A Cointegration Test for Market Efficiency

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INTRODUCTION

The efficiency of the forward or futures market is often examined based on the model

$$S_t = a + bF_{t-1,t} + \epsilon_t \tag{1}$$

where S_t is the spot price at time t, $F_{t-1,t}$ is the price at time $t - 1$ for the forward or futures contract maturing at time t, ϵ_t is an error term with mean zero and finite variance, and a and b are constant coefficients. Under the hypothesis of market efficiency, the market price should fully reflect available information so that there exists no strategy from which traders can profit consistently by speculating in the forward or futures market on future levels of the spot price. Efficiency, so defined, implies a testable restriction that $a = 0$ and $b = 1$ in eq. (1), which is generally referred to as the unbiasedness hypothesis. This hypothesis is called the "simple efficiency" hypothesis by Hansen and Hodrick (1980) and the "speculative efficiency" hypothesis by Bilson (1981), since the test for unbiasedness represents a joint test of market efficiency and no-risk premium.[1]

In testing the parameter restriction in eq. (1), the issue arises regarding whether or not the price series is stationary. The stationarity property is important, since asymptotic distribution theory invoked to construct a test of the hypothesis relies critically upon it. Elam and Dixon (1988) observe that financial price series are generally found to be not stationary and they contain a unit root. As a result, the standard F-test of the hypothesis $a = 0$ and $b = 1$ is no longer appropriate. Using Monte Carlo experiments, Elam and Dixon illustrate that the F-test tends to bias toward incorrectly rejecting market efficiency.

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[1]The notion of "simple efficiency," defined by Hansen and Hodrick (1980), or of "speculative efficiency," defined by Bilson (1981), means that traders have rational expectations and charge no risk premium in the forward exchange market. Bilson (1981) notes that if a market is speculatively efficient, the supply of speculative funds is infinitely elastic at the forward price that equals the expected future spot price.

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In response to Elam and Dixon (1988), Shen and Wang (1990) suggest that the technique of cointegration developed by Engle and Granger (1987) may be used to test for market efficiency. The cointegration approach is attractive in that it can properly account for the nonstationarity in price series. Following Engle and Granger (1987), a test for an equilibrium relationship between S_t and $F_{t-1,t}$ proceeds as follows: estimate eq. (1) as the cointegrating or equilibrium regression, and check its least squares residual for stationarity using a unit-root test. If the residual is found to be stationary, the null hypothesis of no equilibrium relationship between S_t and $F_{t-1,t}$ is rejected. A limitation of the Engle–Granger procedure is that no strong statistical inference can be drawn with respect to the parameters a and b which are of main interest here. Although the coefficient estimator can be shown to be consistent, the estimated standard errors may be misleading for hypothesis testing (see Stock, 1987).[2]

Recent developments in the cointegration analysis by Johansen (1988,1990) provide a new technique for testing market efficiency. Johansen devises a statistical procedure for testing cointegration using the maximum likelihood method. The procedure allows one to formally conduct likelihood ratio tests of the parameters of the equilibrium relationship between nonstationary variables. In contrast to the Engle–Granger single equation procedure, the Johansen procedure is based on a vector autoregressive model that allows for possible interactions in the determination of spot prices and forward or futures prices.[3]

This study suggests the use of the Johansen statistical procedure to test for market efficiency. The statistical analysis is illustrated with an application of the technique to examine the efficiency of the forward currency market.

COINTEGRATION AND MARKET EFFICIENCY

A time series is integrated of order d, denoted $I(d)$, if the series can achieve stationarity only after differencing d times. An $I(0)$ series is thus, by definition, stationary; whereas, an $I(1)$ series contains a unit root and is nonstationary. The simplest example of an $I(1)$ series is a random walk.[4] When the spot price, S_t, and the forward or futures price, $F_{t-1,t}$, are both $I(1)$, the linear combination

$$z_t = S_t - a - bF_{t-1,t} \tag{2}$$

is generally also $I(1)$. However, if there exist a and b such that z_t is stationary or $I(0)$, then S_t and $F_{t-1,t}$ are said to be cointegrated, and the relationship

$$S_t - a - bF_{t-1,t} = 0 \tag{3}$$

is the cointegrating or equilibrium relationship with z_t in eq. (2) representing the equilibrium error (Engle and Granger, 1987).

Cointegration between S_t and $F_{t-1,t}$ is a necessary condition for market efficiency. The hypothesis of market efficiency suggests that $F_{t-1,t}$ is an unbiased predictor of S_t on average, i.e., the forward or futures prices do not consistently over- or under-

[2]Stock (1987) suggests a way to "correct" the estimated standard errors. The statistical test is, however, very sensitive to the nuisance parameters of the underlying series.

[3]Since the Johansen procedure takes into account the error structure of the underlying date generating process, the procedure can provide potentially more precise parameter estimates than the Engle–Granger procedure.

[4]Elam and Dixon (1988) and Shen and Wang (1990) discuss the problem in testing market efficiency when the spot price follows a random walk. Usual F-tests are not valid as long as $I(1)$ series are considered in general.

predict the spot prices. If S_t and $F_{t-1,t}$ are not cointegrated, z_t is nonstationary and S_t and $F_{t-1,t}$ tends to deviate apart without bound. It follows that $F_{t-1,t}$ has little predictive power about the movement of S_t. This is clearly inconsistent with the market efficiency hypothesis.

The cointegration property is, however, only one of the necessary conditions for market efficiency. Market efficiency also requires that $a = 0$ and $b = 1$ in eq. (3), otherwise, $F_{t-1,t}$ is not an unbiased predictor of S_t, even when S_t and $F_{t-1,t}$ move "closely" together over time. Hence, a test for market efficiency involves formal testing of restrictions on the cointegrating parameters. In this regard, the use of the conventional Engle–Granger procedure, as suggested by Shen and Wang (1990), is not appropriate because hypothesis tests on the cointegrating parameters under this procedure do not follow any standard distribution. In contrast, the Johansen procedure, applied in this study, can handle the problem of statistical inference in cointegrated systems. As discussed below, hypothesis tests on the cointegrating parameters, namely $a = 0$ and $b = 1$ in eq. (3), can be conducted using standard asymptotic chi-square tests under the Johansen approach.[5]

In short, market efficiency implies that S_t and $F_{t-1,t}$ are cointegrated, and for the cointegrating parameters, $a = 0$ and $b = 1$. The test for market efficiency thus consists of two related parts. The nonstationary series S_t and $F_{t-1,t}$ are first examined for cointegration. If they are found to be cointegrated, the restriction on the cointegrating parameters that $a = 0$ and $b = 1$ is then tested under the condition of cointegration using a likelihood ratio test. The statistical procedure is discussed next.[6]

EMPIRICAL METHODOLOGY

The statistical procedure used by Johansen (1988, 1990) is outlined here (see also Johansen and Juselius, 1990). The Johansen procedure provides a likelihood ratio test for cointegration in terms of a vector autoregressive (VAR) model that can incorporate different short- and long-run dynamics of a system of economic variables. The cointegration technique enables one to estimate and test the equilibrium relationship among nonstationary series while abstracting from short-run deviations from equilibrium. The Johansen approach is based on the multivariate technique of canonical correlations (see, e.g., Anderson, 1984). Intuitively, the canonical correlation analysis is to find a linear combination of a set of variables such that the correlation among the variables is maximized. Johansen shows that the hypothesis of cointegration can be formulated as the hypothesis of reduced rank of a regression coefficient matrix, which can be estimated consistently from two vector regression equations. Based on these regressions, the likelihood ratio test for cointegration involves computing the squared canonical correlations between the regression residuals, which require calculation of eigenvalues. Inferences on the cointegrating parameters under linear restrictions can be conducted using the chi-square distribution.

[5]Two recent studies by Baillie and Bollerslev (1989) and Hakkio and Rush (1989) examine the cointegration property between forward and spot exchange rates using the Engle–Granger procedure. However, these studies do not formally test the unbiasedness condition.

[6]Strictly speaking, market efficiency also requires the equilibrium error z_t to be a white noise, whereas cointegration requires z_t to be stationary only, i.e., a weaker condition than a white noise. Nonetheless, cointegration and the condition that $a = 0$ and $b = 1$ are, at least, necessary conditions for market efficiency.

Consider a general VAR model that is written as

$$(1 - L)X_t = \Gamma_1(1 - L)X_{t-1} + \ldots + \Gamma_{k-1}(1 - L)X_{t-k-1} + \Gamma_k X_{t-k} + \mu + v_t \quad (4)$$

where X_t is an $n \times 1$ time series vector, L is the lag operator, μ is some constant vector, and v_t is a vector of white Gaussian noises with mean zero and finite variance. Johansen (1989, 1990) shows that the coefficient matrix Γ_k contains the essential information about the cointegrating or equilibrium relationship between the variables in the data set. Specifically, the rank of the matrix Γ_k indicates the number of cointegrating relationships existing between the variables in X_t. In this case, $X_t = (S_t, F_{t-1,t})'$ and so $n = 2$. The hypothesis of cointegration between S_t and $F_{t-1,t}$ is equivalent to the hypothesis that the rank of $\Gamma_k = 1$. If the rank $= 0$, then the two variables are not cointegrated.

The statistical analysis proceeds with the two simple least squares regressions

$$(1 - L)X_t = \Gamma_1(1 - L)X_{t-1} + \ldots + \Gamma_{k-1}(1 - L)X_{t-k-1} + u_{1t} \quad (5a)$$

$$X^*_{t-k} = \Gamma_1(1 - L)X_{t-1} + \ldots + \Gamma_{k-1}(1 - L)X_{t-k-1} + u_{2t} \quad (5b)$$

where $X^*_{t-k} = (X'_{t-k}, 1)'$. Define the product moment matrices of the residuals, $\hat{u}_{1t}$ and $\hat{u}_{2t}$, as $S_{ij} = T^{-1} \Sigma_{t=1}^{T} \hat{u}_{it}\hat{u}_{jt}'$ for $i, j = 1, 2$. The likelihood ratio test statistic for the hypothesis of at most r equilibrium relationships can be shown to be given by

$$-2\ln Q_r = -T \sum_{j=r+1}^{n} \ln(1 - \phi_j) \quad (6)$$

where $\phi_1 > \ldots > \phi_n$ are the eigenvalues of $S_{21}S_{11}^{-1}S_{12}$ with respect to S_{22}. The eigenvalues are also called the squared canonical correlations of $\hat{u}_{2t}$ with respect to $\hat{u}_{1t}$. The limiting distribution of the $-2\ln Q_r$ statistic is given in terms of an $(n - r)$-dimensional Brownian motion process, and the quantiles of the distribution are tabulated in Johansen and Juselius (1990) using simulations. Moreover, the maximum likelihood estimate of the cointegrating vector, given by α of the equilibrium relationship

$$\alpha' X^*_t = 0 \quad (7)$$

can be computed as the first eigenvector of $S_{21}S_{11}^{-1}S_{12}$ with respect to S_{22}.

It is useful to note that eq. (7) includes eq. (3) as a special case where $X^*_t = (S_t, F_{t-1,t}, 1)'$ and $\alpha' = (1, -b, -a)$, which normalizes the coefficient of the S_t variable to be unity. A test for market efficiency can, therefore, be conducted as a test of a linear restriction on the cointegrating vector, namely the restriction that $\alpha' = (1, -1, 0)$.

Following Johansen's analysis, the hypothesis of linear constraints on the cointegrating vector α can generally be represented by

$$\mathrm{H}_G : \alpha = G\pi \quad (8)$$

where G is an $3 \times m$ matrix of full rank m, specified by the researcher, and π is an $m \times r$ matrix of unknown parameters ($3 \geq m \geq r$). When only one equilibrium relationship is considered, $r = 1$. In addition, the condition for market efficiency can be described in terms of $G = (1, -1, 0)$. In this case, G is a 3×1 vector and, accordingly, π is a scalar, which can be interpreted as some normalization constant.

Johansen (1988, 1990) and Johansen and Juselius (1990) demonstrate that the likelihood ratio test statistic for H_G is, in general, given by

$$-2\ln Q_G = -T\sum_{j=1}^{r} \ln\{(1 - \phi_j)/(1 - \phi_j^*)\} \tag{9}$$

where $\phi_1^* > \ldots > \phi_r^*$ are the eigenvalues of $G'S_{21}S_{11}^{-1}S_{12}G$ with respect to $G'S_{22}G$. Under the null hypothesis that the restriction H_G is valid, the limiting distribution of $-2\ln Q_G$ can be shown to be a chi-square distribution with a degree of freedom equal to the number of constraints being tested.

AN APPLICATION TO FORWARD CURRENCY MARKETS

Consider monthly spot and forward rates for five major currencies (against the U.S. dollar): the British pound (BP), Deutsche mark (DM), Swiss franc (SF), Canadian dollar (CD), and Japanese yen (JY). The entire sample, from July 1973 to December 1989, consists of 198 observations for each series. The data used are nonoverlapping, and they are end-of-the-month observations on the (bid) spot exchange rate and the (bid) one-month forward rate, drawn from the *International Monetary Market Yearbook* for the 1973–1987 period and from *The Wall Street Journal* for the 1988–1989 period. If the last day of the month is a weekend or public holiday, the previous working day is chosen which is consistent with the calendar month settlement arrangement of the foreign exchange market.[7] All series are expressed in terms of U.S. currency units and regressions are based on logarithms of the variables.

While forward exchange rates are considered in the present illustration, the statistical procedure can apply also to study futures rates as predictors of future spot rates. It is known that futures contracts differ in some ways from forward contracts. Forward rates are quoted in an interbank market for delivery at a fixed maturity time, usually one, three, six, or 12 months ahead. In contrast, futures quotes are obtained at an organized exchange, i.e., International Monetary Market in Chicago, and futures contracts are traded for delivery at a fixed maturity date, e.g., the third Wednesday of March, June, September, or December. Since futures contracts do not mature every month, the use of futures data can greatly limit the sample size unless overlapping contracts are considered. Moreover, while in theory, forward prices and futures prices can differ for delivery on the same day in the future (e.g., Cox, Ingersoll, and Ross 1980; Jarrow and Oldfield, 1981), empirical work by, e.g., Cornell and Reinganum (1981) find very little difference between the two prices in foreign exchange markets.

Each series of the spot rate and forward rate is first checked for a unit root using the augmented Dickey–Fuller (ADF) test (Dickey and Fuller, 1979) and the Phillips–Perron Z_t test (Phillips, 1987; Phillips and Perron, 1988). The ADF(p) test statistic for a series x_t is given by the t-statistic, t_a, for a in

$$(1 - L)x_t = a_0 + ax_{t-1} + a_1(1 - L)x_{t-1} + \ldots + a_p(1 - L)x_{t-p} + w_t \tag{10}$$

[7]For some month-end observations, the one-month forward rate may not exactly be matched up with the spot rates when the forward contract matures. In those cases, there may be a discrepancy of one business day. Nevertheless, the discrepancy is not systematic across observations, and it does not seem to be significant.

For the Z_t test, regress $(1 - L)x_t = b_0 + bx_{t-1} + e_t$, and compute

$$Z_t(q) = (s_o/s_{Tq})t_b - (1/2)(s_{Tq}^2 - s_0^2)\left[T^{-1}s_{Tq}\left(\sum_{t=2}^{T} x_{t-1}^2\right)^{1/2}\right]^{-1} \tag{11}$$

where T is the sample size, $s_0^2 = T^{-1}\Sigma_{t=1}^{T}\hat{e}_t^2$ and s_{Tq}^2 is a consistent variance estimator given by

$$s_{Tq}^2 = T^{-1}\sum_{t=1}^{T}\hat{e}_t^2 + 2T^{-1}\sum_{\tau=1}^{q}\sum_{t=\tau+1}^{T}[1 - \tau/(q + 1)]\hat{e}_t\hat{e}_{t-\tau} \tag{12}$$

Table I contains the results of the unit-root tests for $p, q = 1,3$, and 5. The results are apparently not sensitive to the choice of the lag parameters p and q. While the unit-root null hypothesis cannot be rejected for all level series even at the 10%

Table I
STATISTICS FOR UNIT-ROOT TESTS ON EXCHANGE RATE SERIES

Series	Monetary Unit	The Dickey–Fuller ADF(p) Test			The Phillips–Perron $Z_t(q)$ Test		
		$p = 1$	$p = 3$	$p = 5$	$q = 1$	$q = 3$	$q = 5$
Levels:							
Spot							
	BP	−1.698	−1.659	−1.716	−1.747	−1.777	−1.803
	DM	−0.922	−1.278	−1.452	−0.991	−1.111	−1.161
	SF	−1.589	−1.964	−2.132	−1.591	−1.649	−1.674
	CD	−1.452	−1.512	−1.674	−1.458	−1.453	−1.457
	JY	−0.445	−0.685	−1.104	−0.426	−0.480	−0.564
Forward							
	BP	−1.701	−1.738	−1.711	−1.771	−1.803	−1.830
	DM	−1.067	−1.340	−1.566	−1.093	−1.213	−1.277
	SF	−1.622	−1.895	−2.184	−1.434	−1.511	−1.547
	CD	−1.474	−1.510	−1.584	−1.513	−1.508	−1.510
	JY	−0.423	−0.624	−1.029	−0.360	−0.434	−0.523
First Differences:							
Spot							
	BP	−8.900	−6.612	−5.390	−13.479	−13.500	−13.521
	DM	−8.432	−6.648	−5.351	−14.648	−14.644	−14.662
	SF	−8.452	−6.629	−5.716	−14.149	−14.170	−14.194
	CD	−10.857	−6.573	−5.093	−14.965	−14.989	−14.948
	JY	−9.333	−5.948	−5.278	−13.644	−13.659	−13.718
Forward							
	BP	−8.835	−6.523	−5.474	−13.342	−13.363	−13.391
	DM	−8.508	−6.621	−5.640	−14.995	−14.959	−14.965
	SF	−8.383	−6.586	−5.830	−14.334	−14.339	−14.360
	CD	−11.109	−6.722	−5.094	−15.242	−15.085	−15.040
	JY	−9.094	−5.965	−5.502	−13.277	−13.304	−13.379

Notes: Critical values are tabulated in Fuller (1976, p. 373); for a sample size of 200, they are given by −2.57 (10%), −2.88 (5%), and −3.46 (10%). The last ten rows correspond to results from tests on first differences of the series.

For "First Differences," all statistics are significant at the 1% level.

significance level, tests performed on first differences strongly indicate that the first difference of each series is stationary. Unit-root tests that allow for a time trend are conducted also. They yield qualitatively the same results (not reported), and the time trend variables are found to be statistically insignificant. The results, on the whole, indicate that the spot and forward exchange rate series examined are all $I(1)$.[8]

The Johansen test is next performed to detect cointegration between the spot exchange rate and the forward exchange rate. The efficient pricing condition, $a = 0$ and $b = 1$, is then tested as a restriction on the cointegrating vector $\alpha = (1, -1, 0)$. In addition, a test is conducted to see if $b = 1$, i.e., $\alpha = (1, -1, -a)$, individually. If $b = 1$ and $a \neq 0$, forward prices can explain movements in spot prices, though forward prices are biased forecasts of future spot prices (see Martin and Garcia, 1981).

The statistical results of cointegration tests are reported in Table II, which illustrates that null hypothesis of no cointegrating relationship, i.e., $r = 0$, is rejected at the 1% level for all currencies under consideration. The results uniformly suggest that S_t and $F_{t-1,t}$ are cointegrated. The estimates of the cointegrating vector $\alpha = (1, -b, -a)$ normalize the coefficient on S_t to be unity, and the implied equilibrium relationship is given by $S_t - bF_{t-1,t} - a = 0$. The reported estimates for a and b appear rather close to 0 and 1, respectively. In testing the hypothesis $b = 1$, the $-2\ln Q_G$ test statistic has a chi-square distribution with one degree of freedom. In no case can the hypothesis of $b = 1$ be rejected statistically, even at the 10% level. However, tests of the hypothesis $a = 0$ and $b = 1$, indicate that this hypothesis is rejected for all five currencies at the 5% significance level or better; since, in this case, the corresponding $-2\ln Q_G$ test statistic has a chi-square distribution with two degrees of freedom. Hence, while the forward exchange rate seems able to explain movements in the spot exchange rate in the sense of Martin and Garcia (1981), the forward rate appears to be a biased predictor of the future spot rate.

Table II
RESULTS OF COINTEGRATION TESTS

Series	k	Parameter Estimates for $\alpha = (1, -b, -a)$	The $-2\ln Q_r$ Statistic for H: $r = 0$	Hypothesis Tested on α: H: $b = 1$	Hypothesis Tested on α: H: $a = 0$ and $b = 1$
BP	2	(1, −1.0024, −0.0014)	36.75[a]	0.6226	15.472[a]
DM	2	(1, −0.9975, −0.0046)	60.63[a]	2.2916	36.738[a]
SF	3	(1, −0.9970, 0.0065)	28.06[a]	1.4271	28.784[a]
CD	1	(1, −1.0012, −0.0012)	73.17[a]	2.5709	37.644[a]
JY	2	(1, −0.9940, 0.0340)	33.15[a]	2.5709	5.129[b]

Notes: The optimal lag, k, is selected based on Schwarz's (1978) information criterion using a maximum lag of 8. The Schwarz criterion tells us to choose the model so as to maximize $-2\ln L - N\ln T$, where L is the likelihood, N the number of the model parameters, and T the number of observations. The $-2\ln Q_r$ statistic is for testing the null hypothesis of no cointegration or $r = 0$; its critical values are given by 19.796 (10%), 21.894 (5%), and 26.409 (1%), based on Johansen and Juselius (1990, Table A3). The statistics for testing the hypothesis $b = 1$ and the hypothesis $a = 0$ and $b = 1$ have a chi-square distribution with the respective degree of freedom equal to one and two.

[a]Significant at the 1% level.

[b]Significant at the 5% level.

[8]Formal evidence that spot and forward exchange rate series generally contain a unit root is reported also in Corbae and Ouliaris (1986) and Meese and Singleton (1982).

The findings reported here are apparently not supportive of the unbiasedness hypothesis. As mentioned earlier, the evidence can be interpreted as a violation of market efficiency, or of a risk premium, or both. The joint hypothesis of market efficiency and no-risk premium has yet to be separated into testable alternative hypotheses. The same point applies to most of the previous empirical results in the literature on market efficiency. A recent study by Lewis (1989) argues also that the forward prediction bias can be consistent with the hypothesis that the market is learning rationally about a process shift in the exchange rate process. Lewis presents empirical evidence that about one half of the underprediction of the dollar value implied by the forward market during the early 1980s can be accounted for as Bayesian forecast errors. While learning about a change in the exchange rate process may possibly explain the forward rate bias for some currencies during specific episodes, its relevance for explaining the prolonged periods of systematic forward rate bias of many different currencies is likely to be limited.

CONCLUSION

A problem in testing forward or futures market efficiency is that financial price series are generally not stationary. When the series are nonstationary, conventional statistical procedures are no longer valid in providing a test for market efficiency. Hence, development of an appropriate statistical procedure that can account for the nonstationarity in the price series is important. This study suggests the use of the cointegration technique, recently developed by Johansen (1988, 1990), to test the "simple efficiency" or "speculative efficiency" hypothesis. The analysis is illustrated with an application to five major forward currency markets, and evidence is found not favorable to the joint hypothesis of market efficiency and no-risk premium.

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