EFFICIENCY TESTS IN THE SPANISH FUTURES MARKETS

CHUN I LEE
IKE MATHUR*

The Spanish futures markets, the MEFF RENTA FIJA, and the MEFF RENTA VARIABLE, are among the fast-growing futures markets in the world. These markets are known for their cutting-edge technological innovations related to trading, providing information, clearing, and settlement. The growing importance of these markets for both foreign and domestic investors motivated the examination of their efficiency. Test results from serial correlations, unit root tests, and variance ratio tests overwhelmingly show that the random walk hypothesis cannot be rejected, indicating that the MEFF markets are efficient. © 1999 John Wiley & Sons, Inc. Jrl Fut Mark 19: 59–77, 1999

INTRODUCTION

The Spanish futures markets are part of a single company called MEFF Sociedad Holding de Productos Financieros Derivados (MEFF Holding Corporation of Financial Derivatives, referred to as MEFF hereafter). The Spanish derivative markets include the MEFF RENTA FIJA (located in Barcelona and responsible for interest rate and foreign exchange futures) and the MEFF RENTA VARIABLE (situated in Madrid and in charge of stock index futures and options, and equity options).

We thank two anonymous JFM referees for their helpful comments, and Sharon Rose for her assistance with preparing the manuscript.

*Correspondence author, Department of Finance, Southern Illinois University at Carbondale, Carbondale, IL 62901-4626.

■ *Chun I Lee is an Assistant Professor at Texas Southern University.*

■ *Ike Mathur is a Professor at Southern Illinois University at Carbondale.*

The Journal of Futures Markets, Vol. 19, No. 1, 59–77 (1999)

CCC 0270-7314/99/010059-19

MEFF was incorporated on 20 December 20 1991 by purchasing all shares of the two existing Spanish derivative exchanges. Functioning both as an exchange and as a clearing house, the shareholder-owned organization has 40 shareholders that include the major Spanish banks, savings banks, and the four Spanish stock exchanges. It is supervised by the Comision Nacional del Mercado De Valores (National Securities Market Commission) and regulated by the Market Regulations, the Royal Decree 1814/1991 and the Securities Market Law of 1988. MEFF's reputation as a cutting edge technology innovator has steadily gained ground as evidenced by the decision of Portugal's futures and options market to adopt its trading system.[1] The system allows the integration of trading, clearing, and settlement, in addition to making available up-to-date real time information to traders and for the supervision and control of risk by the exchanges. Traders can request information on the prevailing state of the market at all times. Simply by pressing two keystrokes on their terminal, they can cancel, resend, and modify orders quickly and easily should there be an unexpected change on the market. The terminals also have a ticker for recording the orders, cancelations, and transactions undertaken by traders. For supervision, consultations and reports can be obtained at any point in real time, during and outside trading hours, on open positions, state of orders that have been transmitted, margins, realized and unrealized profits and losses, and commissions. All of this information can be listed based on the type of member, client, account, contract, and so forth.

These trading innovations have attracted both domestic and foreign investors. The result has been that, in their relatively short history since inception in 1989, the Spanish futures markets have become one of the fastest growing emerging futures markets in the world. Being one of the emerging markets and with such a phenomenal growth, the issue of market efficiency is certainly of concern to many participants in these markets, including non-Spanish investors.[2]

The above discussion suggests that an analysis of the efficiency of the Spanish futures markets would be of interest to both traders and researchers. Thus the purpose of this paper is to conduct comprehensive efficiency tests on the prices and returns of the contracts traded on the

[1]See Financial Times Survey: Spanish Banking and Finance. Financial Times, October 15, 1996.
[2]According to the MEFF RENTA FIJA, about 26% of the transactions conducted during 1995 are related to trades on behalf of nonresidents. Furthermore, the recognition by the U.K. Securities Investment Board and the U.K. Treasury as a "Designated Investment Exchange" and "Recognized Overseas Investment Exchange," respectively, and the authorization by the U.S. Commodities Futures Trading Commission to promote the sale of its products in the USA have made the MEFF RENTA FIJA a market appealing to non-Spanish investors.

Spanish futures markets. These tests are significant for two reasons. First, there is a paucity of research in emerging futures markets, such as is the case for Spain. Second, even among studies that test for efficiency in well-established markets, the evidence is mixed. By examining the Spanish markets employing the refined variance estimator of serial correlations of Diebold (1986) and the variance ratio tests of Lo and McKinlay (1988) that adjust for heteroscedasticity, further evidence on market efficiency is obtained that can shed additional light on the price behavior of assets.

THE RANDOM WALK HYPOTHESIS AND MARKET EFFICIENCY

Since 1900, when Louis Bachelier advanced the concept of efficient markets, financial economists have described and tested the random walk model of stock prices. In studies going back to work by Working (1934), Cowles and Jones (1937), Kendall (1953), and Fama (1965), researchers have tested whether stock prices follow a random walk by examining whether serial correlation coefficients for successive price changes are zero. The general finding of serial correlations that are close to zero has led these researchers to conclude that the stock market is weak form efficient.

Although this evidence of randomness is consistent, stock prices do not conform to the statistician's ideal of a random walk. For example, Merton (1980) shows that the changes in the variance of a stock's returns can be predicted from its variance in the recent past, although this evidence does not violate the weak form of the efficient market hypothesis. What is disturbing to many is the evidence of calendar anomalies—e.g., the January effect documented by Keim (1983) and Barone (1990), the weekend effect by French (1980), and so forth. Specifically with regard to the Spanish equity market, Alonso and Rubio (1990) show that there is overreaction in the market, whereas Rubio (1989) shows that the intertemporal capital asset pricing model does not apply to Spanish equities. Failing to fully explain the anomalies, most financial economists concede that such departures from randomness are minor.

In addition to serial correlations and nonparametric run tests, further tests of market efficiency using unit root tests developed by Dickey and Fuller (1979, 1981), and extended by others, have also been performed. The unit root tests are superior to serial correlation tests because the tests control for the effect of a upward trend in prices. However, because the null hypothesis of the augmented Dick-Fuller (ADF) test commonly used in these tests is that the series has a unit root, the ADF test has been shown to fail to distinguish between a unit root and weakly-stationary alternatives. Consequently, the failure to reject the null hypothesis of a unit root does not conclusively convince one that a unit root exists. To address this shortcoming of the ADF, Kwiatkowski, Phillips, Schmidt, and Shin (1992) (KPSS) have proposed an alternative test in which the null hypothesis is stationarity, with a unit root as the alternative hypothesis.

Despite their superiority over serial correlations, unit root tests fail to detect some important departures from the random walk. As a better alternative to these tests, variance ratio tests have recently been used to test for random walk (e.g., Liu and He, 1991). Using the variance ratio tests developed by Lo and MacKinlay (1988), Liu and He tested and rejected the random walk hypothesis for five weekly exchange rates. Ayadi and Pyun (1994) apply the same test to stocks in the Korean Stock Exchange and show that after adjusting for serial correlation and heteroscedasticity the random walk hypothesis cannot be rejected.

DATA AND METHODOLOGY

The MEFF Exchange

The activities of the MEFF have been divided into fixed income products and equity based products, with MEFF RENTA FIJA (located in Barcelona) responsible for interest rate and foreign exchange futures, and MEFF RENTA VARIABLE (situated in Madrid) in charge of stock index futures and options, and equity options.

In their brief history since inception, the markets have seen dramatic growth. Annual trading volume in futures and options has risen from 188,744 in 1990 to more than 20,000,000 in 1995, while the daily trading average has increased from 938 in 1990 to more than 80,000 in 1995. Among the contracts traded on the MEFF are futures on 10-year and 3-year Government bonds, futures on 90-day and 360-day Madrid Interbank Deposit Rates (MIBOR), and options on these futures. The settlement at expiration involves physical delivery for the 3- and 10-year government bonds, whereas the 90- and 360-day MIBOR contracts (MIBOR'90 and MIBOR'360) are cash settled based on spot 3-month and 360-day MIBOR. To be in line with the size of international interbank contracts, two new MIBOR contracts (MIBOR'90 Plus and Mibor'360 Plus) with a nominal value of 100 million Pesetas were launched on 15 June 15 1995. Except for the nominal value, the new contracts are identical to the existing MIBOR contracts, which have a nominal value of 10 million Pesetas and ceased to exist after 26 June 1995. The Deutsche Mark/Spanish Peseta contract was delisted in 1993 due to a lack of market demand and liquidity. There is no foreign exchange contract traded currently on the MEFF. Similarly, there are no intermediate-term interest rate contracts available subsequent to the delisting of the 5-year Government bond futures contract due to a lack of interest from traders.

Trading in the MEFF is conducted via its computerized trading system. This electronic trading system was chosen over the open outcry system mainly to enable traders who are geographically dispersed to trade the contracts. The system allows the integration of trading, clearing and settlement of trades. It also provides the exchange with better means of supervision and control of risk.

Data

Daily opening and closing prices, from the Knight-Ridder Database, of the nearby contracts of six futures traded on the MEFF are used in the present study. The six contracts, along with the beginning and ending dates of available data, are IBEX 35 Index (IBEX), 921005-951127, MIBOR, 90-Day (MIBOR90), 901022-950626, National Bond, 10-Year (LBOND), 920413-951127, National Bond, 3-Year (SBOND), 900319-950313, Deutsche Mark/Spanish Peseta (DM/SP), 910923-930614, and MIBOR, 360-Day (MIBOR360), 931004-950626. The nearby contracts (i.e., contracts nearest to delivery) are examined because, in general, they are the most active contracts. A switch from the nearby contracts to the contracts next nearest to delivery is made during the delivery month of the nearby contracts.

The Deutsche Mark/Spanish Peseta contract was delisted in June 1993. The inclusion of this contract makes it possible to examine whether this short-lived contract is any less efficient than contracts that have continued to be traded, thereby providing a more complete picture for the evidence of efficiency.

Methodology

Returns

To examine the difference in volatility during trading and nontrading periods, both close-to-close ($R_{C\text{-}C}$) and open-to-open ($R_{O\text{-}O}$) returns are used and calculated as follows:

$$R_{C-Ct} = \text{Ln}\ (P_{Ct}/P_{Ct-1}) \text{ and}$$

$$R_{O-Ot} = \text{Ln}\ (P_{Ot}/P_{Ot-1}),$$

where P_{Ct} and P_{Ot} are the closing and opening price of day t. For each day these daily returns are further decomposed into close-to-open and open-to-close returns and calculated as follows:

$$R_{C-Ot} = \text{Ln}\ (P_{Ot}/P_{Ct-1}) \text{ and}$$

$$R_{O-Ct} = \text{Ln}\ (P_{Ct}/P_{Ot}).$$

Unit Root Tests

The efficiency of the Spanish futures markets is tested using the unit root tests of ADF, and KPSS, and the variance-ratio tests of Lo and MacKinlay (1988). Given the complementary nature of these three tests, it is appropriate to employ all of them to test for efficiency in the Spanish futures markets. It is expected that by doing so, the robustness of the conclusions can be well established.

The ADF tests developed by (among others) Dickey and Fuller (1979, 1981) have been used in the literature to test for the predictability of prices and returns, and hence the efficiency of markets (see, e.g., Annuar et al., 1994; Arshanapalli and Doukas, 1994; and Szakmary et al., 1995). With a unit root, i.e., an I(1) process, as the null hypothesis, the following regression is conducted on the natural logarithm of prices:

$$\Delta p_t = \eta_0 + \eta_1 T + \eta_2 p_{t-1} + \sum_{i=1}^{L} \gamma_t \Delta p_{t-i} + \mu_t \tag{1}$$

where T is the number of observations.

By controlling for the effect of market price trends, the ADF procedures are appropriate for testing for a unit root. However, the null hypothesis of the ADF test, and the way it is tested, dictates that the null hypothesis be accepted unless there is strong evidence against it. The main problem with this is that most time series are not very informative about whether or not there is a unit root. In other words, the standard ADF tests are not very powerful against relevant alternatives. To overcome this lack of power in rejecting the null hypothesis of a unit root, one needs to perform tests of stationarity in addition to the tests of the null hypothesis of a unit root. This issue is addressed by using the KPSS test, which is specifically designed to perform a test that has stationarity as the null hypothesis and a unit root as the alternative hypothesis. The test statistic is calculated as:

$$\eta_u = T^{-2} \sum_{t=1}^{T} S_t^2/S^2\,(L), \tag{2}$$

where L is the lag parameter, S_t is the cumulative sum of the residuals (e_t) from a regression of the series on a constant and a linear trend (i.e., $S_t = \Sigma\, e_t$, $t = 1, 2, \ldots, T$) and where

$$S^2(L) = T^{-1} \sum_{t=1}^{T} e_t^2 + 2T^{-1} \sum_{s=1}^{L} (1 - s/(L + 1)) \sum_{t=s+1}^{T} e_t e_{t-s} \tag{3}$$

The null hypothesis of stationarity is rejected in favor of the unit root alternative if the calculated test statistic exceeds the critical values estimated in Kwiatkowski et al. (1992, Table 1, P. 166).

Variance Ratio Tests

The detection of a unit root in a return series has been used as a basis for supporting the random walk hypothesis, hence the efficiency of the underlying market. Whereas a unit root is one of the implications of the random walk, many studies (e.g., Liu and He, 1991) have found that unit root tests cannot detect certain important departures from random walk. As an alternative, the variance ratio test developed by Lo and MacKinlay (1986) has been used to test the random walk hypothesis. It is shown by Lo and MacKinley (1989) to be more powerful than either the Box-Pierce or ADF tests against several interesting alternative hypotheses, including an AR(1), an ARIMA(1,1,1), and an ARIMA(1,1,0). Based on the premise that the variance of random walk increments in a finite sample increases linearly with the sampling interval, variance ratio tests involve testing whether the ratio of variances of different intervals weighted by their length is one. Specifically, let P_t be the natural logarithm of price series. Under the random walk hypothesis, p_t follows the following form:

$$p_t = \psi + p_{t-1} + \varepsilon_t, \tag{4}$$

and the variance of its qth-differenced series, $\{p_t - p_{t-q}\}$, would be q times the variance of its first- differenced series, $\{p_t - p_{t-1}\}$. Therefore, given q+1 observations of the price series, $p_1, p_2, p_3, \ldots, p_{nq+1}$, the ratio of 1/q of the variance of the series to the variance of the series should equal one.

The variance ratio of q-differenced series is defined as $VR(q) = \Sigma\sigma_c^2(q)/\Sigma\sigma_a^2(q)$.Definitions and details are provided in the Appendix.

EMPIRICAL EVIDENCE

Basic Statistics

Panel A, Table I presents the basic statistics of daily close-to-close and open-to-open returns.[3] With the exception of IBEX, the returns are not normally distributed, and are characterized by both high skewness and kurtosis. Based on the nonparametric sign test, neither close-to-close nor open-to-open returns have an average significantly greater than zero. The open-to-open returns have higher standard deviation than the close-to-close returns, suggesting higher volatility during trading hours as documented in the literature (e.g., French and Roll, 1986).

Returns of close-to-open and open-to-close periods are reported in Panel B, Table I. The open-to-close returns seem to have higher average returns than open-to-close returns, although both are insignificantly different from zero except for the close-to-open returns in MIBOR90 and the open-to-close returns in DM/SP. The normality of the close-to-open returns on IBEX cannot be rejected, indicating that the normality of IBEX daily returns reported in Panel A is due to the normality of the returns in the non-trading period.

Day-of-the-Week Effects

Table II reports the close-to-close returns based on the day of the week (Cyr and Llewellyn, 1994). The normality of the IBEX reported in Table I can be further attributed to all weekdays except Monday. Returns on Monday are unique in that they are negative, on average, for all contracts. In addition, on Monday the volatility is the highest for three contracts, IBEX, MIBOR90, and MIBOR360, whereas for the other three contracts it is certainly not the smallest. In addition to this Monday effect, there appears to be a Thursday effect also, when returns, on average, are positive.

Serial Correlations

Table III reports the serial correlations of close-to-close and open-to-open returns. Based on the homoscedastic standard deviations reported in row

[3]The results are similar when returns are measured as the ratio of prices minus 1: $P_t/P_{t-1} - 1$.

TABLE I

Basic Statistics for Returns

Panel A: Close-to-close and open-to-open returns

	IBEX		MIBOR90		LBOND		SBOND		DM/SP		MIBOR360	
	C-C	O-O	C-C	O-O	C-C	O-O	C-C	O-O	C-C	O-O	C-C	O-O
Observations (N)	805	797	1164	1150	924	911	1057	1050	427	425	424	417
Mean*	6.28	7.95	0.16	0.16	1.12	1.14	0.74	0.78	4.34	4.34	-0.13	-0.10
Std. Dev.*	125.0	135.0	15.0	16.0	65.0	73.0	68.0	68.0	50.0	52.0	12.0	13.0
Skewness	-0.20	0.16	5.19	7.46	-0.22	-0.27	7.10	7.15	3.98	3.60	-1.08	-0.43
Kurtosis	1.37	1.05	78.88	123.5	3.40	4.14	219.1	220.9	38.85	33.67	27.23	23.32
T-stat.	1.42	1.64	0.37	0.35	0.53	0.47	0.35	0.37	1.80**	1.73**	-0.21	-0.16
Prob(T>0)	0.15	0.10	0.71	0.73	0.60	0.64	0.72	0.71	0.07	0.08	0.83	0.88
Prob(S)	0.12·	0.16	0.55	0.23	0.30	0.45	0.34	0.19	0.33	0.35	0.61	0.77
Prob(Normal)	0.95	0.28	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00

Panel B: Close-to-open and Open-to-close Returns

	IBEX		MIBOR90		LBOND		SBOND		DM/SP		MIBOR360	
	C-O	O-C	C-O	O-C	C-O	O-C	C-O	O-C	C-O	O-C	C-O	O-C
Observation (N)	806	805	1165	1164	925	924	1058	1057	428	427	425	424
Mean*	2.38	4.03	-0.42	0.58	-0.68	1.78	0.16	0.60	-0.21	4.53	0.59	-0.70
Skewness	-0.13	-0.00	7.71	8.98	-0.03	-1.66	0.56	7.40	-2.61	3.84	2.66	-0.54
Kurtosis	1.43	4.72	167.2	172.1	3.18	19.24	25.37	231.7	38.89	36.62	22.81	53.45
T-stat	0.63	1.51	-1.34	2.04**	-0.37	1.46	0.36	0.29	-0.24	1.86**	1.65	-1.36
Prob(T>0)	0.53	0.13	0.18	0.04	0.71	0.14	0.72	0.77	0.81	0.06	0.10	0.17
Prob(S)	0.42	0.13	0.01	0.23	0.73	0.00	0.72	0.15	0.59	0.35	0.25	0.19
Prob(Normal)	0.14	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00
Std. Dev.*	107.0	76.0	11.0	10.0	56.0	37.0	14.0	68.0	18.0	40.0	7.0	11.0

Notes:
O-O: Open-to-open returns
C-C: Close-to-close returns
C-O: Close-to-open returns
O-C: Open-to-close returns
*: Expressed in basis points.
**: Significant at better than 10% level

TABLE II

Close-to-Close Returns by the Day-of-the-Week

	IBEX					MIBOR90					LBOND				
	M	Tu	W	Th	F	M	Tu	W	Th	F	M	Tu	W	Th	F
Observations (n)	155	163	164	162	161	230	235	236	232	231	181	188	187	185	183
Mean	-11.89	11.16	-8.88	19.09	21.39	-0.31	-0.55	-1.24	1.77	1.15	-4.37	1.78	-4.01	6.14	6.04
Std. Dev.	137.6	127.1	118.1	125.8	113.5	19.4	12.2	9.5	17.9	11.6	69.1	58.3	59.8	75.2	59.5
Skewness	-0.88	-0.19	0.13	0.42	-0.16	4.60	0.54	-1.20	8.16	0.86	-0.35	0.08	-1.34	0.14	0.07
Kurtosis	3.12	0.58	1.06	0.19	0.24	53.99	11.17	4.69	98.05	5.88	2.14	1.50	5.33	4.38	1.04
T-stat.	-1.08	1.12	-0.96	1.93	2.39	-0.24	-0.69	-2.01	1.50	1.51	-0.85	0.42	-0.92	1.11	1.37
Prob(T>0)	0.28	0.26	0.34	0.06	0.02	0.81	0.49	0.05	0.13	0.13	0.40	0.68	0.36	0.27	0.17
Prob(S)	0.47	0.23	0.26	0.15	0.01	0.35	0.37	0.12	0.50	0.18	0.66	0.65	0.97	0.36	0.17
Prob(Normal)	0.01	0.64	0.90	0.15	0.86	0.00	0.00	0.00	0.00	0.00	0.07	0.36	0.00	0.00	0.23
	SBOND					DM/SP					MIBOR360				
	M	Tu	W	Th	F	M	Tu	W	Th	F	M	Tu	W	Th	F
Observations (n)	214	214	210	207	212	86	88	86	85	82	77	86	86	87	88
Mean	-0.22	5.98	1.28	2.92	-6.22	-0.74	-1.58	12.1	5.73	6.43	-3.83	0.77	0.96	1.32	-0.26
Std. Dev.	73.7	98.3	15.0	54.2	70.9	46.7	27.4	64.7	54.7	48.5	16.4	14.7	9.0	8.6	9.2
Skewness	2.39	13.06	0.43	10.54	-10.35	-0.66	-0.04	6.33	3.06	2.79	-4.36	2.92	0.68	0.10	-0.96
Kurtosis	79.86	185.5	11.08	135.9	122.2	6.27	14.29	48.33	24.21	18.29	25.39	18.73	3.71	3.83	4.50
T-stat	-0.04	0.89	1.21	0.78	-1.28	-0.15	-0.54	1.74	0.96	1.20	-2.05	0.49	0.99	1.43	-0.26
Prob(T>0)	0.96	0.37	0.23	0.44	0.20	0.88	0.59	0.09	0.34	0.23	0.04	0.63	0.33	0.16	0.79
Prob(S)	0.74	0.49	0.05	0.90	0.74	0.92	0.18	0.08	0.49	0.26	0.14	1.00	0.46	0.25	0.52
Prob(Normal)	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00

Note: M: Monday
Tu: Tuesday
W: Wednesday
Th:Thursday
F: Friday

TABLE III

Serial Correlations of Returns

Panel A: Close-to-Close Returns

Lag	IBEX	MIBOR90	LBOND	SBOND	DM/SP	MIBOR360
1	-10.005	0.088	0.057	-0.006	0.221	-0.233
	(0.035)[a]	(0.029)	(0.032)	(0.030)	(0.048)	(0.048)
	(0.042)[b]	(0.072)	(0.047)	(0.006)	(0.192)	(0.181)
2	0.020	0.008	0.005	0.018	0.079	0.073
	(0.035)	(0.029)	(0.033)	(0.030)	(0.050)	(0.051)
	(0.045)	(0.050)	(0.038)	(0.011)	(0.077)	(0.058)
3	-0.070	0.034	0.008	0.010	0.005	0.049
	(0.035)	(0.029)	(0.033)	(0.030)	(0.050)	(0.051)
	(0.042)	(0.084)	(0.042)	(0.009)	(0.036)	(0.051)
4	-0.008	-0.030	0.021	0.001	0.014	-0.076
	(0.035)	(0.029)	(0.033)	(0.030)	(0.050)	(0.051)
	(0.037)	(0.063)	(0.036)	(0.007)	(0.055)	(0.062)
5	-0.025	-0.012	-0.026	0.003	-0.007	0.073
	(0.035)	(0.029)	(0.033)	(0.030)	(0.051)	(0.051)
	(0.037)	(0.044)	(0.037)	(0.004)	(0.033)	(0.071)
6	-0.025	0.020	-0.023	-0.142	-0.042	-0.022
	(0.035)	(0.029)	(0.033)	(0.030)	(0.051)	(0.052)
	(0.035)	(0.040)	(0.035)	(0.139)	(0.026)	(0.061)
Box-Pierce Q(6)	5.57	12.36*	4.77	22.11***	24.68***	31.65***
Adjusted Box-Pierce Q(6)	4.20	2.27	2.89	6.83	5.28	6.96

(Continued)

2 for lag 1, it appears that lag-one serial correlations are significant for the close-to-close returns on MIBOR90, DM/SP, and MIBOR360 and for the open-to-open returns on IBEX, LBOND, DM/SP, and MIBOR360. Based on these seemingly significant correlations, one may be tempted to conclude that serial correlations are not zero, and conclude that there is significant inefficiency in the MEFF. However, this turns out to be a false conclusion due to the heteroscedasticity of the returns. As reported in Table I and evidenced in the literature, returns are heteroscedastic, especially for SBOND, which has the largest kurtosis in daily returns. This deviation from homoscedasticity leads to the underestimation of variance and overstatement of the significance of the serial correlations. Following Hsieh (1988), the variance estimator proposed by Diebold (1986), which adjusts for heteroscedasticity, is estimated and reported in the third row for each lag. It is clear from these adjusted

TABLE III Continued

Serial Correlations of Returns

Panel B: Open-to-Open Returns

Lag	IBEX	MIBOR90	LBOND	SBOND	DM/SP	MIBOR360
1	-0.086	-0.012	-0.065	0.008	0.168	-0.271
	(0.035)	(0.029)	(0.032)	(0.030)	(0.048)	(0.048)
	(0.042)	(0.044)	(0.044)	(0.009)	(0.184)	(0.175)
2	0.007	0.069	0.018	0.007	0.066	0.028
	(0.035)	(0.029)	(0.033)	(0.030)	(0.049)	(0.052)
	(0.039)	(0.091)	(0.040)	(0.006)	(0.075)	(0.063)
3	-0.034	-0.050	-0.006	0.005	0.014	0.160
	(0.035)	(0.029)	(0.033)	(0.030)	(0.049)	(0.052)
	(0.039)	(0.079)	(0.051)	(0.007)	(0.040)	(0.067)
4	-0.039	-0.028	-0.025	0.003	0.049	-0.132
	(0.035)	(0.029)	(0.033)	(0.030)	(0.049)	(0.053)
	(0.037)	(0.037)	(0.036)	(0.008)	(0.058)	(0.076)
5	0.013	0.038	0.062	0.004	-0.054	0.042
	(0.035)	(0.029)	(0.033)	(0.030)	(0.050)	(0.053)
	(0.042)	(0.036)	(0.040)	(0.000)	(0.040)	(0.058)
6	0.001	-0.053	-0.043	-0.139	-0.004	-0.030
	(0.035)	(0.029)	(0.033)	(0.030)	(0.050)	(0.054)
	(0.039)	(0.040)	(0.041)	(0.141)	(0.031)	(0.054)
Q(6)	8.44	14.81***	10.37	21.09***	16.53***	51.47***
	6.25	4.61	6.58	13.16**	4.31	12.18*

Notes:
- [a] Standard deviation of serial correlation
- [b] Heteroscedasticity-adjusted standard deviation based on Diebold [1986]
- * Significant at 10% level
- ** Significant at 5% level
- *** Significant at 1% level

estimates of standard deviation that none of the daily returns is serially correlated.

The Box-Pierce Q statistics confirm the above findings. Before adjusting for heteroscedasticity, the white noise hypothesis is rejected for all but IBEX and LBOND for the close-to-close returns. After adjusting for heteroscedasticity, the white noise hypothesis is not rejected for any of the returns. Similarly, for the open-to-open returns in Table III, Panel B, the white noise hypothesis is rejected for all futures except IBEX and LBOND. However, after adjusting for heteroscedasticity, the white noise hypothesis is rejected only for SBOND and MIBOR 360. Judging from the extremely large kurtosis reported in Table I above, these rejections should be viewed as inconsequential to the overall conclusion of zero serial correlations.

Unit Root Test Results

Table IV reports the results of the unit root tests. Panel A shows that the null hypothesis of one unit root cannot be rejected for any of the six contracts. The evidence from the KPSS tests in Panel B shows that the null hypothesis of no unit root is significantly rejected for all six contracts, thus further supporting the finding of unit roots in the price series. This overwhelming evidence of unit roots presented in Table IV provides further support for the efficiency in the MEFF.

Variance-Ratio Test Results

Table V provides further strong evidence that the random walk hypothesis cannot be rejected for the price series of futures contracts traded on the MEFF. The variance ratios for lags 2, 4, 8, and 16, i.e., VR(Q), Q=2, 4, 8, and 16, are estimated and reported in Table 5. It appears that the ratios are close to 1 for IBEX, LBOND, and SBOND, while MIBOR90 and DM/SP have values much greater than 1, and MIBOR360 has values much smaller than 1. The values of the Z statistic based on the homoscedasticity assumption leads one to reject the random walk hypothesis for MIBOR90, DM/SP, and MIBOR360 at the 5% significant level. However, similar to the tests of serial correlation reported in Table III, they lead to false conclusions due to the violation of the assumption of homoscedasticity in the price series. Adjusting for heteroscedasticity, none of the adjusted Z statistics reported in Table V is significantly different from zero, indicating overwhelmingly that the null hypothesis of unit-variance-ratio cannot be rejected.

SUMMARY

In summary, the results of the above unit root tests and variance-ratio tests show overwhelming evidence supporting the existence of a unit root in the returns of MEFF contracts. The ADF tests cannot reject the null hypothesis of a unit root, which is supported by the KPSS test results that indicate that one can reject the hypothesis of no unit root. Combining these two test results, the evidence of the existence of a unit root is robust. The robustness of this evidence of random walk is further supported by the variance-ratio test results.

CONCLUSION

The growing importance of Spanish futures markets for both domestic and foreign investors suggests a critical examination of their efficiency. A

TABLE IV

Unit Root Tests

Panel A: ADF Test[1]

Test statistic is the t-statistic on η_2 from the regression

$$\Delta p_t = \eta_0 + \eta_1 T + \eta_2 p_{t-1} + \sum \gamma_i \Delta p_{t-i} + \mu_t$$

	IBEX	MIBOR90	LBOND	SBOND	DM/SP	MIBOR360
Open-to-Open Returns						
Test Statistics	-2.613	-2.932	-1.349	-1.849	-2.090	-2.657
Lags in ADF	1	1	1	1	1	1
Close-to-Close Returns						
Test Statistics	-2:728	-3.009	-1.379	-1.818	-2.144	-2.449
Lags in ADF[2]	1	1	1	1	1	1

Panel B: KPSS Test[2]

The test statistic is $T^{-2}\sum S_t^2 / S^2(L)$, where

$$S_t = \sum e_t,\ t=1,2,\ldots,T, \text{ and}$$
$$S^2(L) = T^{-1}\sum e_t^2 + 2T^{-1}\sum (1-S/(L+1))\sum e_t e_{t-s}$$

The e_ts are the residuals from a regression of the series being tested on a constant and trend.

	IBEX	MIBOR90	LBOND	SBOND	DM/SP	MIBOR360
Open-to-Open Returns						
Test statistics[2]: L=0	14.021*	9.186*	12.079*	5.657*	5.288*	5.326*
Test statistic: L=9	1.457*	0.968*	1.229*	0.587*	0.579*	0.581*
Test statistic: L=29	0.523*	0.353*	0.423*	0.209*	0.226*	0.239*
Close-to-Close Returns						
Test statistic: L=0	13.951*	9.343*	11.708*	5.666*	5.422*	5.309*
Test statistic: L=9	1.449*	0.983*	1.191*	0.588*	0.582*	0.579*
Test statistic: L=29	0.520*	0.359*	0.411*	0.210*	0.227*	0.239*

1. Critical values are -3.13, -3.41, and -3.96 at the 10%, 5%, and 1% levels, respectively. The null hypothesis that the series is I(1), i.e., non-stationary, is rejected if the test statistic is greater than the critical value.

2. Critical values are 0.119, 0.146, and 0.216 at the 10%, 5%, and 1% levels, respectively. The null hypothesis of stationarity is rejected if the test statistic exceeds the critical values.

TABLE V

Estimate of Variance-Ratio VR(Q) and Variance-Ratio Test Statistics Z(Q) and Z*(Q)

Lag(q)	IBEX 2	IBEX 4	IBEX 8	IBEX 16	MIBOR90 2	MIBOR90 4	MIBOR90 8	MIBOR90 16	LBOND 2	LBOND 4	LBOND 8	LBOND 16
VR (q)	1.01	0.99	0.90	0.93	1.09	1.16	1.18	1.00	1.06	1.10	1.13	1.09
Z (q)	0.15	-0.11	-0.93	-0.47	3.05	2.97	2.03	-0.03	1.79	1.64	1.33	0.60
Z* (q)	0.13	-0.08	-0.51	-0.20	1.24	1.07	0.57	-0.01	1.25	1.01	0.66	0.23

Lag(q)	SBOND 2	SBOND 4	SBOND 8	SBOND 16	DM/SP 2	DM/SP 4	DM/SP 8	DM/SP 16	MIBOR360 2	MIBOR360 4	MIBOR360 8	MIBOR360 16
VR (q)	1.00	1.02	0.97	0.88	1.23	1.43	1.49	1.48	0.77	0.76	0.74	0.78
Z (q)	-0.14	0.35	-0.28	-0.87	4.69	4.68	3.42	2.24	-4.73	-2.64	-1.78	-1.03
Z* (q)	-0.75	1.18	-0.30	-0.34	1.21	1.18	0.83	0.50	-1.28	-0.70	-0.45	-0.24

VR(q), variance ratio, is calculated as $\sum\sigma^{\sigma}_{c}(q)/\sum\sigma^{2}_{a}(q)$, where $\sum\sigma^{\sigma}_{c}(q)$ is an unbiased estimator of 1/q of the variance of the qth difference of prices and $\sum\sigma^{2}_{a}(q)$ is an unbiased estimator of the variance of the first difference of price.

Z (q): Standard-normal-distributed homoscedasticity test statistic.

Z* (q): Standard-normal-distributed heteroscedasticity-adjusted test statistic.

variety of efficiency tests were conducted. Serial correlation tests that utilize heteroscadasticity-adjusted variance estimators show that, in general, there are zero serial correlations for the six Spanish futures contracts analyzed in this study. Two different types of unit root tests prove the existence of unit roots in the price series of Spanish futures contracts. Finally, powerful variance-ratio tests support the random walk hypothesis for futures contracts trading on the MEFF.

The overwhelming evidence presented in this paper suggests that the MEFF futures markets are efficient. Considering its relatively short history, this evidence of efficiency in the MEFF futures markets is remarkable, because some studies on other emerging futures markets (see, e.g., Bühler and Kempf ,1995), and Martikainen et al. (1995) have shown that the assumption of the random walk hypothesis is violated, thus providing evidence of inefficiency in those markets. The evidence presented in this paper helps explain why the MEFF futures markets have experienced phenomenal growth in its short history since inception. Are factors such as its computerized trading system contributing to its efficiency? Questions like this are interesting and need to be addressed in future studies.

APPENDIX

This Appendix provides details of the variance ratio tests. The variance ratio of q-differenced series is defined as

$$VR(q) = \frac{\Sigma\sigma_c^2(q)}{\Sigma\sigma_a^2(q)} \tag{1}$$

where $\Sigma\sigma_c^2(q)$ is an unbiased estimator of 1/q of the variance of the qth-differenced series and $\Sigma\sigma_a^2(q)$ is an unbiased estimator of the variance of the first-differenced series. More specifically,

$$\Sigma\sigma_c^2(q) = \frac{1}{m}\sum_{t=q+1}^{nq+1} (p_t - p_{t-1} - q\mu)^2 \tag{2}$$

where

$$m = q(nq - q + 1)(1 - 1/n) \text{ and } \mu = \frac{1}{nq}(p_{nq+1} - p_1) \tag{3}$$

and

$$\Sigma\sigma_a^2(q) = \frac{1}{nq - 1} \sum_{t=2}^{nq+1} (p_t - P_{t-1} - \mu)^2 \tag{4}$$

The standard Z test statistic is

$$Z(q) = \frac{VR(q) - 1}{[\phi(q)]^{1/2}} \tag{5}$$

where $\phi(q) = [2\ (2q - 1)\ (q - 1)]/[3q(nq)]$.

Lo and MacKinlay propose a refined test statistic, Z*(q), which adjusts for heteroscedasticity:

$$Z^*(q) = \frac{VR(q) - 1}{[\phi^*(q)]^{1/2}} \tag{6}$$

where

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

and

$$\delta(j) = \frac{\sum_{t=j+2}^{nq+1} (p_t - p_{t-1} - \mu)^2\, (p_{t-j} - p_{t-j-1} - \mu)^2}{\left[\sum_{t=2}^{nq+1} (p_t - p_{t-1} - \mu)^2\right]^2}$$

Both Z(q) and Z*(q) are asymptotically normally distributed with mean zero and standard deviation 1.

BIBLIOGRAPHY

Alonso, A., & Rubio, G. (1990). Overreaction in the Spanish Equity Market. Journal of Banking and Finance, 14, 469–481.

Annuar, M.N., Ariff, M., & Shamsher, M. (1994). Is Kuala Lumpur's Emerging Share Market Efficient? Journal of International Financial Markets, Institutions & Money, 4(1/2), 89–100.

Arshanapalli, B., & Doukas, J. (1994). Common Stochastic Trends in a System of Eurocurrency Rates. Journal of Banking and Finance, 18, 1047–1061.

Ayadi, O.F., & Pyun, C.S. (1994). An Application of Variance Ratio Test to the Korean Securities Market. Journal of Banking and Finance, 18, 643–658.

Bachelier, L. (1967). Theorie de la speculation. Annales de l'Ecole normale Superieure (3rd. series, 17, 21–86). Translated by A.J. Boness in P.H. Cootner

(Ed.), The random character of stock market prices, Cambridge, MA: MIT Press.

Barone, E. (1990). The Italian Stock Market: Efficiency and Calendar Anomalies. Journal of Banking and Finance, 14, 483–510.

Bühler, W., & Kempf, A. (1995). DAX Index Futures: Mispricing and Arbitrage in German Markets. The Journal of Futures Markets, 15, 833–859.

Cowles, A., & Jones, H. (1937). Some Posteriori Probabilities in Stock Market Action. Econometrica, 5, 280–294.

Cyr, D., & Llewellyn, T. (1994). A Time Series Test of Calendar Seasonalities in the S&P 500 Index Since the Introduction of Index Derivative Securities. The Journal of Futures Markets, 14, 511–529.

Dickey, D.A., & Fuller, W.A. (1979).Distribution of the Estimators for Autoregressive Time Series with A Unit Root. Journal of American Statistical Association, 74 (366), 427–431.

Dickey, D.A., & Fuller, W.A. (1981). Likelihood Ratio Statistics for Autoregressive Time Series with a Unit Root. Econometrica, 49, 1057–1072.

Diebold, F.X. (1986). Modelling the Persistence of Conditional Variance: Comment. Econometric Reviews, 5, 51–56.

Fama, E.F. (1965). The Behavior of Stock Market Prices. Journal of Business, 38, 34–105.

French, K.R. (1980). Stock Returns and the Weekend Effect. Journal of Financial Economics, 8, 55–69.

French, K.R., & Roll, R. (1986). Stock Return Variances: The Arrival of Information and the Reaction of Traders. Journal of Financial Economics, 17, 5–26.

Hsieh, D.A. (1988). The Statistical Properties of Daily Foreign Exchange Rates: 1974–1983. Journal of International Economics, 24, 129–145.

Keim, D.B. (1983). Size Related Anomalies and Stock Return Seasonality. Journal of Financial Economics, 12, 13–32.

Kendall, M. (1953). The Analysis of Economic Time Series, Part I: Prices. Journal of the Royal Statistical Society, 96, 11–25.

Kwiatkowski, D.P., Phillips, C.B., Schmidt, P., & Shin, Y. (1992). Testing the Null Hypothesis of Stationarity Against the Alternative of a Unit Root: How Sure Are We That Economic Time Series Have a Unit Root? Journal of Econometrics, 54, 159–178.

Liu, C.Y., & He, J. (1991). A Variance-Ratio Test of Random Walks in Foreign Exchange Rates. Journal of Finance, 46, 773–785.

Lo, A.W., & MacKinlay, C. (1988). Stock Market Prices Do Not Follow Random Walks: Evidence from a Simple Specification Test. Review of Financial Studies, 1, 41–66.

Martikainen, T., Perttunen, J., & Puttonen, V. (1995). Finnish Turn-of-the-Month Effects: Returns, Volume, and Implied Volatility. The Journal of Futures Markets, 15, 605–615.

Merton, R. (1980). On Estimating the Expected Return on the Market: An Exploratory Investigation. Journal of Financial Economics, 8, 323–361.

Rubio, G. (1989). An Empirical Evaluation of the Intertemporal Capital Asset Pricing Model: The Stock Market in Spain. Journal of Business Finance and Accounting, 16, 729–743.

Szakmary, A.C., Mathur, I., & Yu, S.H. (1995). Overshooting Behavior in Foreign Exchange Markets: Evidence from Cointegration Tests. Research in International Business and Finance, 12, 117–132.

Working, H. (1934). A Random Difference Series for Use in the Analysis of Time Series. Journal of the American Statistical Association, 29, 11–24.