat{\alpha}) / \hat{\sigma}_{b,t} \leq x\} \quad (3)$$

Let $c_{b,0.025}$ and $c_{b,0.975}$ be the 0.025 and 0.975 quantiles of the subsampling distribution,

⁴There is a well-known difficulty in constructing confidence intervals for the sum of the AR coefficients: conventional asymptotic or bootstrapped confidence intervals are not valid for this key measure of persistence when the data are generated by a nearly integrated process (Basawa et al., 1991).

⁵A potential advantage of the Romano and Wolf (2001) subsampling procedure is that, unlike the Hansen (1999) grid-bootstrap procedure, it does not require the assumption that e_t is independently and identically distributed (i.i.d.) in equations (1) or (2), as it is still valid for dependent error processes.

equation (3). A 95% two-sided equal-tailed confidence interval for α is given by

$$[\hat{\alpha} - (1/\tau_T)s(\hat{\alpha})c_{b,0.975}, \hat{\alpha} + (1/\tau_T)s(\hat{\alpha})c_{b,0.025}] \quad (4)$$

where $\tau_T = T^{1/2}$. Romano and Wolf (2001) also discuss the construction of a two-sided symmetric subsampling interval. Instead of equation (3), we generate the empirical approximating distribution,

$$L_{b,|\cdot|}(x) = \frac{1}{T-b+1} \sum_{t=1}^{T-b+1} 1\{\tau_b|\hat{\alpha}_{b,t} - \hat{\alpha}|/\hat{\sigma}_{b,t} \leq x\} \quad (5)$$

Let $c_{b,|\cdot|,0.05}$ be the 0.05 quantile for the empirical distribution, equation (5). A 95% two-sided symmetric confidence interval is given by

$$[\hat{\alpha} - (1/\tau_T)s(\hat{\alpha})c_{b,|\cdot|,0.05}, \hat{\alpha} + (1/\tau_T)s(\hat{\alpha})c_{b,|\cdot|,0.05}] \quad (6)$$

We follow algorithm 5.1 (minimizing confidence interval volatility) in Romano and Wolf (2001, p. 1297) in order to select b . The algorithm proceeds as follows:

1. Compute a subsampling 95% confidence interval for α for each b in $b = b_{small}$ to $b = b_{big}$, yielding the endpoints $I_{b,low}$ and $I_{b,up}$. Set $b_{small} = c_1T^\eta$ and $b_{big} = c_2T^\eta$ for $0 < c_1 < c_2$ and $0 < \eta < 1$. Romano and Wolf (2001) recommend $c_1 \in [0.5, 1]$, $c_2 \in [2, 3]$ and $\eta = 0.5$. (We set $c_1 = 1$, $c_2 = 3$ and $\eta = 0.5$.)
2. For each b , compute a volatility index, VI_b , where the volatility index is the standard deviation of the interval endpoints in a neighbourhood of b . That is, for a small integer k , let VI_b equal the standard deviation of $\{I_{b-k,low}, \dots, I_{b+k,low}\}$ plus the standard deviation of $\{I_{b-k,up}, \dots, I_{b+k,up}\}$. Romano and Wolf recommend $k = 2$ or $k = 3$. (We set $k = 2$.)

Select the value for b , b^* , with the smallest volatility index and report $[I_{b^*,low}, I_{b^*,up}]$ as the final 95% subsampling confidence interval.

3 Data and Results

We use the stock price indices for 16 OECD countries, i.e., Australia, Austria, Canada, France, Germany, India, Ireland, Italy, Japan, South Korea, Netherlands, New Zealand, Switzerland, Swiss, the UK and the USA in our empirical study. The data set is obtained from the OECD Main Economic Indicators at <http://stats.oecd.org/mei/>. For all countries the data are quarterly from different starting date (see column (2) in Table 3) but they are all ended in 2008Q1. Log transformation for stock prices are used throughout the study.

We begin by applying various unit root tests, including of the Augmented Dickey-Fuller (1979, hereafter ADF) test, Schmidt and Phillips (1992, hereafter SP) test, the Kwiatkowski et al. (1992, thereafter KPSS) test, the Elliott et al. (1996, hereafter DF-GLS) modified ADF test and the Ng and Perron (2001, thereafter NP) test, to ascertain the order of integration of the variables. The key here is to account for serial correlation in conducting the unit root test. We set the maximum lag order $k = 12$, which is the lagged difference, and use the Ng and Perron (2001) modified Akaike Information Criterion (MAIC) to select the optimal lag length. In order to take into account possible shift in regime in the unit root test, we also consider the Zivot and Andrews (1992, thereafter ZA) test that allows an endogenous structural break. We adopt the Akaike Information Criterion (AIC) to select the optimal lag length for the ZA.

The ADF, SP, DF-GLS, ZA and NP testing principles share the same null hypothesis of a unit root. In contrast, the KPSS procedure tests for level (η_μ) or trend stationary (η_τ) against the alternative of a unit root. In this sense, the KPSS principles involve different maintained hypothesis from the ADF or the DF-GLS unit root test. The NP test is a modified version of the Phillips and Perron (1988) test which allows, first, to correct the size distortions (as suggested by Perron and Ng, 1996), second, to improve the power (as suggested by Elliott et al, 1996).

We summarize the various unit root test results in Table 1. Basically, we find no additional evidence against the unit root hypothesis at the 5% significance based on the ADF,

SP, DF-GLS and NP tests in their level data but with minor exceptions. For example, the results of the ADF τ_τ statistic for Canada and NP \overline{MZ}_t statistic for Ireland, but they are insignificant at the 1% level, indicating that the null hypothesis of a unit root cannot be rejected. However, as Perron (1989) pointed out, in the presence of a structural break, the power to reject a unit root diminishes if the stationary alternative is true and the structural break is ignored. To address this, we use model C of Zivot and Andrews' (1992) sequential trend break model to investigate the order of the empirical variables. The results from Column (9) in Table 1 generally suggest empirical variables are non-stationary in their levels. When we apply the ADF, SP, KPSS and DF-GLS tests to the first difference of these series (see Table 2), we must reject the null hypothesis of a unit root at the 5% level or better. This implies that the stock prices of these 16 OECD countries have a unit root.

As aforementioned in the introduction, conventional unit root tests alone provide limited information on the degree of persistence in stock prices, as they concentrate solely on testing the null hypothesis that the sum of the autoregressive coefficients is unity in an AR representation of a series against the alternative hypothesis that the sum of the AR coefficients is less than unity. We therefore turn to apply two recently developed econometric procedures, due to Hansen (1999) and Romano and Wolf (2001), in order to estimate 95% confidence intervals for the sum of the AR coefficients in AR representations of international stock prices.

Column (3) of Table 3 reports the OLS estimate of α in equation (2) for each country. The OLS point estimates of α are greater than or equal to 0.90 for every country, with the exception of Canada. Of course, these point estimates are biased downwards and are of limited value. In order to provide more informative measures of persistence, we calculate Hansen (1999) grid-bootstrap and Romano and Wolf (2001) subsampling 95% confidence intervals for α for every country. As pointed out by Rapach and Wohar (2004), these confidence intervals provide valid asymptotic first-order coverage and appear to have good coverage in finite samples.

We report the Hansen (1999) grid-bootstrap 95% confidence interval for α in column (4) of Table 3.⁶ Observe that the lower bound of the grid-bootstrap confidence interval is greater than or equal to 0.90 for every country, with the exceptions of Canada and Ireland. For Canada and Ireland, the lower bound is still quite close to 0.90. The upper bound of the 95% confidence interval is greater than unity for every country, so the data are not inconsistent with a unit root in the stock price for every country. According to the Hansen (1999) grid-bootstrap 95% confidence intervals for α , the lower bounds for the 95% confidence intervals indicate that stock price indices display a high degree of persistence.

Romano and Wolf (2001) equal-tailed and symmetric subsampling 95% confidence intervals for α are found in columns (5) and (6) of Table 3. For the most part, these confidence intervals are similar to the grid bootstrap confidence intervals. For the equal-tailed intervals reported in column (5), the lower bounds are greater than 0.90, with the exception of, again, Canada and Ireland, and the lower bound is quite close to 0.9 for Canada and Ireland. The upper bound is greater than unity for every country with the exception of Canada (and the upper bound is still very close to unity for Canada). The symmetric intervals, reported in column (6), are similar to the equal-tailed subsampling intervals, although both the upper and lower bounds for the symmetric intervals appear to be somewhat smaller on average than those for the equal-tailed intervals. Nevertheless, with the exception of Australia, Austria, Canada, Ireland, South Korea, Switzerland and the UK, the lower bounds for symmetric intervals are greater than 0.90. For Austria and the UK, the lower bound is still quite close to 0.9. The upper bound is greater than unity for every country with the exception of Canada and Swiss, and the upper bound is still very close to unity for the two countries. Therefore, we can conclude that most OECD stock price indices are characterized by a random walk based on the Romano and Wolf (2001) equal-

⁶All computations are implemented by using the GAUSS program available from Professor David E. Rapach's homepage at <http://pages.slu.edu/faculty/rapachde/Research.htm>. We thank him for making his computer code publicly available in his homepage.

tailed and symmetric subsampling 95% confidence intervals for α , with the exception of Canada stock price index.

Finally, we compute percentile grid-bootstrap 95% confidence intervals for the half-life of the impulse response function using the procedure outlined in Gospodinov (2004). The estimated results is reported in column (7) of Table 3, and the grid-bootstrap 95% confidence interval for the half-life is reported in column (8). Note that the half-lives are measured in years in Table 3. The confidence intervals are very wide for all of the half-lives, and an infinite upper bound (corresponding to a nonstationary stock price) is found for all countries. Again, the data are consistent with a high degree of persistence and in line with the RWH.

4 Concluding Remarks

The purpose of this study is to investigate the random walk hypothesis by examining the degree of persistence in international stock prices of 16 OECD countries based on the recently developed Hansen (1999) grid-bootstrap and Romano and Wolf (2001) subsampling procedures. These procedures generate confidence intervals for the sum of the AR coefficients with correct first-order asymptotic coverage that also have good coverage in finite samples. Our results indicate a high degree of persistence in quarterly stock prices. The lower bound for the 95% confidence interval for the sum of the AR coefficients is often greater than 0.90, while the upper bound is almost always greater than unity, for the countries we consider. Grid-bootstrap 95% confidence intervals for the half-lives have lower bounds ranging from approximately 2 to 8 years, while the upper bound is infinite for every country. Overall, the lower bounds for the 95% confidence intervals for the sum of the AR coefficients and the half-lives indicate that international stock prices are highly persistent and in line with the random walk hypothesis. In contrast to Chaudhuri and Wu (2003) who find considerable evidence of mean reversion in emerging markets, but consistent with Narayan and Smyth's (2005, 2006, 2007) findings that most OECD stock

price indices are characterized by a random walk, the only country for which one can reject the random walk hypothesis is Canada.

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Table 1: Results of the Unit Root Tests for the Level Data

(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)
	ADF		SP	KPSS		DF-GLS		ZA	NP	
country	τ_μ	τ_τ	τ	η_μ	η_τ	$Z(t) = (1)$	$Z(t) = (1, t)$	model C	\overline{MZ}_α	\overline{MZ}_t
Australia	0.223	-3.076	-2.29	4.004**	0.392**	2.265	-2.255	-4.406	-9.789	-2.203
Austria	-0.885	-3.109	-1.51	3.603**	0.393**	2.324	-1.386	-4.506	0.496	-2.330
Canada	0.286	-3.927**	-2.82	4.188**	0.373**	2.465	-2.186	-4.869	-6.783	-1.756
France	-0.483	-2.233	-1.50	4.037**	0.698**	1.954	-1.448	-3.383	-6.696	-1.822
Germany	-0.140	-2.907	-1.76	3.609**	0.479**	1.045	-1.545	-3.719	-8.521	-2.040
India	1.474	-1.384	-1.18	3.960**	0.759**	3.785	-0.493	-2.909	-2.499	-0.955
Ireland	-0.824	-3.204	-1.98	4.256**	0.308**	2.608	-2.093	-4.431	-17.126	-2.922**
Italy	-0.785	-1.880	-1.28	3.594**	0.560**	1.360	-1.268	-3.291	-7.678	-1.958
Japan	-1.933	-1.277	-1.15	3.547**	0.731**	1.152	-0.436	-2.738	-3.538	-1.139
Korea	-1.246	-1.908	-1.19	1.470**	0.334**	0.992	-1.263	-3.577	-6.163	-1.778
Netherlands	-0.349	-2.097	-1.41	3.733**	0.672**	1.846	-1.396	-3.360	-8.924	-2.110
New Zealand	-1.212	-1.885	-1.58	2.927**	0.344**	0.716	-1.477	-4.668	-12.804	-2.489
Switzerland	-0.229	-1.829	-1.54	4.605**	0.746**	3.022	-1.490	-4.033	-5.274	-1.619
Swiss	-0.703	-2.253	-1.70	3.910**	0.498**	2.169	-1.709	-4.174	-10.269	-2.262
UK	-0.953	-2.926	-1.91	4.060**	0.368**	2.135	-1.785	-3.835	-12.006	-2.384
USA	0.727	-2.109	-1.48	4.008**	0.811**	3.302	-1.330	-3.872	-6.997	-1.840

(1) The critical values for the ADF τ_μ and τ_τ statistics at the 10%, 5%, 1% significance levels are -2.57, -2.88, -3.46 and -3.13, -3.43, -3.99, respectively.

(2) The critical value for the SP τ statistic at the 5% significance level is -3.04.

(3) The critical values for the KPSS η_μ and η_τ statistics at the 10%, 5%, 1% significance levels are 0.347, 0.463, 0.739 and 0.119, 0.146, 0.216, respectively.

(4) The critical values for the DF-GLS $Z(t)=(1)$ and $Z(t)=(1, t)$ statistics at the 10%, 5%, 1% significance levels are -1.62, -1.95, -2.58 and -2.57, -2.89, -3.48, respectively.

(5) The critical values for the ZA test for model C at the 5% and 1% significance levels are -5.08 and -5.57, respectively.

(6) The critical values for the NP \overline{MZ}_α and \overline{MZ}_t statistics at the 5% significance levels are -17.3 and -2.91, respectively.

(7) *, ** and *** denote significant at the 10%, 5% and 1% level, respectively.

Table 2: Results of the Unit Root Test for Differenced Data

(1) Differenced data	(2) ADF	(3) SP	(4) KPSS	(5) DF-GLS
Country	τ_{μ}	τ	η_{μ}	$Z(t) = (1)$
Australia	-4.869**	-9.39**	0.043	-11.535**
Austria	-5.737**	-9.21**	0.086	-5.690**
Canada	-7.456**	-10.15**	0.075	-11.463**
France	-10.460**	-8.66**	0.086	-10.165**
Germany	-3.582**	-6.26**	0.094	-6.965**
India	-7.271**	-10.83**	0.432	-7.460**
Ireland	-6.673**	-9.26**	0.040	-9.992**
Italy	-5.847**	-7.87**	0.082	-9.500**
Japan	-9.259**	-6.88**	0.348	-7.490**
Korea	-7.288**	-5.56**	0.115	-5.338**
Netherlands	-4.582**	-6.59**	0.083	-6.683**
New Zealand	-3.993**	-8.38**	0.087	-9.044**
Switzerland	-5.385**	-9.64**	0.086	-11.137**
Swiss	-10.948**	-8.95**	0.052	-10.563**
UK	-11.088**	-8.09**	0.083	-6.921**
USA	-10.523**	-7.55**	0.125	-9.391**

(1) The critical values for the ADF τ_{μ} and τ_{τ} statistics at the 10%, 5%, 1% significance levels are -2.57, -2.88, -3.46 and -3.13, -3.43, -3.99, respectively.

(2) The critical value for the SP τ statistic at the 5% significance level is -3.04.

(3) The critical values for the KPSS η_{μ} and η_{τ} statistics at the 10%, 5%, 1% significance levels are 0.347, 0.463, 0.739 and 0.119, 0.146, 0.216, respectively.

(4) The critical values for the DF-GLS $Z(t)=(1)$ and $Z(t)=(1, t)$ statistics at the 10%, 5%, 1% significance levels are -1.62, -1.95, -2.58 and -2.57, -2.89, -3.48, respectively.

(5) *, ** and *** denote significant at the 10%, 5% and 1% level, respectively.

Table 3: Point estimates and 95% confidence intervals for measures of stock price persistence

(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Country	Sample period	$\hat{\alpha}_{OLS}$	Grid-bootstrap 95% CI	Subsampling equal-tailed 95% CI	Subsampling symmetric 95% CI	HL_{IRF}	Grid-bootstrap 95% CI
Australia	1958 Q 1–2008 Q 1	0.927	[0.910, 1.017]	[0.924, 1.028]	[0.851, 1.023]	2.889	[2.027, ∞]
Austria	1957 Q 1–2008 Q 1	0.950	[0.945, 1.013]	[0.939, 1.015]	[0.890, 1.007]	4.699	[3.299, ∞]
Canada	1956 Q 1–2008 Q 1	0.895	[0.885, 1.015]	[0.895, 0.979]	[0.857, 0.979]	2.018	[1.426, ∞]
France	1955 Q 1–2008 Q 1	0.969	[0.960, 1.017]	[0.969, 1.016]	[0.929, 1.015]	6.463	[4.296, ∞]
Germany	1960 Q 1–2008 Q 1	0.945	[0.939, 1.022]	[0.931, 1.008]	[0.918, 1.003]	3.968	[2.739, ∞]
India	1957 Q 1–2008 Q 1	0.980	[0.974, 1.019]	[0.983, 1.033]	[0.933, 1.025]	9.755	[7.294, ∞]
Ireland	1955 Q 1–2008 Q 1	0.916	[0.896, 1.017]	[0.898, 1.022]	[0.851, 1.005]	2.167	[1.670, ∞]
Italy	1957 Q 1–2008 Q 1	0.976	[0.960, 1.013]	[0.980, 1.017]	[0.929, 1.016]	6.330	[4.272, ∞]
Japan	1959 Q 1–2008 Q 1	0.986	[0.981, 1.017]	[0.987, 1.032]	[0.948, 1.024]	12.439	[7.996, ∞]
South Korea	1981 Q 1–2008 Q 1	0.939	[0.912, 1.028]	[0.954, 1.041]	[0.863, 1.014]	3.594	[2.514, ∞]
Netherlands	1957 Q 1–2008 Q 1	0.969	[0.968, 1.019]	[0.962, 1.013]	[0.941, 1.013]	7.876	[5.064, ∞]
New Zealand	1967 Q 1–2008 Q 1	0.963	[0.945, 1.023]	[0.953, 1.021]	[0.911, 1.014]	5.100	[3.461, ∞]
Switzerland	1950 Q 1–2008 Q 1	0.967	[0.952, 1.015]	[0.940, 1.070]	[0.826, 1.043]	5.956	[3.899, ∞]
Swiss	1955 Q 1–2008 Q 1	0.956	[0.937, 1.011]	[0.947, 1.000]	[0.915, 0.997]	4.250	[3.058, ∞]
UK	1958 Q 1–2008 Q 1	0.952	[0.939, 1.020]	[0.944, 1.043]	[0.890, 1.027]	4.324	[2.897, ∞]
USA	1957 Q 1–2008 Q 1	0.970	[0.957, 1.016]	[0.967, 1.027]	[0.919, 1.022]	6.151	[4.111, ∞]

⁽¹⁾ Lag length k in equation (2) selected using the Ng and Perron (2001) modified AIC.

⁽²⁾ OLS estimate for the sum of the AR coefficients in equation (2).

⁽³⁾ Estimate of the half-life (measured in years) based on the impulse response function.