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Some evidence of random walk behavior of Euro exchange rates using ranks and signs

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Abstract

This study utilises recently developed tests based on ranks and signs, in addition to the traditional variance ratio test, to examine the behavior of Euro exchange rates. We show that adjustments for multiple tests must be employed in order to avoid size distortions. Overall, such adjustments provide evidence consistent with random walk behavior of Euro exchange rates.

JEL classification: G12; G14; G15
Keywords: Euro exchange rates; Variance ratio; Ranks; Signs; Adjustment for multiplicity

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1. Introduction

Past evidence suggests that nominal exchange rate series follow a random walk process (see Meese and Singleton, 1982; Baillie and Bollerslev, 1989; Giddy and

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Dufey, 1975; Hsieh, 1988; among others). This implies, therefore, that the behavior of nominal exchange rates are weak form efficient, and hence not predictable (Fama, 1970, 1991). Nevertheless, a recent contribution by Jamaleh (2002) suggests that economic fundamentals effectively drive the dynamics of the Euro/Dollar exchange rate. This suggestion implies that the notion of random behavior of Euro/Dollar exchange rate return series is rejectable due to the impact of economic factors.

The examination of the behavior of asset returns, in the context of a weak form efficient market, is of interest to not only academics, but also, practitioners and regulators. While academicians seek to understand the behavior of asset returns over time, practitioners and investors are often interested in identifying market inefficiencies that produce exploitable patterns in exchange rate returns. Regulators in contrast, are interested in improving the informational efficiency of the securities market in which exchange rates are traded. Knowledge of the behavior of exchange rates, particularly in relation to efficiency/randomness issues, are therefore of considerable interest to a large number of interest groups.

The present work uses the alternative variance ratio tests based on ranks and signs proposed by Wright (2000), in addition to the tests suggested by Lo and MacKinlay (1989), to examine the behavior of some Euro exchange rates. On January 1, 1999, the Euro became the currency of 11 European member states namely Austria, Belgium, France, Finland, Germany, Ireland, Italy, Luxembourg, Portugal and Spain. On this date, the national currencies of the member countries became non-decimal subunits of the Euro and conversion rates between each of them and the Euro became irrevocably fixed. National banknotes and coins of old currencies were finally withdrawn from use (end of dual circulation) and replaced by Euros from February 28, 2002.

The scope of this paper is limited to a preliminary test of aggregate market transactions of the Euro. As Lo (1997) claims, a meaningful test for the efficient market hypothesis should specify additional elements, such as information structure and investors/'traders' preferences. [2] In any case, this is a timely addition to the extant literature of a new financial product, the Euro from its birth. Another contribution of this paper is that, to the best knowledge of the authors, this study will be the first of its kind on the Euro. Also, the large volume of studies that have examined the behavior of asset returns are based on parametric tests, whereas this paper reports some findings based on a new non-parametric methodology. As pointed out by Wright (2000), the non-parametric based tests will be robust to many forms of conditional heteroskedasticity and ought to have power against a wide range of models of serial correlation such as autoregressive moving average and fractionally integrated alternatives with heavy tailed innovations. Finally, we apply p-value adjustments for multiple tests in order to avoid size distortions due to sequential testing.

The remainder of the paper is organized as follows. Section 2 discusses some previous research on the issue. The research methodology, some Monte Carlo

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[2] We are grateful to an anonymous referee for pointing this out to us.

simulations and the empirical results are discussed in Section 3. Section 4 provides a summary of the main findings and some concluding remarks.

2. Previous studies

A number of studies have tested the hypothesis that exchange rate series follow a random walk behavior. Among such studies are Meese and Singleton (1982) and Baillie and Bollerslev (1989) who report a unit root component in the exchange rate series and Giddy and Dufey (1975), Cornell and Dietrich (1978), Logue et al. (1988) and Hsieh (1988) who suggest that exchange rate series contain uncorrelated increments.

Lo and MacKinlay (1988) criticized the traditional random walk tests of asset returns and introduced a more robust volatility-based specification test. Since most asset returns often possess time-varying volatilities and deviations from normality, the importance of developing a test which is robust to heteroskedasticity and non-normality becomes important. Lo and MacKinlay developed tests based on assumptions of both homoskedasticity and heteroskedasticity which they applied to examine the validity of the random walk hypothesis for weekly stock returns. Other studies have applied the methodology to test the random walk hypothesis in world money rates (e.g. Liu and He, 1991; Chou et al., 1996; Pan et al., 1996; Fong et al., 1997).

Liu and He (1991) applied variance ratio tests based on Lo and MacKinlay (1988) and provided evidence that rejected the random walk hypothesis for five pairs of nominal exchange rates. Their reported results suggest that autocorrelations are present in weekly increments in nominal exchange rate series. Ayadi and Pyun (1994) applied the variance ratio test to the daily Korean stock market prices. Their findings rejected the random walk hypothesis under the assumption of homoskedasticity. However, when heteroskedasticity is assumed, their findings supported the random walk hypothesis for daily data and longer horizons of data intervals. Chou et al. (1996) use variance ratio to test the random walk hypothesis of interest rates for eight world currencies. Their results indicate that most of the interest rates in the countries examined do not follow a random walk in the short run and vice versa for the long run.

3. Data and methodology

The behavior of the return series of the Euro-based exchange rates is examined by first applying the traditional variance ratio test, and then, by application of the non-parametric tests suggested by Campbell and Dufour (1997), Wright (2000) and Luger (2003).

The data for the study consists of the daily nominal exchange rates for the Australian dollar, Canadian dollar, New Zealand dollar, Japanese yen, British pound, Norwegian kroner, Singapore dollar, Swedish krona, Swiss franc and United States dollar, all relative to the Euro from 5th January 1999 to 11 November 2002 (almost 4 years of data). The number of daily observations totaled 1005. The exchange rate data for the series examined were obtained from Datastream/Primark. The short historical data available for the Euro is of course likely to have some impact on the results.

The variance ratio test proposed by Lo and MacKinlay (1988, 1989) is based on the fact that, for a random walk series, the variance of its kth difference is k times the variance of its first difference. The M_1 test is valid under the assumption of independent and identically distributed returns, whereas the M_2 is robust against heteroskedasticity.

Table 1 shows the results of Lo and MacKinlay's variance ratio tests on the logarithm of the daily Euro-based nominal exchange rates, for $k = 2$, 5, 10 and 30.

Results for individual k values suggest that the random walk assumption is violated in the Canada, Norway, Singapore and Switzerland cases, using M_1 test. However, using the M_2 test, the null is rejected just for Canada, Norway and Singapore exchange rates.

Table 1
Lo and MacKinlay's VR tests results

	Australia	Canada	Japan	UK	US
M_1					
$k = 2$	0.716	−2.296*	−0.824	−1.049	−0.415
$k = 5$	−0.487	−2.318*	−0.935	−1.296	−1.355
$k = 10$	−1.060	−1.902	−0.814	−1.309	−0.979
$k = 30$	−0.806	−1.228	−0.670	−1.220	−0.333
M_2					
$k = 2$	0.623	−1.877	−0.753	−0.863	−0.390
$k = 5$	−0.454	−2.028*	−0.807	−1.102	−1.276
$k = 10$	−0.994	−1.734	−0.706	−1.113	−0.922
$k = 30$	−0.773	−1.177	−0.605	−1.053	−0.314
	New Zealand	Norway	Singapore	Sweden	Switzerland
M_1					
$k = 2$	−0.404	0.725	−2.558*	1.852	0.293
$k = 5$	−1.277	−1.907	−2.877**	−0.017	−1.223
$k = 10$	−1.178	−1.960*	−2.217*	−0.803	−1.682
$k = 30$	−1.520	−2.059*	−1.293	−1.548	−2.267*
M_2					
$k = 2$	−0.373	0.649	−2.484*	1.691	0.168
$k = 5$	−1.232	−1.744	−2.778**	−0.015	−0.735
$k = 10$	−1.126	−1.799	−2.117*	−0.703	−1.061
$k = 30$	−1.463	−1.993*	−1.233	−1.396	−1.633

The entries are Lo and MacKinlay's VR tests at aggregation interval k. * indicates significance at (individual) 5% level, whereas ** indicates significance at 5% level using the SMM(α;m;∞) asymptotic distribution, $\alpha = 0.05$ and $m = 4$.

However, the application of VR tests for multiple k values leads to overrejection of the null hypothesis, above the nominal size. Therefore, as suggested by Chow and Denning (1993), we compare the maximum M_1 and M_2 statistics (in absolute value) with the asymptotic α-point critical value of the studentized maximum modulus, $\mathrm{SMM}(\alpha; m; \infty)$, where m is the number of k values. [3] The results allow us to reject the null just for the Euro/Singapore nominal exchange rate, using both M_1 and M_2 tests.

3.1. Ranks and signs-based random walk tests

In a recent paper, Wright (2000) proposes the use of signs and ranks of differences in place of the differences in the Lo and MacKinlay tests. Wright demonstrates that his non-parametric variance ratio tests based on ranks (R_1 and R_2) and signs (S_1 and S_2), can be more powerful than the tests suggested by Lo and MacKinlay. They have high power against a wide range of models displaying serial correlation, including fractionally integrated alternatives. The tests based on ranks are exact under the independence and identical distribution assumption, whereas the tests based on signs are exact even under conditional heteroskedasticity. Moreover, Wright (2000) shows that ranks-based tests display low size distortion, under conditional heteroskedasticity.

Given T observations of asset returns $\{y_1, \ldots, y_T\}$, Wright's proposed R_1 and R_2 are defined as:

$$R_1 = \left(\frac{\frac{1}{Tk}\sum_{t=k}^{T}(r_{1t} + \cdots + r_{1t-k+1})^2}{\frac{1}{T}\sum_{t=1}^{T} r_{1t}^2} - 1 \right) \times \phi(k)^{-1/2}, \tag{1}$$

$$R_2 = \left(\frac{\frac{1}{Tk}\sum_{t=k}^{T}(r_{2t} + \cdots + r_{2t-k+1})^2}{\frac{1}{T}\sum_{t=1}^{T} r_{2t}^2} - 1 \right) \times \phi(k)^{-1/2}, \tag{2}$$

where

$$\begin{aligned} r_{1t} &= \left(r(y_t) - \frac{T+1}{2} \right) \Big/ \sqrt{\frac{(T-1)(T+1)}{12}}, \\ r_{2t} &= \Phi^{-1}(r(y_t)/(T+1)), \\ \phi(k) &= \frac{2(2k-1)(k-1)}{3kT}, \end{aligned}$$

$r(y_t)$ is the rank of y_t among $y_1, \ldots, y_T$, and Φ^{-1} is the inverse of the standard normal cumulative distribution function. The tests based on the signs of returns are given by:

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[3] For a number of k values equal to four, the asymptotic critical value is 2.491, at the 5% significance level.

$$S_1 = \left(\frac{\frac{1}{Tk}\sum_{t=k}^{T}(s_t + \cdots + s_{t-k+1})^2}{\frac{1}{T}\sum_{t=1}^{T}s_t^2} - 1 \right) \times \phi(k)^{-1/2}, \tag{3}$$

$$S_2 = \left(\frac{\frac{1}{Tk}\sum_{t=k}^{T}(s_t(\overline{\mu}) + \cdots + s_{t-k+1}(\overline{\mu}))^2}{\frac{1}{T}\sum_{t=1}^{T}s_t(\overline{\mu})^2} - 1 \right) \times \phi(k)^{-1/2}, \tag{4}$$

where $s_t = 2u(y_t, 0)$, $s_t(\overline{\mu}) = 2u(y_t, \overline{\mu})$, and

$$u(x_t, q) = \begin{cases} 0.5 & \text{if } x_t > q, \\ -0.5 & \text{otherwise.} \end{cases}$$

Thus, S_1 assumes a zero drift value. If the value of the drift parameter is unknown, the procedure described in Luger (2003), based on Campbell and Dufour (1997), is applied to compute S_2. This method consists of a two-step strategy.

First, an exact $1-\alpha_1$ level confidence interval for the drift parameter μ, valid under the null hypothesis, is established. The second step consists of computing the S_2 statistic, for each candidate value b for the drift parameter in the confidence interval. The value of the S_2 statistic (retaining the sign) at aggregation interval k is then defined as

$$S_2(k) = \inf\{|S_2(k, b)| : b \in CI_\mu(\alpha_1)\},$$

where $S_2(k,b)$ is computed by defining $s_t(b) = 2u(y_t, b)$. The chosen S_2 value is compared to the appropriate critical values for an α_2 level test, such that the overall level of the strategy is bounded by $\alpha = \alpha_1 + \alpha_2$. In this paper, we have set $\alpha_1 = 0.01$ and $\alpha_2 = 0.04$.

The results for individual k values reported in Table 2 suggest that the null hypothesis of random walk can be rejected for all nominal exchange rates, with the exception of Australia, Japan and UK. However, as pointed out by Wright (2000), using several k values would lead to an overrejection of the null hypothesis, as in Lo and MacKinlay's tests context.

This phenomenon is illustrated through the following experiment. 1,000 samples of three alternative models with zero population autocorrelation have been generated:

1. Model 1.

   $$x_t \sim N(0, 1).$$

2. Model 2.

   $$x_t \sim t_3.$$

Table 2
Wright's VR tests results

	Australia	Canada	Japan	UK	US
R_1					
$k = 2$	0.516	−2.317*	0.529	−0.348	−0.565
$k = 5$	−0.585	−2.940*	−0.483	−1.146	−2.183*
$k = 10$	−1.353	−2.172*	−0.572	−0.831	−1.598
$k = 30$	−1.233	−1.321	−0.425	−0.819	−0.607
R_2					
$k = 2$	0.658	−2.322*	−0.293	−0.874	−0.574
$k = 5$	−0.645	−2.674*	−0.857	−1.449	−1.845
$k = 10$	−1.322	−2.115*	−0.852	−1.362	−1.329
$k = 30$	−1.138	−1.399	−0.646	−1.366	−0.595
S_1					
$k = 2$	−0.252	−1.389	1.641	−0.063	−0.379
$k = 5$	−0.588	−2.800*	0.449	−1.026	−1.095
$k = 10$	−1.234	−1.918	0.991	0.067	−0.168
$k = 30$	−1.789	−1.321	0.912	0.476	1.071
S_2					
$k = 2$	0.000	−1.389	1.262	−0.063	−0.126
$k = 5$	−0.380	−2.362*	−0.012	0.035	−1.602
$k = 10$	−0.964	−0.841	−0.004	0.059	−1.006
$k = 30$	−0.246	0.035	−0.004	−0.139	0.053

	New Zealand	Norway	Singapore	Sweden	Switzerland
R_1					
$k = 2$	−0.963	0.517	−3.276*	2.027*	0.117
$k = 5$	−1.880	−2.142*	−3.382*	−0.246	−1.604
$k = 10$	−2.302*	−2.299*	−2.761*	−0.976	−1.643
$k = 30$	−2.359*	−2.128*	−1.769	−1.591	−2.478*
R_2					
$k = 2$	−0.660	0.445	−2.812*	1.870	0.308
$k = 5$	−1.607	−2.190*	−3.085*	−0.075	−1.433
$k = 10$	−1.818	−2.269*	−2.466*	−0.842	−1.700
$k = 30$	−2.090*	−2.327*	−1.651	−1.677	−2.506*
S_1					
$k = 2$	−0.505	2.209*	−2.777*	1.831	0.821
$k = 5$	−1.371	0.012	−2.731*	0.357	−1.026
$k = 10$	−1.477	−0.494	−2.157*	−0.254	−0.979
$k = 30$	−0.856	−0.054	−0.610	−0.755	−1.547
S_2					
$k = 2$	−0.379	0.316	−2.399*	1.326	0.316
$k = 5$	−1.348	−1.095	−2.039*	−0.012	−0.104
$k = 10$	−1.477	−1.443	−1.282	0.015	−0.153
$k = 30$	−0.856	−0.816	−0.459	0.037	−0.915

The entries are Wright's VR tests at aggregation interval k. * indicates significance at (individual) 5% level.

3. Model 3.

$$h_t = 0.95h_{t-1} + \xi_t,$$
$$x_t = \exp(0.5h_t)\varepsilon_t,$$

where $\varepsilon_t \sim N(0,1)$ and $\xi_t \sim N(0,0.1)$ are independent random variables, and Model 3 is a stochastic volatility model.

For each artificial sample, Wright's VR ranks and signs tests are computed for two k values (2 and 5) when $T = 100$, three k values (2, 5 and 10) when $T = 500$ and four k values (2, 5, 10 and 30) when $T = 1000$, and we analyse the impact on the empirical size of each individual test. The null hypothesis of random walk is rejected by test j if, for some k value, the statistic is larger or lower than the 97.5% or 2.5% percentile of the corresponding tabulated distribution, respectively. [4]

Results in Table 3 confirm the argument that using VR tests at various aggregation intervals leads to rejection rates larger than the nominal size. [5]

In order to avoid this perverse effect, we apply p-value adjustments for multiplicity, in line with Psaradakis (2000).

We compute the Sidack-adjusted p-value for each test j as:

$$\tilde{p}_{ji}^{(S)} = 1 - (1 - p_{ji})^m,$$

$i = 1,\ldots,m$, where p_{ji} is the p-value corresponding to the VR test j computed for an individual k value, and m is number of k values. Given a significance level α, the decision rule states that, using the VR test j, the null is rejected if $\tilde{p}_j^{(S)} = \min_{1\leqslant i\leqslant m}\tilde{p}_{ji}^{(S)} \leqslant \alpha$. The results in Table 4 have been produced using Monte Carlo simulations, following Wright's procedures to compute the empirical quantiles, using 100,000 replications in order to compute the empirical p-values for each VR test.

The results show that the suggested p-value adjustment procedure leads to an undersized testing strategy, most probably at the cost of a loss of power, with the exception of R_1 and R_2 tests against Model 3. In fact, despite the satisfactory results with R_1 test against Model 3, results concerning both R_1 and R_2 tests are not reliable when the time series are conditionally heteroskedastic. On the other hand, S_2 test is clearly undersized in all cases (with and without Sidack-type correction), well beyond the nominal level, since the two-step procedure always leads to a conservative test.

This method, however, assumes that the test statistics computed at different intervals are uncorrelated. In order to take into account possible correlations among the statistics, bootstrap-adjusted p-values can be computed as described in Westfall and Young (1993) and Psaradakis (2000). The goal of the procedure is to obtain an approximation to the null sampling distribution of $\min_{1\leqslant i\leqslant m} p_{ji}$, by resampling with

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[4] Critical values can be found in Wright (2000, Table 1, p. 3). Critical values for the S_2 test at $\alpha_2 = 4\%$ level have been simulated through 100,000 replications.

[5] Note that false rejections notably increase for R_1 and R_2 tests against Model 3, since those tests are not robust in the presence of heteroskedasticity.

Table 3
Wright's individual VR tests size

	R_1	R_2	S_1	S_2
$T = 100$				
Model 1	0.071	0.073	0.080	0.002
Model 2	0.080	0.078	0.088	0.000
Model 3	0.092	0.131	0.086	0.002
$T = 500$				
Model 1	0.108	0.107	0.108	0.017
Model 2	0.093	0.111	0.110	0.007
Model 3	0.129	0.183	0.102	0.013
$T = 1000$				
Model 1	0.135	0.138	0.128	0.014
Model 2	0.118	0.133	0.128	0.019
Model 3	0.160	0.236	0.139	0.018

The entries are Wright's VR tests rejection rates at (individual) 5% level.

Table 4
Wright's individual VR tests Sidack-corrected size

	R_1	R_2	S_1	S_2
$T = 100$				
Model 1	0.031	0.034	0.024	0.000
Model 2	0.034	0.038	0.038	0.000
Model 3	0.056	0.069	0.031	0.000
$T = 500$				
Model 1	0.037	0.036	0.046	0.001
Model 2	0.027	0.031	0.040	0.002
Model 3	0.054	0.082	0.040	0.001
$T = 1000$				
Model 1	0.041	0.043	0.038	0.002
Model 2	0.035	0.035	0.035	0.000
Model 3	0.051	0.082	0.037	0.002

The entries are Wright's VR tests rejection rates at (individual) 5% level.

replacement from the original returns. Table 5 displays size results, using $N = 200$ bootstrap samples.

These new results deserve special attention. Since the empirical distributions have been constructed under the null hypothesis of *iid* returns, which is clearly violated by Model 3, now the VR tests which are sensitive to deviations from this assumption (R_1 and R_2) are oversized. Nevertheless, the S_2 test, which was clearly undersized, now displays an empirical size closer to the 5% nominal level, mostly at larger sample sizes. This is explained by the fact that the uncorrelation of the original series is equivalent to the *iid* behavior of the *signs* of the data, regardless of the (absolute) *size* of the returns. Therefore, the bootstrap approach for S_1 and S_2 does not need to rep-

Table 5
Wright's individual VR tests bootstrap−corrected size

	R_1	R_2	S_1	S_2
$T = 100$				
Model 1	0.045	0.039	0.038	0.028
Model 2	0.054	0.052	0.040	0.045
Model 3	0.045	0.074	0.036	0.048
$T = 500$				
Model 1	0.034	0.037	0.032	0.029
Model 2	0.052	0.051	0.039	0.034
Model 3	0.074	0.113	0.047	0.052
$T = 1000$				
Model 1	0.048	0.041	0.038	0.045
Model 2	0.054	0.051	0.050	0.051
Model 3	0.067	0.125	0.043	0.051

The entries are Wright's VR tests rejection rates at (individual) 5% level.

Table 6
Wright's VR tests Sidack-corrected p-values

Currency	$\tilde{p}_{R_1}^{(S)}$	$\tilde{p}_{R_2}^{(S)}$	$\tilde{p}_{S_1}^{(S)}$	$\tilde{p}_{S_2}^{(S)}$
Australia	0.630	0.656	0.218	0.853
Canada	0.010**	0.026**	0.014**	0.058*
Japan	0.961	0.915	0.341	0.597
UK	0.755	0.528	0.809	0.999
US	0.117	0.262	0.639	0.385
New Zealand	0.034**	0.114	0.472	0.472
Norway	0.075*	0.038**	0.103	0.497
Singapore	0.002**	0.006**	0.017**	0.075*
Sweden	0.136	0.194	0.241	0.552
Switzerland	0.017**	0.015**	0.404	0.908

The entries are Sidack-adjusted p-values of Wright's VR tests, corresponding to four k values. * and ** indicate significance at 10% and 5% level, respectively.

licate the empirical characteristics of the conditional variance since it suffices to simply bootstrap the exchange rate returns.

Table 6 displays the Sidack-corrected p-values for each test. For instance, using just the S_1 test without p-value adjustment, the null hypothesis could be rejected for Canada, Norway and Singapore, at different significance levels; however, using the tests robust against heteroskedasticity (S_1 and S_2), with Sidack-adjusted p-values, the null could be clearly rejected at the 5% significance level just for Canada and Singapore.[6] Hence, for 8 out of 10 currencies the null hypothesis could not be rejected.

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[6] Note that the R_1 and R_2 tests allow us to reject the null in a larger number of cases, but we prefer to prevent our conclusions from false rejections due to conditional heteroskedasticity.

Table 7
Wright's VR tests bootstrap-corrected p-values

Currency	$\tilde{p}_{R_1}^{(N)}$	$\tilde{p}_{R_2}^{(N)}$	$\tilde{p}_{S_1}^{(N)}$	$\tilde{p}_{S_2}^{(N)}$
Australia	0.489	0.500	0.161	0.322
Canada	0.008**	0.024**	0.017**	0.008**
Japan	0.903	0.811	0.252	0.145
UK	0.582	0.386	0.756	0.973
US	0.087*	0.202	0.633	0.067*
New Zealand	0.024**	0.089*	0.409	0.101
Norway	0.063*	0.033**	0.244	0.096*
Singapore	0.001**	0.006**	0.020**	0.006**
Sweden	0.101	0.133	0.191	0.133
Switzerland	0.012**	0.009**	0.309	0.382

The entries are bootstrap-adjusted p-values of Wright's VR tests, under the null hypothesis of *iid* returns, corresponding to four k values and setting $N = 1000$ replications. * and ** indicate significance at 10% and 5% level, respectively.

Table 7 shows the bootstrap-adjusted p-values, setting $N = 1000$ bootstrap replications under the null hypothesis of *iid* returns. The results with S_2 provide stronger evidence against the weak form efficient market hypothesis for Canada and Singapore: the Sidack-adjusted p-values reject the null at the 10% level, whereas now the hypothesis can be rejected at the 1% level. Moreover, the null can be rejected in two additional cases, but at the 10% level (US and Norway). Following our conservative approach, we can claim that these results provide no evidence inconsistent with the contention that, with few exceptions, Euro exchange rates markets are weak form efficient.

In addition to Wright's tests, we have applied the exact and approximate versions of Campbell and Dufour (1997) and Luger (2003) tests. The inference drawn from these additional tests is consistent with the weak form efficiency of Euro exchange rates for the major trading countries in our study. [7]

4. Conclusions

This study utilises recently developed non-parametric based tests, in addition to the traditional variance ratio test to examine the behavior of daily Euro exchange rate returns for 10 countries.

The signs of the traditional variance ratio test suggest a negative dependence in most of the exchange rate return series examined, providing a strong rejection of weak form market efficiency for Euro exchange rates for the Canadian and Singapore dollar, and a weak rejection for the Swiss franc and Norwegian kroner exchange rates. However, following Chow and Denning (1993), adjusted critical values must be used in order to prevent size distortions, due to sequential testing at

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[7] In the interest of space, the results are not reported in this paper. However, they are available from the authors upon request.

various intervals. Using such critical values, the null hypothesis can be rejected in just the Canadian dollar case.

Wright's (2000) signs and ranks based variance ratio tests suggest the rejection of the null hypothesis for the Canadian, Singapore and New Zealand dollar, and for the Swiss franc and Norwegian kroner exchange rates. However, we show that, similarly to the context of the traditional variance ratio tests, the empirical size of sequential testing is larger than the nominal level. Then, two different kind of *p*-value adjustments are applied: the Sidack correction, and the bootstrap correction. In the latter case, empirical size is very close to the nominal level, although its validity in the context of conditional heteroskedasticity is confined to the signs based tests. Empirical evidence using adjustments for multiplicity shows that the null hypothesis can be clearly rejected just in two cases: the Canadian and Singapore dollar.

Comparison of the results of our study produced from various tests demonstrate that there is potential for considerable variation when examining the weak form efficiency of exchange rates series across countries. While our results suggest that the behavior of Euro exchange rates for the major trading currencies is weak form efficient, this may not necessarily be the case for non-major trading currencies and more studies in the area will be useful.

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