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TESTING FOR RANDOM WALK BEHAVIOR IN EURO EXCHANGE RATES

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ABSTRACT. This study examines the random walk behavior of major Euro exchange rates. The hypothesis is tested with new variance ratio tests based on power transformation and multiple ranks from daily and weekly data. We find that Euro exchange rates for the major trading countries follow the random walk hypothesis, and therefore are significantly weak-form efficient. This outcome is not necessarily the case for non-major trading currencies, especially for the Swedish kroner, where the random walk hypothesis is rejected at the daily and weekly frequencies.

JEL Classification: G14; G15; C14.
Keywords: Exchange Market Efficiency; Euro Exchange Rates; Random Walk; Variance Ratio Test.

RÉSUMÉ. Cet article examine l'hypothèse selon laquelle les principaux taux de change de l'euro évoluent selon un processus de marche aléatoire. À cette fin, nous appliquons de nouveaux tests du rapport de variances, basés sur une transformation puissance et des rangs multiples, sur des données quotidiennes et hebdomadaires. Les résultats obtenus pour les principaux partenaires commerciaux montrent que les taux de change de l'euro vérifient l'hypothèse de marche aléatoire, confirmant l'efficience au sens faible des marchés des changes. Ceci n'est pas forcément vérifié pour les devises des petits pays, notamment la couronne suédoise, pour laquelle l'hypothèse de marche aléatoire est rejetée tant à fréquence quotidienne qu'hebdomadaire.

Classification *JEL:* G14 ; G15 ; C14.
Mots-clefs : Efficience du marché des changes ; taux de change de l'euro ; marche aléatoire ; test du rapport de variances.

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1. Introduction

On January 1, 1999, the Euro became the currency of eleven European Union member states, namely Austria, Belgium, France, Finland, Germany, Ireland, Italy, Luxembourg, the Netherlands, 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. Greece joined the Eurozone on January 1, 2001. On February 28, 2002, national banknotes and coins of old currencies were withdrawn from use, signaling the end of dual circulation, and were replaced by Euros. Slovenia was admitted to the Eurozone on January 1, 2007, Cyprus and Malta on January 1, 2008, and Slovakia on January 1, 2009.[2] A symbol of the strength of the Economic and Monetary Union, the Euro has become an important international currency. Indeed, since its introduction, the Euro has been the second most widely held international reserve currency after the US dollar. Recently, due to the weak dollar and the strong Euro, many countries have held or plan to hold more Euros than US dollars in their currency reserves. Since December 2006, the Euro has surpassed the US dollar in terms of combined value of cash in circulation. The possibility that the Euro will become the primary international reserve currency in the near future is now gaining wide attention, suggesting the importance of understanding the behavior of Euro-based exchange rates.[3]

In this paper we analyze the efficiency market hypothesis [EMH thereafter] of major Euro exchange rates. The EMH states that, at any time, prices fully reflect all available and relevant information. Therefore, given only past price and return data, the current price is the best predictor of the future price, and the price change or return is expected to be zero (Fama, 1970; 1991). This is the essence of the weak-form EMH, which implies a random walk, and which has been the most commonly tested hypothesis in the empirical literature.[4]

If the nominal exchange rate follows a random walk process, then the market is weak-form efficient, and therefore not predictable. This means that it is impossible for an exchange trader to generate excess returns over time through speculation. Alternately, if the nominal exchange rate is predictable, then the market is not weak-form efficient, which means that exchange traders can generate abnormal returns through speculation.

Since the seminal work of Lo and MacKinlay (1988, 1989) and Poterba and Summers (1988), the standard variance ratio [VR thereafter] test or its improved modifications have been used to test market efficiency. They have been used in many empirical applications

2. The following European Union countries plan to join the Eurozone: Estonia, Latvia, Lithuania, Poland, the Czech Republic and Hungary on January 1, 2012, and Bulgaria and Romania on January 1, 2014.

3. To some extent, Euro-based exchange rates could behave differently from dollar-based exchange rates due to the differences in principal policy objectives, institutional constraints, and implementation practices between the European Central Bank and the US Federal Reserve (see, e.g., Wang *et al.*, 2007).

4. Note that if the random walk hypothesis is based on the theory of efficiency, the EMH does not imply that prices follow a random walk. Therefore, if prices do not follow a random walk, this does not imply inefficiency of the market. See Lo and MacKinlay (2001) for a discussion on random walk hypothesis and efficiency market hypothesis.

on foreign exchange rates, by Liu and He (1991), Fong, Koh and Ouliaris (1997), Wright (2000), Yilmaz (2003), Lima and Tabak (2007), and Azad (2009), among others. However, all studies on the random walk hypothesis [RWH thereafter] examine foreign exchange rates against the US dollar. To the best of our knowledge, Belaire-Franch and Opong (2005), Al-Khazali and Koumanakos (2006), and Chen (2008) are the only studies that evaluate the RWH for Euro-based nominal exchange rates.

Therefore, we re-examine the random walk behavior of major Euro exchange rates in two ways. First, this study is based on an extensive sample. We study daily and weekly data for major Euro exchange rates over the 1999–2008 period. Consistent with Azad (2009), we analyze the weak-form from both data frequencies because a market can be considered to be perfectly weak-form efficient if it is found to behave randomly at any level of data frequency. We thus avoid the shortcomings of high and medium-low frequency data. Indeed, using high frequency data to estimate the market efficiency of developed markets allows us to take into account some of their characteristics, such as high trading volume and better information. However, it also implies biases for developing markets, such as non-trading and asynchronous prices, which are overcome with medium-low frequency data. Second, we apply new VR tests proposed by Chen and Deo (2006) and Belaire-Franch and Contreras (2004), which are more powerful than those applied in previous studies, to examine the behavior of major Euro exchange rates. More precisely, Chen and Deo (2006) suggested a power transformation of the VR statistic, which corrects a well-known problem with the VR test, namely that the VR statistic is biased and right skewed in finite samples. Wright (2000) proposes an alternative to conventional asymptotic VR tests using rank tests, which are exact tests whose sampling distributions do not entail asymptotic approximations. Belaire-Franch and Contreras (2004) developed a multiple version of Wright's (2000) rank tests that overcomes the problem of size distortions when applying individual VR tests for several holding periods.

The remainder of this paper is organized as follows: Section 2 presents a brief literature survey on exchange rate efficiency; Section 3 discusses the VR tests; Section 4 summarizes the characteristics of the data and reports the empirical results. The conclusion is drawn in Section 5.

2. Brief literature survey

A number of studies tested the hypothesis that exchange rate series exhibit a random walk behavior. For instance, Meese and Singleton (1982) and Baillie and Bollerslev (1989) report a unit root component in the exchange rate series, and Giddy and Dufey (1975), Cornell and Dietrich (1978) and Hsieh (1988) show, using correlation tests, that exchange rate series contain uncorrelated increments.

Table 1 - Selected studies on the EMH for exchange rate markets

Studies	Sample	Methodologies	Comments
Relative to US dollar			
Liu and He (1991)	1974–1989 (W)	Lo-MacKinlay VR	not EMH for CA, FR, GE, JP, UK.
Fong *et al.* (1997)	1974–1989 (W)	Richardson-Smith VR	not EMH for CA, FR, GE, JP, UK. not EMH for FR, GE, JP; EMH for CA, UK.
		Chow-Denning VR	
Wright (2000)	1974–1996 (W)	Wright VR	not EMH for CA, FR, GE, JP, UK.
		Chow-Denning VR	
Chang (2004)	1974–1998 (D)	Lo-MacKinlay VR (bootstrap)	not EMH for JP. inconclusive for CA, FR, GE, UK.
Ajayi and Karemera (1996)	1986–1991 (D, W)	Lo-MacKinlay VR	EMH for HK, IN, KO, PH, SI. not EMH for MA, TA, TH.
Lee *et al.* (2001)	1988–1995 (D)	Cecchetti-Lam VR (bootstrap)	EMH for KO; not EMH for AU, HK, MA, NZ, PH, SI, TA, TH.
Lima and Tabak (2007)	(D, W)	Lo-MacKinlay VR	EMH for AR, BR, IN, HK, MA, PH, MA, RU, SI, TH.
		Cecchetti-Lam VR (bootstrap)	
Azad (2009)	1998–2007 (D,W)	Wright VR	EMH for AU, IN, JP, NZ, PH, SI, KO. not EMH for CH, HK, MA, TA, TH.
Relative to Euro			
Belaire-Franch and Opong (2005)	1999–2002 (D)	multiple Wright VR (size adjustment)	EMH for AU, NZ, JP, UK, NO, SE, SW, US; not EMH for CA, SI.
Al-Khazali and Koumanakos (2006)	2000–2004 (D)	Wright VR	EMH for KU, EM; not EMH for BA, EG, JO, MO, OM, QA, SA, TU.
Chen (2008)	1999–2008 (D)	Lo-MacKinlay VR	EMH for US.
		Chow-Denning VR	
		Wright VR	

Notes:

i) D: daily; W: weekly; M: monthly. EMH: efficiency market hypothesis.
AU: Australian dollar; AR: Argentina peso; BA: Bahraini dinar; BR: Brazilian real; CA: Canadian dollar; CH: Chinese renminbi; EG: Egyptian pound; EM: Emirate dirham; FR: French franc; GE: German mark; HK: Hong Kong dollar; IN: Indonesian rupiah; IT: Italian lira; JO: Jordanian dinar; JP: Japanese yen; KO: Korean won; KU: Kuwaiti dinar; MA: Malaysian ringgit; MO: Moroccan dirham; NO: Norwegian kroner; NZ: New Zealand dollar; OM: Omani rial; PH: Philippine peso; QA: Qatari rial; RU: Russian rouble; SA: Saudi rial; SI: Singaporean dollar; SE: Swedish krona; SW: Swiss franc; TA: Taiwanese dollar; TH: Thai baht; TU: Tunisian dinar; UK: British pound. US: United States dollar.

ii) The results displayed for Azad (2009) and Ajayi and Karemera (1996) are based on daily data.

Since the seminal work of Lo and MacKinlay (1988, 1989) and Poterba and Summers (1988), the standard VR test or its improved modifications have been used to test the random walk hypothesis for nominal exchange rates (see Table 1).[5] These tests are robust to heteroskedasticity and non-normality, which are present in exchange rate series.[6] Furthermore, the VR tests have important economic implications. As suggested by Lo and MacKinlay (1989) and Liu and He (1991), use of the VR provides a convenient way to differentiate between the overshooting and undershooting phenomena in exchange rates.

Liu and He (1991) applied VR tests based on Lo and MacKinlay (1988) and provided evidence that the RWH was rejected for five major foreign exchange rates (Canadian dollar, French franc, German mark, Japanese yen and British pound). Their results suggested that autocorrelations are present in weekly increments in nominal exchange rate series. Fong, Koh and Ouliaris (1997), Wright (2000), Yilmaz (2003) and Chang (2004) re-examined the same five exchange rates using various VR tests. Wright (2000) and Yilmaz (2003)[7] applied non-parametric sign and rank-based VR tests and multiple Chow-Denning (1993) and Richardson-Smith (1991) VR tests and confirmed the results obtained by Liu and He (1991). Fong, Koh and Ouliaris (1997) found that the Richardson-Smith test failed to reject RWH for all five currencies, whereas the Chow-Denning test continued to reject the hypothesis for the French franc, German mark and Japanese yen.[8] Chang (2004) provides evidence rejecting the RWH for the Japanese yen, while the results for the other four currencies were inconclusive when employing the Lo-MacKinlay VR test with a bootstrap resampling technique and daily data.

Ajayi and Karemera (1996), Lee, Pan and Liu (2001), Lima and Tabak (2007) and Azad (2009) analyzed foreign exchange rates of Asian countries (Hong Kong dollar, Indonesian rupiah, Korean won, Malaysian ringgit, Philippine peso, Singaporean dollar, Taiwanese dollar and Thai baht).[9] Ajayi and Karamera (1996) rejected the RWH in the majority of markets employing the Lo-MacKinlay VR test. Lee, Pan and Liu (2001) applied the multiple Cecchetti-Lam (1994) VR tests with bootstrap method and found little evidence of serial correlations for all the currencies, with the exception of the Korean won. Using the Lo-MacKinlay and multiple Cecchetti-Lam (1994) VR tests with a bootstrap technique, Lima

5. See Azad (Appendix A, 2009) for a survey of the different methodologies used to explain random walk behavior of exchange rate markets.

6. Lo and MacKinlay (1989) examined the VR, Dickey-Fuller unit root and Box-Pierce serial correlation tests and found that the VR test was more powerful than other tests for the heteroskedastic random walk. In other words, when the focus is the absence of correlation among the increments, the VR test is preferred. See Hoque *et al.* (2007) and Charles and Darné (2009) for a review of the VR tests.

7. Yilmaz (2003) also examined daily data for the Swiss franc and Italian lira.

8. The difference between the results obtained by Fong *et al.* (1997) and Yilmaz (2003), even if they employed the same VR tests, can be explained by the fact that Fong *et al.* (1997) investigated weekly data between October 1979, and March 1989, whereas Yilmaz (2003) studied daily data between January 2, 1974, and February 12, 2001.

9. Lee *et al.* (2001) and Azad (2009) also examined the Australian dollar and New Zealand dollar; Azad (2009) analyzed the Chinese renminbi and Japanese yen. Lima and Tabak (2007) studied the Brazilian real, Argentine peso and Russian ruble. See Table 1.

and Tabak (2007) found that the RWH cannot be rejected for the exchange rates of these emerging markets. Azad (2009) failed to reject the RWH in most of the currencies using the Wright (2000) tests and daily data, except for the Hong Kong dollar and Malaysian ringgit, and found that the Taiwanese dollar and Thai baht are efficient in the short-term but not with long holding periods. However, the RWH was rejected only for the Malaysian ringgit when employing weekly data.

Anthony and MacDonald (1999) reported mixed evidence refuting the RWH for many daily bilateral exchange rates (Belgian-Luxembourg franc, Danish krone, German mark, French franc, Italian lira, Irish punt and Dutch guilder) in the European Monetary System using the Lo-MacKinlay VR test.

Belaire-Franch and Opong (2005) used nonparametric ranks and sign-based VR tests, suggested by Wright (2000), with size adjustment for multiple tests to examine the RWH of ten daily Euro-based nominal exchange rates (Australian dollar, Canadian dollar, New Zealand dollar, Japanese yen, British pound, Norwegian kroner, Singapore dollar, Swedish krona, Swiss franc and United States dollar). Their results indicated that the behavior of Euro exchange rates for the major trading currencies is weak-form efficient. Chen (2008) also found that the US dollar exchange rate market is weak-form efficient using the Lo-MacKinlay, Chow-Denning and Wright VR tests. Finally, Al-Khazali and Koumanakos (2006) employed the Wright VR tests to examine the RWH of the Euro exchange rates for ten Middle Eastern and North African (MENA) currencies (Bahraini dinar, Egyptian pound, Emirate dirham, Jordanian dinar, Kuwaiti dinar, Moroccan dirham, Omani rial, Qatari rial, Saudi rial and Tunisian dinar). They rejected the RWH for all currencies except for the Emirate dirham and Kuwaiti dinar.

3. The variance ratio tests

The VR methodology consists of testing the RWH against stationary alternatives by exploiting the fact that the variance of random walk increments is linear in all sampling intervals, i.e. the sample variance of k-period return (or k-period differences), $y_t - y_{t-k}$, of the time series y_t, is k times the sample variance of the one-period return (or the first-difference), $y_t - y_{t-1}$. The VR at lag k is then defined as the ratio between $(1/k)^{th}$ of the k-period return (or the k^{th} difference) to the variance of the one-period return (or the first-difference). Thus, for a random walk process, the variance computed at each individual lag interval k (k = 2, 3, ...) should be equal to unity.

In testing the null hypothesis of random walk, the VR test evaluates the hypothesis that a given time series or its first-difference (or return), $x_t = y_t - y_{t-1}$, is a collection of independent and identically distributed observations (i.i.d.) or that it follows a martingale difference sequence. We define the VR of k-period return as:

$$V(x) = \frac{Var(x_t + x_{t-1} + \ldots + x_{t-k+1})/k}{Var(x_t)}$$
$$V(x) = \frac{Var(y_t - y_{t-k})/k}{Var(y_t - y_{t-1})} \cong 1 + \sum_{i=1}^{k-1} \left(\frac{2(k-i)}{k} \right) \rho_i$$

where ρ_i is the i^{th} lag autocorrelation coefficient of $\{x_t\}$.

The VR tests consider statistics based on an estimator of $V(k)$. Given T observations of first-differences of a variable, $\{x_1, \ldots, x_T\}$. It is assumed that x_t is a realization of the underlying stochastic process X_t, which follows a martingale difference sequence. The VR statistic is defined as:

$$VR(x) = \frac{\hat{\sigma}^2(k)}{\hat{\sigma}^2(1)}$$

where $\hat{\sigma}^2(1)$ is the unbiased estimator of the one-period return variance, using the one-period returns x_t, and is defined as:

$$\hat{\sigma}^2(1) = (T-1)^{-1} \textstyle\sum_{t=1}^{T} (x_t - \hat{\mu})^2 = (T-1)^{-1} \sum_{t=1}^{T} (y_t - y_{t-1} - \hat{\mu})^2$$

with $\hat{\mu} = T^{-1} \sum_{t=1}^{T} x_t$ is the estimated mean. For the estimator of k-period return variance $\hat{\sigma}^2(k)$, using k-period returns $(x_t + \ldots + x_{t-k+1})$, there are many ways to do it. Due to limited sample size and the desire to improve the power of the test, this estimator is often performed using overlapping long-horizon returns (k-period), as advocated by Lo and MacKinlay (1988)[10], and it is defined as:

$$\hat{\sigma}^2(k) = m^{-1} \sum_{t=k}^{T} (x_t + x_{t-1} + \ldots + x_{t-k+1} - k\hat{\mu})^2 = m^{-1} \sum_{t=k}^{T} (y_t - y_{t-k} - k\hat{\mu})^2$$

where $m = k(T - k + 1)(1 - kT^{-1})$. The value of m is chosen such as $\hat{\sigma}^2(k)$ is an unbiased estimator of the k-period return variance when σ_t^2 is constant over time.

Following Wright (2000), the VR statistic can be written as:

$$VR(x;k) = \left\{ (Tk)^{-1} \sum_{t=k}^{T} (x_t + \ldots + x_{t-k+1} - k\hat{\mu})^2 \right\} \div \left\{ T^{-1} \sum_{t=1}^{T} (x_t - \hat{\mu})^2 \right\}$$

This is an estimator for the unknown population VR, denoted as $V(k)$, which is the ratio of $1/k$ times, the variance of the k-period differences to the variance of the first-difference. If the data-generating process of time series is a random walk, the expected value of $VR(x;k)$ should be equal to unity for all horizons k. If returns are positively (negatively) autocorrelated, the VR should be higher (lower) than unity. Time series is said to be mean-reverting if $VR(x;k)$ is significantly lower than unity at long horizons k.[11] However, time series are mean-averting, i.e. explosive, if $VR(x;k)$ is significantly higher than unity at long horizons.

10. Lo and MacKinlay (1988) and Campbell *et al*. (1997) argued that using overlapping data in estimating the variances allows to obtain a more efficient estimator and hence a more powerful test.

11. In the case of exchange rates, there are several alternative explanations for mean-reversion behavior including the overshooting phenomenon (Dornbusch, 1976) and the undershooting phenomenon (Frenkel and Rodriguez, 1982), and central bank intervention (Liu and He, 1991).

3.1. Lo and MacKinlay (1988) test

Lo and MacKinlay (1988) proposed the asymptotic distribution of $VR(x;k)$ by assuming that k is fixed when $T \to \infty$. Under the assumption of conditional heteroskedasticity, Lo and MacKinlay (1988) proposed the heteroskedasticity[12] robust test statistic $M(x;k)$ given by:

$$M(x;k) = \frac{VR(x;k) - 1}{\phi^*(k)^{1/2}}$$

which follows the standard normal distribution asymptotically under the null hypothesis that $V(k) = 1$, where:

$$\phi^*(k) = \sum_{j=1}^{k-1} \left[\frac{2(k-j)}{k}\right]^2 \delta(j)$$

$$\delta(j) = \left\{\sum_{t=j+1}^{T} (x_t - \hat{\mu})^2 (x_{t-j} - \hat{\mu})^2\right\} \div \left\{\left[\sum_{t=1}^{T} (x_t - \hat{\mu})^2\right]^2\right\}$$

3.2. Chen and Deo (2006) tests

It is critical to note that the Lo-MacKinlay test is an asymptotic test in that its sampling distribution is approximated by its limiting distribution. Indeed, the practical use of the statistic is impeded by the fact that the asymptotic theory provides a poor approximation to the small-sample distribution of the VR statistic. More specifically, rather than being normally distributed (when standardized by $\sqrt{T}$) as the theory states, the statistics are severely biased and right skewed for large k (relative to T) (Lo and MacKinlay, 1989), which makes application of the statistic problematic. Therefore, the finite-sample null distribution of the test statistic is quite asymmetric and non-normal. These finite sample deficiencies may give rise to serious size distortions or low power, which can cause misleading inferences. This is especially true when the sample size is not large enough to justify asymptotic approximations (Cecchetti and Lam, 1994).

To circumvent this problem, Chen and Deo (2006) suggested a simple power transformation of the VR statistic that, when k is not too large, provides a better approximation to the normal distribution in finite samples and is able to solve the well-known right skewness problem. They showed that the transformed VR statistic leads to significant gains in power against mean-reverting alternatives.[13] Furthermore, the distribution of the transformed VR statistic is shown, both theoretically and through simulations, to be robust to conditional heteroskedasticity.

First, they defined the VR statistic based on the periodogram as:

$$VR_p(k) = \frac{1}{(1 - k/T)} \frac{4\pi}{T\hat{\sigma}^2} \sum_{j=1}^{(T-1)/2} W_k(\lambda_j) I_{\Delta X}(\lambda_j)$$

12. Lo and MacKinlay (1988) also proposed a test statistic which is robust under homoskedasticity and follows the standard normal distribution asymptotically.

13. A solution is also provided in a series of theoretical papers such as those by Richardson and Stock (1989), Deo and Richardson (2003), Tse, Ng and Zhang (2004), Perron and Vodounou (2005).

where $I_{\Delta X}(\lambda_j) = (2\pi T)^{-1} \left| \sum_{t=1}^{T} (y_t - y_{t-1} - \hat{\mu}) \exp(-i\lambda_j t) \right|^2,$

and $W_k(\lambda_j) = \sum_{|j| \leq k} \left(1 - \frac{|j|}{k}\right) \exp(-ij\lambda) = k^{-1} \left\{ \frac{\sin\left(\frac{k\lambda}{2}\right)}{\sin\left(\frac{\lambda}{2}\right)} \right\}^2$

To obtain their transformed VR statistic, noted $VR_p^{\beta}(k)$, they applied the following power transformation[14] to $VR_p(k)$:

$$\beta = 1 - \frac{2}{3} \frac{\left(\sum_{j=1}^{[(T-1)/2]} W_k(\lambda_j)\right)\left(\sum_{j=1}^{[(T-1)/2]} W_k^3(\lambda_j)\right)}{\left(\sum_{j=1}^{[(T-1/2]} W_k^2(\lambda_j)\right)^2} \tag{1}$$

This power-transformed VR test is an individual test where the null hypothesis is tested for an individual value of k. To answer the question as to whether or not a time series is mean-reverting requires that the null hypothesis hold true for all values of k. Therefore, it is necessary to conduct a joint test where a multiple comparison of VRs is done over a set of different time horizons. However, conducting separate individual tests for a number of k values may be misleading because it leads to over-rejection of the null hypothesis of a joint test above the nominal size (Chow and Denning, 1993). Thus, Chen and Deo (2006) also proposed a joint VR test based on their individual power transformed VR statistic. They suggested the following Wald statistic:

$$QB(k) = (\mathbb{V}_{p,\beta}(k) - \mu_\beta)'\Sigma_\beta^{-1}(\mathbb{V}_{p,\beta}(k) - \mu_\beta)$$

where $\mathbb{V}_{p,\beta}$ a column vector sequence of VR statistics $\mathbb{V}_{p,\beta}(k) = [VR_p^{\beta}(2), \ldots, VR_p^{\beta}(k)]$ with $VR_p^{\beta}(k)$ the power transformed VR as in (1), μ_β and $\Sigma_\beta(k)$ are a measure of the expectation and covariance matrix of $\mathbb{V}_{p,\beta}$, respectively. The join VR $QP(k)$ statistic follows a χ^2 distribution with k degrees of freedom.

3.3. Belaire-Franch and Contreras (2004) tests

Wright (2000) advocates a nonparametric alternative to conventional asymptotic VR tests using ranks and signs. Wright's (2000) tests have two advantages over the Lo-MacKinlay test when sample size is relatively small: (i) the rank ($R1$ and $R2$) and sign ($S1$) tests have exact sampling distribution, there is no need to resort to asymptotic distribution approximation, and (ii) the tests may be more powerful than the conventional VR tests in 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 (i.i.d.) assumption, whereas the tests based on signs are exact even under conditional heteroskedasticity. Wright (2000) shows that rank-based tests display low size distortions under conditional heteroskedasticity.

Given T observations of first-differences of a variable, $\{x_1,\ldots,x_T\}$, and let $r(x)$ be the rank of x_t among $\{x_1,\ldots,x_T\}$. Under the null hypothesis that x_t is generated from an i.i.d. sequence,

14. See Chen and Deo (2004) for a discussion on power transformations.

$r(x)$ is a random permutation of the numbers of $1,\ldots,T$ with equal probability. Wright (2000) suggested the $R1$ and $R2$ statistics, defined as:

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

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

where the standardized ranks $r_{1,t}$ and $r_{2,t}$, obtained from alternative standardization to have mean zero, are given by:

$$r_{1,t}=\frac{r(x_t)-\frac{T+1}{2}}{\sqrt{(T-1)(T+1)/12}}$$

$$r_{2,t}=\Phi^{-1}\frac{r(x)}{T+1}$$

where $\phi(k)=2(2k-1)(k-1)(3kT)^{-1}$, and Φ^{-1} is the inverse of the standard normal cumulative distribution function. The R_1 and R_2 statistics follow the same exact sampling distribution. The critical values of these tests can be obtained by simulating their exact distributions.

The test based on the signs of first-differences is given by:

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

where $\phi(k)$ is defined as above, $s_t=2u(x_t,0)$, $s_t(\bar{\mu})=2u(x_t,\mu)$, and:

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

Similarly to R_1 and R_2 tests, the critical values of the S_1 test can be obtained by simulating its exact sampling distribution.

However, as Wright (2000) observes, using several k values would lead to an over-rejection of the null hypothesis. To overcome this problem, Belaire-Franch and Contreras (2004) propose multiple rank and sign VR tests by substituting Wright's rank and sign-based tests in the definition of the Chow-Denning (1993) multiple test procedure.[15] The statistics are defined as:

$$CD_{(R_1)}=\max_{1\le i\le m}\left|R_1(k_i)\right| \qquad CD_{(R_2)}=\max_{1\le i\le m}\left|R_2(k_i)\right| \qquad CD_{(S_1)}=\max_{1\le i\le m}\left|S_1(k_i)\right|$$

The rank-based procedures are exact under the i.i.d. assumption whereas the signs-based procedures are exact under both the i.i.d. and martingale difference sequence assumption.

15. Belaire-Franch and Contreras (2004) show that the rank-based tests are more powerful than the sign-based tests from Monte Carlo simulations based on various models (i.i.d. processes, stochastic volatility models, correlated models, and fractionally integrated processes). Colletaz (2005) and Kim and Shamsuddin (2008) also propose an extension to Wright's VR methodology following Chow-Denning, but only for the rank (R_1) and sign (S_1) tests, respectively.

4. Empirical findings

4.1. Data and descriptive statistics

The data examined consist of the daily and weekly nominal exchange rates for the Australian dollar, British pound, Canadian dollar, Japanese yen, Korean won, New Zealand dollar, Norwegian kroner, Singapore dollar, Swiss franc, Swedish kroner and US dollar relative to the Euro, which includes all the important Euro-based exchange rates that are classified as independently floating by the International Monetary Fund. The data span January 4, 1999, to May 30, 2008, namely 2409 and 486 observations for the daily and weekly data, respectively. For weekly data, the prices are observed on Wednesday or on the following day if the markets are closed on Wednesday. The nominal exchange rate data were compiled by the European Central Bank and were obtained from Thomson Financial Datastream. We use both frequencies to overcome issues such as bias with daily data (e.g., non-trading, bid-ask spread, asynchronous prices) and assumptions about weekly data (alternate day price in the case of non-trading on the day of the week observed), especially for developed markets.

We first present descriptive statistics for the return series, calculated as the first-differences in the logarithms of the nominal exchange rates for daily and weekly data in Tables 2 and 3, respectively. For the daily data (Table 2), the US dollar has the best performance although it is also one of the most volatile exchange rates, as the standard deviation shows. Higher standard deviation is observed for the Japanese yen and the New Zealand dollar. The Swiss franc displays the worst performance and the Swedish kroner the lowest standard deviation.

Table 2 - Descriptive statistics for daily log exchange rate returns

	Mean	SD	Skewness	Kurtosis	JB	LB(10)	$LB^2(10)$	LM(10)
Australia	–6.49	0.64	0.43*	6.05*	1006.95*	18.37*	138.88*	93.30*
Canada	–6.02	0.65	0.41*	5.35*	622.80*	12.81*	78.57*	52.55*
Japan	8.50	0.71	–0.12*	6.64*	1334.60*	3.035	227.34*	137.60*
United Kingdom	4.54	0.44	0.40*	5.35*	615.25*	19.74*	382.77*	190.23*
United States	11.80	0.61	0.23*	4.51*	249.58*	5.07	78.34*	60.18*
New Zealand	–4.50	0.71	0.42*	5.44*	668.52*	14.76*	98.63*	71.48*
Norway	–4.93	0.36	0.45*	4.72*	378.62*	7.59	54.23*	40.73*
Korea	6.00	0.69	0.41*	5.69*	789.17*	4.56	347.34*	176.44*
Singapore	3.67	0.55	0.34*	5.89*	880.03*	6.04	179.21*	112.39*
Sweden	0.22	0.23	–0.41*	7.76*	2337.38*	6.01	382.30*	199.99*
Switzerland	–0.65	0.33	–0.02	5.34*	551.19*	3.75	398.69*	198.13*

* Means significant at the 5% level. The mean values are multiplied by 10^5.

The standard error values are multiplied by 10^2.

The Jarque-Bera statistic is significant at the 1% level for all series, suggesting that foreign exchange returns are highly non-normal. The excess kurtosis and skewness indicate that the empirical distributions of the foreign exchange returns have fat tails and are skewed. The Ljung-Box LB statistics for testing serial correlation show that all the series are not significantly serially correlated, except for the Australian dollar, the British pound and the New Zealand dollar.

We also compute the Ljung-Box LB^2 statistic and the LM test of Engle (1982) to test heteroskedasticity.[16] These two statistics are significant, indicating that all currencies show strong conditional heteroskedasticity. Accordingly, statistical inference for randomness using the VR tests should be based on the heteroskedasticity-adjusted statistic.

For the weekly data (Table 3), all the returns show evidence of significant excess skewness and excess kurtosis and are not normally distributed, except for the US dollar. All the exchange returns are not significantly serially correlated, except for the Swiss franc. Only the New Zealand dollar and the Norwegian kroner do not exhibit conditional heteroskedasticity. Thus, these two currencies do not need to employ the heteroskedasticity-adjusted statistic for the VR tests.

Table 3 - Descriptive statistics for weekly log exchange rate returns

	Mean	SD	Skewness	Kurtosis	JB	LB(10)	$LB^2(10)$	LM(10)
Australia	−2.91	0.14	0.44*	4.81*	82.00*	13.22*	22.04*	20.64*
Canada	−2.65	0.14	0.18*	3.71*	12.91*	3.65	34.63*	37.75*
Japan	4.58	0.16	−0.43*	5.47*	138.33*	8.41	35.20*	23.89*
United Kingdom	2.35	0.01	0.33*	3.94*	26.69*	5.81	77.63*	51.60*
United States	5.93	0.14	−0.03	3.13	0.40	6.19	37.71*	25.74*
New Zealand	−1.92	0.15	0.30*	3.75*	18.84*	9.23	12.58	13.30
Norway	−2.16	0.08	0.32*	3.40*	11.65*	6.48	10.56	12.52
Korea	3.56	0.15	0.46*	5.18*	113.29*	4.25	72.12*	60.73*
Singapore	1.67	0.12	−0.10*	3.49*	5.48*	6.13	111.61*	58.27*
Sweden	0.04	0.08	−0.16*	4.14*	28.49*	18.11*	111.59*	67.73*
Switzerland	0.18	0.05	−0.22	4.15*	30.81*	21.61*	75.98*	56.77*

* Means significant at the 5% level. The mean values are multiplied by 10^4.
The standard error values are multiplied by 10.

4.2. Testing the efficient market hypothesis

Tables 4-7 display the results of the individual and multiple VR tests for daily and weekly exchange markets. The holding periods (k_i's) considered are (2, 4, 8, 16). As advocated by Deo and Richardson (2003), we use relatively short holding periods when testing for the mean reversion using VR tests. The test statistic is displayed for the Belaire-Franch and Contreras ($CD_{(R_1)}$, $CD_{(R_2)}$ and $CD_{(S_1)}$) tests and for the Chen and Deo ($VR_p^\beta(k)$ and $QB(k)$) tests. The p-values are only reported for the Chen and Deo tests.

16. The LB, LB^2 and LM tests are applied on the residuals of the ARMA model, where the lag length is selected based on the Akaike and Schwarz information criterion.

For the daily data, the RWH is rejected only for Singapore at the 5% level from the individual *VR* tests (TABLE 4). Indeed, the *p*-values are less than 5% for the $VR_p^\beta(k)$ test, whatever the horizon k. This can be explained by the fact that the Singaporean foreign exchange market is highly regulated compared with those of other countries. The RWH is also rejected for Australia, Norway and Sweden at $k = 2$, but not all k, implying that these markets are only inefficient over the very short-horizon. This seems to indicate that investors do not take risky speculative positions in the very short-term. The results for the other currencies show that the RWH is not rejected at all lags.

Rejection of the RWH for Singapore is confirmed by the multiple VR tests (TABLE 5), which also reject the null hypothesis for Norway and Sweden at the 5% level. Note that the $CD_{(S_1)}$ does not reject the RWH for Singapore and Sweden, but Belaire-Franch and Contreras (2004) show that the rank-based tests are more powerful than the sign-based tests. This outcome is similar to that obtained by Belaire-Franch and Opong (2005), who find that the RWH is rejected for Canada while this hypothesis is not rejected for Norway and Sweden using the multiple version of Wright's (2000) rank tests, which overcome the problem of size distortions. However, five out of eight size-corrected Wright's VR tests rejected the RWH for Norway.[17] This result can be explained by the fact that their data span from January 5, 1999, to November 11, 2002, while our data span January 5, 1999, to May 30, 2008, which is almost four more years of data. Indeed, we applied the multiple VR tests of Belaire-Franch and Contreras (2004) and Chen and Deo (2006) to their period of interest and we obtained the same results as did Belaire-Franch and Opong (2005).

For the weekly data, the RWH is not rejected for the majority of currencies from the individual VR tests (TABLE 6), except for Sweden and Japan at the 5% and 10% level, respectively. The results of the multiple VR tests (TABLE 7) confirm the rejection of the RWH for Sweden. This indicates that the Swedish exchange market seems to be weak-form inefficient as there is evidence to reject the RWH at daily and weekly frequencies.

For the Singapore and Norwegian exchanges, we find evidence of efficiency from low frequency data (weekly data) and inefficiency from high frequency data (daily data). This indicates a slow price adjustment in response to a shock for these two non-major trading currencies, implying that these markets are more predictable at the short horizon than at the long horizon.

Consistent with Azad (2009), we conclude that eight out of eleven exchange markets (Australia, Canada, Japan, UK, US, New Zealand, Korea and Switzerland) can be considered perfectly weak-form efficient because they behave randomly at all levels of data frequency.[18] This indicates that the daily data reflect the most up-to-date information about prices for these major trading currencies. Thus it is impossible for an exchange trader to generate excess returns over time through speculation.

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17. Moreover, most of Wright's VR tests corrected *p*-values are close to the 10% level in Belaire-Franch and Opong (2005).
18. To take into account structural breaks that can affect the outcomes, we use the modified ICSS algorithm developed by Rapach and Strauss (2008). We identified one break for the Japanese yen and the British pound, and two breaks for the US dollar. We estimate the variance ratio tests for each sub-sample. The outcomes are similar to those for the whole period and are not reported.

Table 4 - Individual variance ratio test results for daily data

	k			
	2	4	8	12
Australia	1.0116	1.0065	0.9986	0.9953
	(1.7015)*	(0.6887)	(–0.0717)	(–0.2032)
	[0.0444]	[0.2455]	[0.4714]	[0.4195]
Canada	0.9955	0.9885	0.9862	0.9872
	(–0.6656)	(–1.2267)	(–1.0673)	(–0.6916)
	[0.2528]	[0.1100]	[0.1429]	[0.2446]
Japan	0.9975	0.9934	0.9837	0.9704
	(–0.3733)	(–0.6602)	(–1.1568)	(–1.4994)**
	[0.3545]	[0.2546]	[0.1237]	[0.0669]
United Kingdom	1.0000	0.9933	0.9928	0.9798
	(0.0082)	(–0.7047)	(–0.5132)	(–1.0077)
	[0.4967]	[0.2405]	[0.3039]	[0.1568]
United States	1.0008	0.9946	0.9954	0.9970
	(0.1625)	(–0.6883)	(–0.3812)	(–0.1428)
	[0.4355]	[0.2456]	[0.3515]	[0.4432]
New Zealand	1.0048	0.9953	0.9924	0.9792
	(0.7634)	(–0.5030)	(–0.5764)	(–1.1103)
	[0.2226]	[0.3075]	[0.2822]	[0.1334]
Norway	1.0124	1.0056	0.9995	1.0016
	(2.1609)*	(0.6859)	(–0.0173)	(0.1309)
	[0.0154]	[0.2464]	[0.4931]	[0.4479]
Korea	0.9955	0.9885	0.9862	0.9872
	(–0.6656)	(–1.2267)	(–1.0673)	(–0.6916)
	[0.2528]	[0.1100]	[0.1429]	[0.2446]
Singapore	0.9869	0.9707	0.9668	0.9629
	(–2.3258)*	(–3.5612)*	(–2.7807)*	(–2.1230)*
	[0.0100]	[0.0002]	[0.0027]	[0.0169]
Sweden	1.0133	1.0040	0.9874	0.9769
	(2.0157)*	(0.4328)	(–0.8657)	(–1.1032)
	[0.0219]	[0.3326]	[0.1933]	[0.1350]
Switzerland	0.9955	0.9885	0.9862	0.9872
	(–0.6656)	(–1.2267)	(–1.0673)	(–0.6916)
	[0.2528]	[0.1100]	[0.1429]	[0.2446]

* and ** Significant at the 5% and 10% levels, respectively.

The estimates of variance ratios are shown in the main row, the $VR_p^{\beta}(k)$ statistics are in parentheses, the *p*-values are in brackets.

Table 5 - Multiple variance ratio test results for daily data

	QB	$CD_{(R_1)}$	$CD_{(R_2)}$	$CD_{(S_1)}$
Australia	5.044	1.708	1.925	1.515
	(0.283)			
Canada	2.161	1.346	1.504	0.959
	(0.706)			
Japan	2.491	0.985	1.475	2.406
	(0.646)			
United Kingdom	4.128	1.115	0.979	0.196
	(0.389)			
United States	2.501	1.257	1.206	1.940
	(0.645)			
New Zealand	7.195	1.933	1.902	0.938
	(0.126)			
Norway	8.210**	2.588*	2.356*	2.854*
	(0.084)			
Korea	2.161	2.038	1.830	1.876
	(0.706)			
Singapore	14.056*	2.835*	3.555*	0.415*
	(0.007)			
Sweden	10.248*	2.347*	2.533*	1.590
	(0.037)			
Switzerland	2.161	1.233	1.580	1.753
	(0.706)			

* and ** Significant at the 5% and 10% levels, respectively.

The *p*-values are only given in parentheses for the Chen-Deo QB test. Critical values used to test the significance of $CD_{(R_1)}$, $CD_{(R_2)}$ and $CD_{(S_1)}$ correspond to 2.33, 2.33 and 2.35 at the 5% level, respectively (Belaire-Franch and Contreras, 2004).

Table 6 - Individual variance ratio test results for weekly data

	k			
	2	4	8	12
Australia	0.9926	0.9900	0.9706	0.9476
	(–0.4848)	(–0.4551)	(–0.9951)	(–1.2568)
	[0.3139]	[0.3245]	[0.1598]	[0.1044]
Canada	0.9948	0.9923	0.9882	0.9976
	(–0.3661)	(–0.3766)	(–0.4038)	(0.0196)
	[0.3571]	[0.3532]	[0.3432]	[0.4922]
Japan	0.9804	0.9712	0.9554	0.9446
	(–1.3343)**	(–1.3954)**	(–1.5701)**	(–1.3734)**
	[0.0910]	[0.0815]	[0.0582]	[0.0848]
United Kingdom	0.9971	0.9854	0.9786	0.9994
	(–0.2230)	(–0.7549)	(–0.7418)	(0.0755)
	[0.4118]	[0.2252]	[0.2291]	[0.4699]
United States	1.0013	1.0058	1.0185	1.0156
	(0.1314)	(0.3729)	(0.8159)	(0.5295)
	[0.4477]	[0.3546]	[0.2073]	[0.2982]
New Zealand	0.9987	0.9821	0.9634	0.9449
	(–0.0785)	(–0.9099)	(–1.3209) **	(–1.3881)**
	[0.4687]	[0.1814]	[0.0933]	[0.0825]
Norway	1.0128	1.0082	1.0129	1.0054
	(1.0340)	(0.4938)	(0.5685)	(0.2383)
	[0.1506]	[0.3107]	[0.2849]	[0.4058]
Korea	1.0105	1.0126	1.0227	1.0444
	(0.6111)	(0.5828)	(0.8169)	(1.1929)
	[0.2706]	[0.2800]	[0.2070]	[0.1164]
Singapore	0.9883	0.9911	0.9904	0.9840
	(–0.8735)	(–0.4346)	(–0.3122)	(–0.3446)
	[0.1912]	[0.3319]	[0.3774]	[0.3652]
Sweden	0.9671	0.9531	0.9357	0.9277
	(–2.0128)*	(–2.0724)*	(–2.0659)*	(–1.6181)**
	[0.0213]	[0.0191]	[0.0194]	[0.0528]
Switzerland	0.9809	0.9863	0.9622	0.9645
	(–1.2238)	(–0.6011)	(–1.2375)	(–0.7743)
	[0.1105]	[0.2739]	[0.1079]	[0.2194]

* and ** Significant at the 5% and 10% levels, respectively. The estimates of variance ratios are shown in the main row, the $VR_p^{\beta}(k)$ statistics are in parentheses, the *p*-values are in brackets.

Table 7 - Multiple variance ratio test results for weekly data

	QB	$CD_{(R_1)}$	$CD_{(R_2)}$	$CD_{(S_1)}$
Australia	2.250	2.103	1.966	1.680
	(0.690)			
Canada	0.827	1.346	0.879	0.898
	(0.935)			
Japan	2.805	1.691	1.686	0.772
	(0.591)			
United Kingdom	3.024	1.312	1.271	1.358
	(0.554)			
United States	1.484	0.905	0.462	1.412
	(0.830)			
New Zealand	2.820	1.933	1.875	1.086
	(0.589)			
Norway	2.352	0.977	0.819	1.413
	(0.671)			
Korea	1.872	0.724	0.563	1.039
	(0.759)			
Singapore	1.064	0.598	1.176	0.461
	(0.900)			
Sweden	5.1990	2.572*	2.687*	2.195*
	(0.268)			
Switzerland	7.403	1.583	1.801	0.500
	(0.116)			

* and ** Significant at the 5% and 10% levels, respectively.

The *p*-values are only given in parentheses for the Chen-Deo *QB* test. Critical values used to test the significance of $CD_{(R_1)}$, $CD_{(R_2)}$ and $CD_{(S_1)}$ correspond to 2.14, 2.14 and 2.12 at the 5% level, respectively (Belaire-Franch and Contreras, 2004).

5. Conclusion

This study employed new variance ratio tests to evaluate the random walk behavior of eleven major Euro exchange rates over the period of January 4, 1999, to May 30, 2008, using daily and weekly data. These tests, which are robust to heteroskedasticity and non-normality, are the Chen and Deo (2006) power-transformed tests and the multiple Belaire-Franch and Contreras (2004) rank and sign-based tests.

The results suggest that Euro-based exchange rates for the major trading countries (Australia, Canada, Japan, UK, US, New Zealand, Korea and Switzerland) follow the random walk hypothesis at both data frequencies, and are therefore significantly and perfectly weak-form efficient, suggesting no excess returns over time through speculation. This outcome is not necessarily the case for non-major trading currencies, especially for the Swedish kroner, where the random walk hypothesis is rejected at daily and weekly frequencies. Finally, the weak-form efficiency is rejected for daily data but is not rejected for weekly data for Singapore and Norway, suggesting the possibility of abnormal returns through speculation on the short horizon.

A. C. & O. D.[19]

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