Cointegration, Unbiased Expectations, and Forecasting in the BIFFEX Freight Futures Market

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The relationship between freight cash and futures prices is investigated using cointegration econometrics. Results illustrate that the BIFFEX futures market is unbiased, and hence efficient for the current, one, two, and quarterly contract horizons. Since the futures contract is based on an index of various shipping routes, which has undergone several changes since its inception, stability in the relationship between the spot and futures rates is investigated using rolling cointegration techniques. Results indicate that the futures contract appears to have become more efficient over time in predicting the spot rate, and that the decrease in trading volume found in the BIFFEX market is not driven by a lack of efficiency in this market. Rather, the decrease in futures trading might be attributed to the growth rate of the freight forward market. This article incorporates the long-run cointegrating relationships between cash and futures prices in a forecasting model and compares the forecasting performance of this model with several alternatives. It is found that while the futures price is the best predictor of future spot rates for the

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current-month contract, time-series models can outperform the futures contract at longer contract horizons. Jrl Fut Mark 20:545–571, 2000.

INTRODUCTION

When the Baltic Freight Futures Contract was launched in May 1985, it was unique in many respects. It was the first futures contract encompassing nearly every part of the globe, and the first contract designed to respond to erratic freight rate fluctuations in the shipping industry. The futures contract was also unique in the sense that it was the first futures contract designed for a service rather than representing an index or a product.[1] However, the volume of trading activity remains low and the number of traders limited. This has disappointed advocates of the use of derivatives in the shipping market. Since the futures contract is currently based on an index of various shipping routes known as the Baltic Panamax Index (BPI) (formerly known as the Baltic Freight Index [BFI]), it often is stated that the freight futures contract has failed to reflect accurately the market conditions on lightly weighted routes. This, industry experts suggest, is the reason why tailor-made, over-the-counter (OTC) derivatives like Freight Forward Agreements (FFAs) were introduced in 1992, and have realized relatively high growth rates in trading activity. Despite this, the freight futures market still exists after fifteen years of trading.

One area of research related to the Baltic International Freight Futures Exchange (BIFFEX) freight futures contract has focused on issues relating to efficiency and forecasting future spot freight rates.[2] For instance, Chang (1991) employed simple regression analysis to test the forward pricing function of freight futures prices. The study concluded that all information was being efficiently utilized in the freight futures market, reinforcing the nonrejection of the unbiasedness hypothesis. Denning, Riley, and Delooze (1994) adopted a variance ratio methodology to examine BIFFEX price, volume, and index quotes for randomness. They found possible evidence of seasonality in the index values, suggest-

[1]At the same time, trading in an identical contract (based on electronic trading, rather than open outcry) also started at the then-existing International Futures Exchange (INTEX), based in Bermuda. Trading at the INTEX ceased because of insufficient trading volume. On February 10, 1986, an attempt also was made to create a tanker market similar to the freight futures contract, named TIFFEX. A major reason for its eventual failure was a lack of volume.

[2]Another area of research has concentrated on the effectiveness of the BIFFEX freight futures contract in a risk-management/hedging setting. A partial listing includes: Thoung and Visscher (1990), Hauser and Neff (1993), and Haigh and Holt (1999, 2000).

ing the possibility of the potential for exploitation by traders. Moreover, the Denning et al. paper presents a useful application of a multiple comparison procedure that is used simultaneously and conveniently to test several hypotheses at the same time.[3] Chang and Chang (1996), looked at the predictability of the dry bulk shipping market using the BIFFEX. Their research utilized univariate time-series analysis, and concluded that the BIFFEX freight futures contract is unbiased at predicting future spot rates up to a one-month horizon. However, according to their research, performance declined dramatically after one-month, and after a six-month horizon the BIFFEX was unable to forecast future spot rates. Their research, like the two previous papers, did not incorporate any long-run relationship between the future and spot prices in the analysis. Kavussanos and Nomikos (1999) used cointegration analysis to test for unbiasedness in the BIFFEX freight futures market, thus incorporating the long-run equilibrium information. They reported evidence of unbiasedness at nearby contract lengths and showed support for rejecting unbiasedness for the relatively popular quarterly contract. Unfortunately, the authors only used a portion of available data, so many of the results presented in the paper seem to have been driven by the small sample sizes used.[4]

The current article has several objectives. First, an evaluation is made of to the extent to which the freight futures price can predict accurately the future spot rate. To this end, Johansen's (1988, 1991) cointegration and error-correction techniques are adopted to test the unbiasedness and efficiency hypothesis in this unique futures market using the most complete data set available. The second purpose of the study is to evaluate the stability of the cointegrating relationship over time. This is of particular interest as there have been several changes in the index since 1985 (in order to attract more futures trading), and as the futures contract is based on the index; it is reasonable to believe that the relation-

[3]As pointed out by an anonymous reviewer, it is unclear how papers that utilize cointegration analysis for unbiasedness testing account for and properly control for the multiplicity effect through some combined or joint measure of inferences (multiple-comparison procedures). While accounting for such an issue is beyond the scope of this paper, the author recognizes problems associated with making several nonindependent hypothesis tests. However, care was taken throughout this analysis to assess each hypothesis test separately using stringent critical values deemed appropriate for that inference. See Hochberg and Tamhane (1987) for further information on multiple-comparison procedures.

[4]At the time of this writing, the current article was the first known research using cointegration analysis to test for unbiasedness in the BIFFEX freight futures contract. Several months after this article was submitted, a similar article by Kavussanos and Nomikos (1999) also was submitted to the same journal. While the Kavussanos and Nomikos article researched a similar issue with regard to the BIFFEX, the unbiasedness results obtained by the authors should be interpreted with caution because of the small sample sizes used.

ship between the futures and spot prices may have changed. Therefore, a rolling cointegration technique is employed to test the stability of the cointegrating relationship over time. Results of the rolling cointegration analysis are tied in with issues of efficiency and liquidity within this market, and recent market developments are discussed based upon this analysis.

The last section of the paper builds upon previous work by Cullinane (1992), Chang and Chang (1996), Denning et al. (1994), Veenstra and Franses (1997), and Kavussanos and Nomikos (1999) in developing several different time-series models that forecast future freight rates. Importantly, the approach in this paper differs from most by incorporating the monthly cointegrating information developed in the first section of the paper in the forecasting model. A comparison is made between performance of the error-correction model (ECM), which incorporates the long-run information to a Vector Autoregressive Model in levels [VAR (levels)], a Vector Autoregressive Model in first differences [VAR (diff)], and the actual futures price in predicting future spot rates.

COINTEGRATION, EFFICIENCY, AND UNBIASEDNESS

The concept of unbiasedness suggests that the current futures price, F_{t-n}, of a contract expiring at time t, should equal the spot price expected to prevail at time t. This unbiasedness hypothesis also can be interpreted as the "simple efficiency" hypothesis proposed by Hansen and Hodrick (1980) and the "speculative efficiency" hypothesis proposed by Bilson (1981), since the test for unbiasedness represents a joint test of market efficiency *and* no-risk premium. As such, the hypothesis that futures prices are unbiased predictors of spot prices is a joint hypothesis that markets are efficient and that a risk premium is not present. Rejecting unbiasedness could therefore be caused by the failure of efficiency, or the presence of a risk premium. Accepting the null hypothesis of unbiasedness is an acceptance of market efficiency with no risk premium. Should this be the case, long-run market efficiency and long-run unbiasedness can be regarded as synonymous.

Therefore, efficiency/unbiasedness implies that the current futures price is indeed the best forecaster of the expected spot price, and that the current futures price therefore should incorporate all available information. The implication of unbiasedness is that traders can be assured that in general, over time (or in the long run), the current futures price is likely to predict accurately future spot rates without the trader having to pay a risk premium for the privilege of trading the contract. Given the above, in order to test long run efficiency/unbiasedness, there have been several attempts to estimate the following:

$$S_t = \delta_0 + \delta_1 F_{t-n} + \mu_t. \tag{1}$$

That is, the current futures price of an asset, F_{t-n}, should equal the expected spot price, S_t, at time t. Long-run unbiasedness, and hence efficiency, thus is confirmed when the joint restriction $\delta_0 = 0$ and $\delta_1 = 1$ holds. Examples of studies on this relationship include Tomek and Gray (1970) and Kahl and Tomek (1986). However, it has been shown that OLS procedures are inappropriate when nonstationary data are used. The unreliability of test statistics in the presence of unit root leads to invalid inferences. Elam and Dixon (1988) demonstrated, using Monte Carlo experiments, that a standard F test of $\delta_0 = 0$ and $\delta_1 = 1$ is biased toward falsely rejecting the null hypothesis of unbiasedness. A solution to the above-outlined problem is offered by cointegration methodology, which deals specifically with the issue of nonstationary time-series data. That is, when two series are linked by an equilibrium relationship, inference regarding the parameters of the equilibrium relationship can be obtained using cointegration techniques.

The development of cointegration modeling stems largely from the work Granger (1986) and Engle and Granger (1987).[5] A limitation of the Engle–Granger procedure is that no strong statistical inference can be drawn with respect to the parameters, δ_0 and δ_1, which is the crux of the unbiasedness testing. Although, for the Engle–Granger procedure, the estimated coefficients can be shown to be consistent, the associated standard errors may be misleading for any hypothesis testing.[6] The other inherent problem with the Engle–Granger two-step procedure is that the cointegrating vector is assumed to be unique.

A solution to the problems of the Engle–Granger approach has been offered by Johansen's multivariate approach (Johansen, 1988, 1991; Johansen & Juselius, 1990) and has been used widely since by numerous researchers in testing for unbiasedness. A partial listing includes: Fortenbery and Zapata (1997), who rejected the unbiasedness hypothesis in the cheddar cheese market; Lai and Lai (1991), who also failed to find evidence in support of the unbiasedness hypothesis in the forward currency markets; and Antoniou and Foster (1994), who found support for long-term unbiasedness in cocoa and coffee markets, but also found evidence of short-term inefficiencies.

[5]Dickey, Jansen, and Thornton (1991) present a survey of cointegration procedures.
[6]See Hall (1986) and Stock (1987).

Because of the advantages of the Johansen methodology, this technique is adopted in the current analysis. First, assume an n-dimensional vector Y_t that is generated by an autoregressive form depicted as:

$$\mathrm{Y_t} = \mathrm{c} + \sum_{i=1}^{k} \Pi_i \mathrm{Y}_{t-i} + \varepsilon_t, \tag{2}$$

where Y_t is an $n \times 1$ vector of the $I(1)$ variables: spot (S_t), and futures, (F_{t-n}) prices, respectively, in this instance. Π_i is an $n \times n$ matrix of parameters, c is a vector of constants, and ε_t is a random error term. Johansen and Juselius (1990) prove that eq. (2) can be rewritten as an error-correction representation as follows:

$$\Delta \mathrm{Y_t} = \sum_{i=1}^{k-1} \Gamma_i \Delta \mathrm{Y}_{t-i} + \Pi \mathrm{Y}_{t-k} + \varepsilon_t, \tag{3}$$

with

$$\Gamma_i = -(\mathrm{I} - \Pi_1 - \Pi_2 - \ldots - \Pi_i) \quad (i = 1, \ldots\ldots, k-1), \tag{4}$$

and

$$\Pi = -[\mathrm{I} - \Pi_1 - \ldots \Pi_k]. \tag{5}$$

Equation (3) resembles a VAR (diff), with an inclusion of the lagged-level component, otherwise known as the error correction term. Since ε_t is stationary, the rank of the 'long-run' matrix determines how many linear combinations of Y_t are stationary. When Π is full rank ($\Pi = 2$), then S_t and F_{t-n} are jointly stationary and a VAR (levels) is the appropriate model to use to study the relationship between the prices. If the rank is zero, then cointegration between S_t and F_{t-n} will not be present, and the VAR (diff) representation is the appropriate representation. If the rank of Π is positive and less than n, then cointegration is present, and there exist matrices $\alpha\beta'$, with dimensions $n \times r$, such that $\Pi = \alpha\beta'$. The cointegration test statistics involve comparing the number of cointegrating vectors under the null and alternative hypotheses. Specifically, if we denote $\lambda_1', \lambda_2', \ldots, \lambda_n'$ and $\lambda_1^*, \lambda_2^*, \ldots, \lambda_n^*$ as the ordered characteristic roots of the unrestricted and restricted models, respectively then to test the restrictions on β, we can form the test statistic:

$$\mathrm{T} \sum_{i=1}^{r} [\ln(1 - \lambda_i^*) - (1 - \lambda_i')]. \tag{6}$$

Asymptotically, this statistic is distributed as a χ^2 with the number of restrictions imposed on β equaling the number of degrees of freedom.

The matrix β is the matrix of cointegrating parameters and the matrix α contains the adjustment parameters.[7] The α parameters act as weights on each cointegrating vector. $\beta' Y_t$ is interpreted as the long-term equilibrium between the Y_t variables, S_t and F_{t-n}. Because of cross-equation restrictions, α and β cannot be estimated by using OLS. Utilizing maximum-likelihood methods, however, it is possible not only to estimate the error-correction model, but also to establish the rank of Π, and use the r most significant cointegrating vectors to create β'. In addition, it also is possible to select α such that $\Pi = \alpha\beta'$. The Johansen procedure also permits testing of the restrictions on the parameters in the cointegrating vectors. Establishing the definition of stationary linear combinations of nonstationary variables $\beta' Y_t$, one may define:

$$\beta' \mathrm{Y}_t = 0, \tag{7}$$

where $Y_t = (S_t, 1, F_{t-n})$, and $\beta' = (1, -\delta_0, -\delta_1)$. Solving eq. (7) gives the cointegrating relation described in eq. (1). The long-term test for unbiasedness then is undertaken by imposing $\beta' = (1, 0, -1)$, which normalizes S_t to unity and gives the restrictions that $\delta_0 = 0$ and $\delta_1 = 1$.[8]

DATA

The BIFFEX freight futures contract began trading in May 1985 on a quarterly month basis, and in July 1988, spot (current) and prompt (one) months were introduced. In October 1991, another 'prompt' contract was introduced, referred to here as the two-month contract. Therefore, from October 1991 onwards, traders could trade the current month, the next two months, and the quarterly contracts for up to 18 months forward. The data set used in this study spans May 1985 to May 1999. Data from May 1985 to December 1990 was provided by Clarkson Research Services, and data from January 1990 to May 1999 was provided by the statistics department at The London International Financial Futures Ex-

[7]This result is known as Granger's Representation Theorem (Engle & Granger, 1987).

[8]However, failure to reject the hypothesis $\delta_0 = 0$ and $\delta_1 = 1$ is a necessary, but not a sufficient, condition for long- *and* short-run market unbiasedness/efficiency. Cointegration-based tests are tests of long-run tendencies rather than of period-by-period equilibria (Fackler & Goodwin, 1999), and short-run inefficiencies generally are considered to present if lagged spot and futures prices (or seasonal factors) help explain movements in the current spot price. The significance of these lagged parameters implies that incorporating these in a forecasting model could help predict future spot and futures prices.

change (LIFFE), thus providing a continuous series of spot and futures prices, and one year of overlap.[9]

Due to problems of residual correlation introduced by overlapping observation intervals (Granger & Newbold, 1977), the futures price is chosen at a forecast horizon equal to the observation interval. Therefore, the futures price data for the current-month contract, which began trading in July 1988, comprises 131 nonoverlapping observations. The futures price corresponding to this contract is denoted $F_{t|t-1}$. The one-month contract has a total of 129 observations, denoted $F_{t|t-2}$, and the two-month contract has a total of 90 observations, and is denoted $F_{t|t-3}$. The quarterly contract, which began trading in May 1985, yields 55 observations. Futures prices for the quarterly contract are denoted $F_{t|t-4}$. For all contracts, the spot price S_t, is constructed as the average of the last five index values at the end of the delivery month. The only exception is the December contract, which uses the average of the last five days before the twentieth of December.[10] For the current, one- and two-month contracts, the last 20 observations were withheld from the cointegration analysis in order to conduct out-of-sample forecasting. All analysis conducted is performed using logarithmic transformations.[11]

A test for the order of integration was performed on the level of each of the spot and futures prices associated with each contract, and the first difference of each of the series. Augmented Dickey–Fuller (1979, 1981) test statistics are presented in Table I. The test statistics were undertaken allowing for the presence of a time trend. Excluding the time trend yielded qualitatively similar results, so the results are excluded to conserve space. Results indicate that all time series are nonstationary, a necessary condition for cointegration. When the tests are applied to the differenced series, however, test statistics clearly reject the null hypothesis of unit root.

MODEL ESTIMATION

An ECM was estimated for each contract using the maximum-likelihood technique, as outlined by Johansen and Juselius (1990). In each case,

[9]Due to the destruction of the Baltic Exchange in April 1992, much of the futures exchange data was lost. The author therefore thanks Clarkson Research Services for providing the data back to 1985.

[10]Replacing the last index value with the average of the last five index values in the delivery month was suggested by an anonymous reviewer. The concept behind using an average of the last five days is done, according to LIFFE, in order to prevent market manipulation. Beginning with the November 1999 contract month, the average of the last six index values in the delivery month has been adopted.

[11]All analysis conducted in this paper also was performed leaving data in the levels. As the results obtained from the levels analysis were similar qualitatively to those using logarithmic transformations, they are excluded to conserve space. These results are available from the author upon request.

TABLE I

Augmented Dickey-Fuller (ADF) Tests for Order of Integration on Futures Prices (F_{t-n}) and Average of the Last Five Index values (S_t)

Test is on the estimated coefficient Θ_1 from the following prototype model:

$$\Delta X_t = \Theta_0 + \Theta_1 X_{t-1} + \sum_{k=1}^{K} \beta_k \Delta X_{t-k}$$

Contract	N	K	HO: I(1) vs HA: I(0) ADF	HO: I(2) vs HA: I(1) ADF
Current				
Ln (F_{ft-1})	111	2	−2.911	−5.749
Ln (S_t)	111	2	−2.851	−5.670
One Month				
Ln (F_{ft-2})	109	2	−2.867	−5.707
Ln (S_t)	*109*	*2*	*−2.757*	*−5.762*
Two Month				
Ln (F_{ft-3})	70	3	−2.349	−3.685
Ln (S_t)	70	3	−1.690	−4.765
Quarterly				
Ln (F_{ft-4})	55	2	−1.727	−5.360
Ln (S_t)	55	2	−2.522	−4.046

Critical values are taken from Fuller (1976). They are −2.57 (10%), −2.88* (5%), and −3.46 (1%). Therefore, based on these results are series are *I*(1). The optimal lag length (*K*) was based on the Akaike Information Criterion (1973). *N* represents the number of in-sample observations.

the estimation was restricted to one cointegrating vector. One lag length was shown to be optimal for the current-month contract, two and quarterly contracts, whereas five lags were shown to be optimal for the one-month contract based on the Akaike Information Criterion (AIC; Akaike, 1973). Each ECM also was estimated by including eleven monthly dummy variables to account for possible seasonality in the price series. While in each model some of the dummy variables were found to be significant statistically, their inclusion did not change qualitatively the sign, magnitude, and hypothesis test results, and so, to conserve space, the long-run parameter results from the ECM estimation excluding the dummy variables are presented.[12] The first ECM describes the relationship between the futures and spot prices for the current contract. The results in the first panel of Table II clearly illustrate that the null hypothesis of no cointegrating relationship, i.e., $r = 0$, is rejected at the 1% level using both λ_{max} and λ_{trace} statistics. The hypothesis that $r = 1$ is not rejected, however. These test results clearly indicate that the futures and

[12]Such a result suggests that there may be short-run inefficiencies in the futures market that may be exploited by incorporating seasonality into time series forecasts, which is undertaken later in this paper. Denning et al. (1994) also found evidence of seasonality.

TABLE II

Johansen Cointegration Tests

λ_{trace}			λ_{max}		
Test Statistic	*Hypothesis*	*99% Critical Value*	*Test Statistic*	*Hypothesis*	*99% Critical Value*
		Current-Month Contract (k = 1)			
125.25	$r = 0$	19.19	117.72	$r = 0$	14.90
7.53	$r \leq 1$	11.65	7.53	$r = 1$	8.18
		One-Month Contract (k = 5)			
38.86	$r = 0$	19.19	28.81	$r = 0$	14.90
10.04	$r \leq 1$	11.65	10.04	$r = 1$	8.18
		Two-Month Contract (k = 1)			
39.62	$r = 0$	19.19	35.62	$r = 0$	14.90
4.00	$r \leq 1$	11.65	4.00	$r = 1$	8.18
		Quarterly Contract (k = 1)			
60.04	$r = 0$	19.19	52.23	$r = 0$	14.90
7.81	$r \leq 1$	11.65	7.81	$r = 1$	8.18

Because there are two series, the number of stochastic trends between the spot and futures prices is $2 - r$, where r denotes the number of cointegrating relations, 0 or 1. Each cointegrating vector is normalized with respect to the cash price, S_t. Tests are on eigenvalues with the Π matrix. The λ_{trace} statistic is $N(\Sigma_{i=r+1}^{2} \ln(1 - \lambda_i))$, where λ_i are ordered (largest ot smallest) eigenvalues on Π, and the λ_{max} statistic is $N(1 - \lambda_{r+1})$. Critical values for the λ_{max} and λ_{trace} statistics are from Osterwald-Lenum (1992). Residuals behave as white noise using lagrange multiplier statistics. The null hypothesis of no ARCH effects is not rejected for each model. The optimal lag length (k) is based on the Akaike Information Criterion (1973).

spot prices associated with the current contract ($F_{t|t-1}$ and S_t) are indeed cointegrated. The estimated coefficients for the current-month contract are:

$$\begin{bmatrix} \Delta F_{t|t-1} \\ \\ \\ \Delta S_t \end{bmatrix} = \begin{bmatrix} -0.993 & 0.965 & 0.194 \\ (-12.249) & (12.249) & (12.249) \\ -0.120 & 0.116 & 0.023 \\ (-0.986) & (0.986) & (0.986) \end{bmatrix} \begin{bmatrix} F_{t-1|t-2} \\ S_{t-1} \\ 1 \end{bmatrix}$$

All figures in parenthesis represent t-values associated with parameter estimates. The π or $\alpha\beta'$ matrix is factored as:

$$\alpha\beta' = \begin{bmatrix} .965 \\ \\ .116 \end{bmatrix} [-1.029 \;\; 1.000 \;\; 0.201],$$

and the corresponding long-run equilibrium between BIFFEX futures and spot prices is represented by $\beta' = (-1.029 + 1.00 + 0.201)$. Perturbations in this equilibrium relationship then are represented as $z_t =$

TABLE III

Tests of Weak Exogeneity of Spot and Futures Prices[a]

Contract	H_O:	$\chi^2(1)$	*p value*	H_O:	$\chi^2(1)$	*p value*
Current	$\alpha_F = 0$	89.23	0.000	$\alpha_S = 0$	0.93	0.33
One-month	$\alpha_F = 0$	18.29	0.000	$\alpha_S = 0$	0.92	0.34
Two-month	$\alpha_F = 0$	22.18	0.000	$\alpha_S = 0$	0.00	0.97
Quarterly	$\alpha_F = 0$	44.40	0.000	$\alpha_S = 0$	2.54	0.11

[a]Tests are based on the following: $T = N(\ln(1 - \lambda_R) - \ln(1 - \lambda_U))$, where λ_R is the eigenvalue calculated with the restriction and λ_U the eigenvalue calculated without the restriction.

$\ln(S_t) - 1.029\ln(F_{t|t-1}) + 0.201$, where z_t represents stationary deviations in the long-run equilibrium between the two sets of prices.

The loading matrix, α, is the measure of the average speed of convergence towards the long-run equilibrium and plays a crucial role in analyzing how each of the price series will respond to deviations from the long-run equilibrium relationship. The larger the value of α, the faster the speed of adjustment. The null hypothesis of weak exogeneity is a hypothesis about the rows of α when the parameters of interest are the long-run parameters α and β. The null hypothesis for weak exogeneity for a price series is: H_O: $\alpha_i = 0$, against the alternative H_A: $\alpha_i \neq 0$. The test statistic is shown by Johansen (1991) to be distributed asymptotically as a χ^2 statistic. Tests for weak exogeneity are presented in Table III. The first row corresponds to tests related to the current-month contract. The hypothesis that the BIFFEX futures price is weakly exogenous with respect to the long-run equilibrium is rejected strongly. Weak exogeneity is not rejected for the spot price, implying that the spot price is *not* affected by transitory perturbations in the long-run relationship. Consequently, when the spot and futures prices fall out of their long-run equilibrium relationship (when they fall off of the line represented by β' above), it is the futures market that adjusts to restore the equilibrium. This is an especially intuitively pleasing result, given that it is much more difficult to trade each of the routes comprising the index than it is to trade the actual BIFFEX freight futures contract.

Therefore, in the case of the current-month contract, the futures price responds through the α vector (.965), such that when the spot price is too high in realtion to the futures price in period t, the futures price will adjust (quickly) in period $t + 1$. As noted by the insignificance of the

TABLE IV

Testing Restrictions on the Cointegrating Relationships

Contract	δ_0	δ_1	*HO:* $\delta_0 = 0$	*p Value*	*HO:* $\delta_1 = 1$	*p Value*	*HO:* $\delta_0 = 0$ $\delta_1 = 1$	*p Value*
Current	0.201	1.029	0.98	0.32	1.06	0.30	4.18	0.12
One-Month	0.613	1.086	1.01	0.32	1.06	0.30	3.70	0.16
Two-Month	0.769	1.110	0.40	0.53	0.43	0.51	1.77	0.41
Quarterly	1.061	1.150	2.70	0.10	2.74	0.10	2.85	0.24

The statistics for testing the individual hypothesis have a χ^2 distribution with one degree of freedom, whereas the joint hypothesis has a χ^2 distribution with two degrees of freedom.

speed-of-adjustment parameter associated with the spot market, the spot price does not respond to shocks in the long-run relationship.

Long-run unbiasedness, and hence efficiency, also can be tested within the long-run cointegrating vector, β. As previously pointed out, the hypothesis that the futures price is an unbiased predictor of the spot price is a joint hypothesis that markets are efficient and that there is no risk premium. In the context that unbiasedness is not rejected, unbiasedness and market efficiency can be regarded as synonymous. Table IV illustrates that the parameter values are very close to their theoretically correct values, $\delta_0 = 0$ and $\delta_1 = 1$, and that each individual test statistic confirms the restrictions. Moreover, the joint test, distributed as a χ^2 statistic with two degrees of freedom, illustrate that the null hypothesis of unbiasedness cannot be rejected at the 1% level of significance. Therefore, the results indicate that for the current-month contract, the futures contract is an unbiased long-run predictor of the spot price. This evidence supports the long-run unbiasedness hypothesis, and hence market efficiency with no risk premium.

The second ECM describes the relationship between the futures and spot prices for the one-month contract. The results in the second panel of Table II clearly illustrate that the null hypothesis of no cointegrating relationship, i.e., $r = 0$, is rejected at the 1% level using both λ_{max} and λ_{trace} statistics and that the hypothesis that $r = 1$ is not rejected using the λ_{trace} statistic. These test results clearly indicate that the spot and futures prices associated with the one-month contract (S_t and $F_{t|t-2}$) are cointegrated. The estimated coefficients for the one-month contract are:

$$\begin{bmatrix} \Delta F_{t|t-2} \\ \\ \Delta S_t \end{bmatrix} = \begin{bmatrix} \begin{matrix} 0.163 & -0.192 \\ (1.463) & (-1.502) \end{matrix} \\ \\ \begin{matrix} 0.304 & -0.124 \\ (1.941) & (-0.693) \end{matrix} \end{bmatrix} \begin{bmatrix} \Delta F_{t-1|t-3} \\ \\ \Delta S_{t-1} \end{bmatrix}$$

$$+ \begin{bmatrix} \begin{matrix} 0.186 & 0.052 \\ (1.792) & (0.466) \end{matrix} \\ \\ \begin{matrix} 0.325 & -0.190 \\ (2.227) & (-1.206) \end{matrix} \end{bmatrix} \begin{bmatrix} \Delta F_{t-2|t-4} \\ \\ \Delta S_{t-2} \end{bmatrix}$$

$$+ \begin{bmatrix} \begin{matrix} -0.015 & -0.178 \\ (-0.182) & (-1.749) \end{matrix} \\ \\ \begin{matrix} 0.128 & -0.233 \\ (1.095) & (-1.630) \end{matrix} \end{bmatrix} \begin{bmatrix} \Delta F_{t-3|t-5} \\ \\ \Delta S_{t-3} \end{bmatrix}$$

$$+ \begin{bmatrix} \begin{matrix} 0.013 & -0.176 \\ (0.190) & (-1.883) \end{matrix} \\ \\ \begin{matrix} -0.279 & -0.126 \\ (-2.807) & (-0.960) \end{matrix} \end{bmatrix} \begin{bmatrix} \Delta F_{t-4|t-6} \\ \\ \Delta S_{t-4} \end{bmatrix}$$

$$+ \begin{bmatrix} \begin{matrix} -0.650 & 0.598 & 0.367 \\ (-5.672) & (5.672) & (5.672) \end{matrix} \\ \\ \begin{matrix} -0.188 & 0.173 & 0.106 \\ (-1.168) & (1.168) & (1.168) \end{matrix} \end{bmatrix} \begin{bmatrix} F_{t|t-3} \\ S_{t-1} \\ 1 \end{bmatrix}$$

and the corresponding π matrix is factored as:

$$\alpha\beta' = \begin{bmatrix} .598 \\ \\ .173 \end{bmatrix} [-1.086 \;\; 1.000 \;\; 0.613].$$

The one-month contract's long-run equilibrium between BIFFEX futures prices and the spot price is represented by $\beta' = (-1.086 + 1.00 + 0.613)$ and perturbations in this equilibrium relationship are represented as $z_t = \ln(S_t) - 1.086\ln(F_{t|t-2}) + 0.613$. Tests for weak exogeneity also are undertaken for the one-month contract, and these results are presented on the second row of Table III. The hypothesis that the BIFFEX futures price is weakly exogenous with respect to the long-run equilibrium is rejected strongly, while weak exogeneity is not rejected for the spot price. Again, this implies that the spot price is not affected by transitory departures from the long-run relationship, and that the futures market adjusts to restore the long-run equilibrium relationship. Unbiasedness and efficiency are tested within the long-run cointegrating vector, β for the one month current. Individual test statistics in Table IV confirm the restrictions necessary to prove the existence of long-run unbiasedness (efficiency). The evidence again supports the long-run unbiasedness hypothesis, and hence market efficiency with no risk premium for the one-month contract.

The magnitude and significance of several of the coefficients presented in the Γ matrices (short-run parameter matrices) for the one-month contract suggest that the spot and futures markets are quite responsive to lagged changes in both markets. This suggests that there could be short-run inefficiencies in the markets that may be exploited by utilizing the lagged relationship in a forecasting setting (see section on forecasting below).

The next ECM describes the relationship between the futures and spot prices for the two-month contract. Results in the third panel of Table II again illustrate that the null hypothesis of no cointegrating relationship, i.e., $r = 0$, is rejected at the 1% level, and that the hypothesis that $r = 1$ is not rejected. This implies, like the previous two contracts, that the spot and futures prices associated with the two-month contract (S_t and $F_{t|t-3}$) are cointegrated. The estimated coefficients for the two-month contract are:

$$\begin{bmatrix} \Delta F_{t|t-3} \\ \\ \\ \Delta S_t \end{bmatrix} = \begin{bmatrix} -0.295 \;\; 0.266 \;\; 0.204 \\ (-5.611)(5.611)(5.611) \\ \\ -0.003 \;\; (0.003) \;\; 0.002 \\ (-0.037)(0.037)(0.037) \end{bmatrix} \begin{bmatrix} F_{t|t-4} \\ \\ S_{t-1} \\ \\ 1 \end{bmatrix},$$

and the corresponding π matrix is:

$$\alpha\beta' = \begin{bmatrix} .266 \\ \\ .003 \end{bmatrix} [-1.110 \;\; 1.000 \;\; 0.769].$$

The two-month contract's long run is represented by $\beta' = (-1.110 + 1.00 + 0.769)$ and perturbations in this equilibrium relationship are represented as $z_t = \ln(S_t) - 1.110\ln(F_{t|t-3}) + 0.769$. Tests for weak exogeneity also are undertaken for the two-month contract, and these results are presented on the third row of Table III. These results suggest that the BIFFEX futures price is not weakly exogenous with respect to the long-run equilibrium, but weak exogeneity is not rejected for the spot price. Tests for unbiasedness and efficiency also are undertaken for the two-

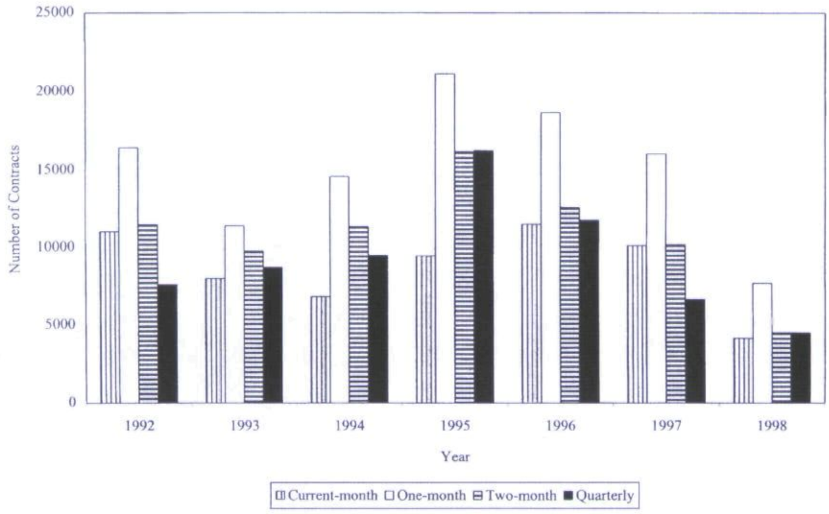

FIGURE 1
Volume of Trading by Contract Month: 1992–1998.

month current. Individual test statistics, again reported in Table IV, support the restrictions necessary to prove the existence of long-run unbiasedness (efficiency) for the two- month contract.

The last ECM to be estimated is for the quarterly contract. Even though this contract is quite 'distant', it is still a popular contract used for hedging by traders, as represented by its relatively large share of volume for the BIFFEX contract (see Fig. 1). The results in the last section of Table II illustrate that S_t and $F_{t|t-4}$ associated with the quarterly month contract also are cointegrated, and that the estimated coefficients for this contract are:

$$\begin{bmatrix} \Delta F_{t|t-4} \\ \\ \\ \Delta S_t \end{bmatrix} = \begin{bmatrix} -0.769 \quad 0.669 \quad 0.710 \\ (-9.381)(9.381)(9.381) \\ \\ -0.211 \quad 0.183 \quad 0.194 \\ (-1.702)(1.702)(1.702) \end{bmatrix} \begin{bmatrix} F_{t|t-5} \\ \\ S_{t-1} \\ \\ 1 \end{bmatrix}.$$

The corresponding π matrix for the quarterly contract is factored as:

$$\alpha\beta' = \begin{bmatrix} .669 \\ \\ .183 \end{bmatrix} [-1.150 \quad 1.000 \quad 1.061],$$

and the quarterly contract's long-run representation is given by β' =

$(-1.150 + 1.00 + 1.061)$. Temporary disturbances in this equilibrium relationship are represented as $z_t = \ln(S_t) - 1.150\ln(F_t) + 1.061$.

Tests for weak exogeneity for the quarterly contract are displayed in the last line of Table III. The hypothesis that the BIFFEX futures price is weakly exogenous with respect to the long-run equilibrium is strongly rejected once again, as with all other contracts. Weak exogeneity is not rejected for the spot price, implying that the spot price is not affected by transitory perturbations in the long-run relationship. Consequently, like all other contracts studied here, when the spot and futures prices fall out of their long-run equilibrium relationship, it is the futures market that adjusts to restore the equilibrium. Therefore, in the case of the quarterly month contract, the futures price responds through the α vector (.669), so that when the spot price is too high in relation to the futures price in period t, the futures price will adjust in period $t + 1$. Once again, as noted by the insignificance of the speed-of-adjustment parameter associated with the spot market, the spot price does not respond to shocks in the long-run relationship.

Table IV illustrates that the parameter values for the quarterly contract are very close to their theoretically correct values, $\delta_0 = 0$ and $\delta_1 = 1$, and that each individual test statistic once again confirms the restrictions. Moreover, the joint test illustrates that the null hypothesis cannot be rejected at the 1% level of significance for long-run unbiasedness (and hence efficiency). Therefore, the results indicate that, for the quarterly month contract, the futures contract is an unbiased long-run predictor of the spot price. This evidence supports the unbiasedness hypothesis, and hence market efficiency with no risk premium. These results strongly contradict the results presented by Kavussanos and Nomikos (1999), who report the quarterly BIFFEX contract to be long-term biased.[13]

CHANGES IN THE INDEX, STABILITY OF THE COINTEGRATING RELATIONSHIP, VOLUME, AND FFA'S

Because freight services differ in terms of size, gear, fuel efficiency, cargo, destination, dates, and so forth, and in a service market like the freight

[13]In their paper, the authors only use 33 observations for the quarterly contract (beginning July 1988). The relatively small number of observations seems to have driven their results of finding long-term biasedness, and hence the motivation behind much of the paper. The quarterly contract was, however, the first contract to be traded at the BIFFEX, and data is available back to the inception day—May 1985. Therefore, in this paper, *all* observations are utilized, creating a more-stable sample size of 55 observations. The more-stable results found in this paper suggest that it is unwise to jump to the conclusion that a futures market is biased when basing the results on an incomplete, and rather small, sample size.

market where there exists no standard uniform unit, the only possible option was to develop an index in order to trade freight futures contracts. In its inception month, May 1985, the index (BFI) was comprised of thirteen routes. However, because the spot freight market is changing continuously, previously important voyages became irrelevant or subject to long-term contracts. New voyages also have become more important. Therefore, as the index was designed to reflect correctly the freight market, its composition had been updated regularly. In fact, since its inception day, the BFI had undergone 13 major refinements. In November 1999, the BFI stood at 11 different routes (see Appendix). The major reason for changes in the index, prompted by LIFFE, was to attract trading activity in the contract brought about by greater hedging activity. Since the various segments of the dry-bulk cargo market may move somewhat independently, the index seemed to have failed to reflect accurately the market condition on lightly weighted routes, and so the hedging effectiveness was reduced. Therefore, the index, over time, has become defined more narrowly and the weightings of each of the routes has changed, with the intention that it become a more-accurate hedging instrument for the more important routes.

The numerous changes in the index over time suggest that the relationship between the index and the futures price has changed over time. Therefore, to test this hypothesis of a stable cointegrating vector, an examination of the relationship between the index and futures prices over time is undertaken for the current-month contract. In order to estimate the stability of cointegration over time, eq. (3) was estimated recursively using a moving 50-month estimation period, beginning with the first month of trading (July 1988) for the current-month contract. Figure 2 illustrates the time series of trace statistics and the critical values for testing for the existence of a cointegrating relationship. As can be seen from the figure, at every date, the null hypothesis of no cointegration is rejected strongly. Moreover, the time-series plot of the trace statistic illustrates that the long-run relationship between the futures and spot prices (index) indeed has become *stronger* over time (the trace statistics increases)—likely the result of a narrowing in the index value.[14]

Despite the apparent strengthening of the relationship between the spot and futures prices over time in the BIFFEX freight futures market, the level of trading activity has remained very low in this unique market. Figure 3 illustrates the yearly volume of futures (not options) that have

[14]Such stability tests also were undertaken for the one- and two-month contracts. Results also suggest evidence of a stable cointegrating relationship, but as the current month yielded the most amount of useable observations, the results presented here focus on this contract.

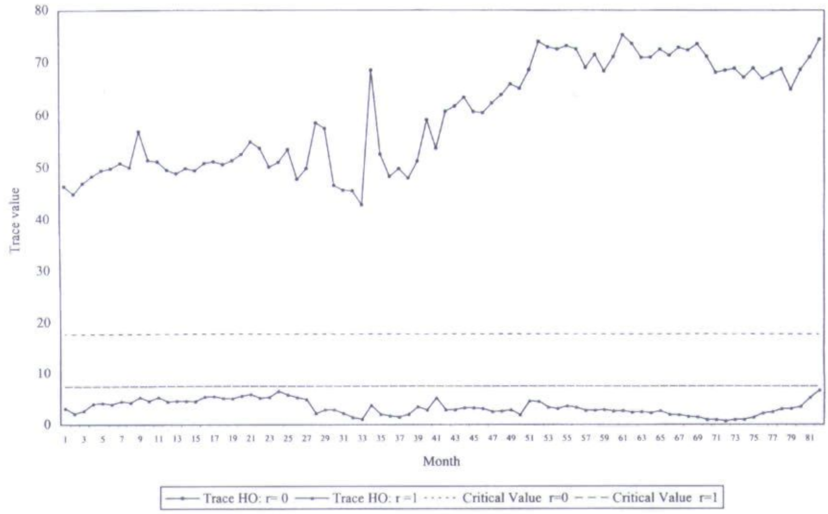

FIGURE 2
Trace Tests for $r = 0$ and $r = 1$ for Current-Month Futures Contract: July 1988–April 1995.

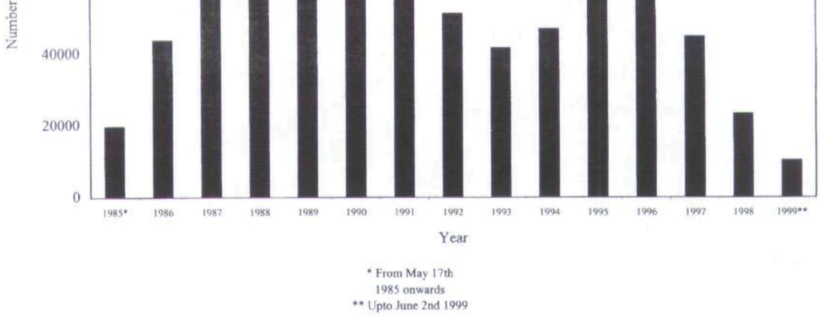

FIGURE 3
Volume of Trading by All Contracts: May 1985–June 1999.

been traded since the inception date of May 1985, and illustrates the rather low levels of trading activity. The consequent decrease in volume in futures contracts has coincided with an explosive growth in freight forward agreements, beginning in 1992.[15] The FFAs are tailor-made hedges between two principals who have different views of the market with a need to hedge a forward commitment. These agreements offer the opportunity of trading individual routes from the index for virtually any time period, offering a more accurate hedging mechanism. Interestingly, market participants attribute the decrease in volume in the futures market to be associated partially with the growth in hedging activity with the forward contracts, not with inefficiencies associated with the BIFFEX freight futures contract.[16] This contradicts previous explanations by Kavussanos and Nomikos (1999), who attributed the lack of volume in the BIFFEX with their apparent finding of long-run unbiasedness in the contract. In fact, the results shown in this paper illustrate that developments in the freight index seem to have made the cointegrating relationship much stronger, and hence more efficient, thus making the freight futures contract a more efficient hedging mechanism for traders on certain routes.

In November 1999, the BFI was replaced with the Baltic Panamax Index (BPI), representing just 7 Panamax routes (see Appendix), with the intention that it will prove to be a more narrowly defined index, which, it is hoped, will attract greater hedging interest. Results shown in this paper suggest that such a narrowing of the index (by adopting the BPI) should strengthen the relationship between the spot and futures markets, thus making it a more viable hedging mechanism for traders on routes that comprise the BPI. Moreover, because of the switch to the BPI, further trading activity is expected to be induced, brought about by increased arbitrage trading activity between the BIFFEX futures markets and the growing forward markets. To date, the major focus of trading FFAs has been confined to four major routes: Routes 1A, 2, 2A, and 7. Each of these routes currently is correlated highly with the freight futures contract, with correlation coefficients of 0.953, 0.952, 0.967, and 0.879, respectively. The narrowing of the index is expected to enable traders to create a forward value of the BPI using forward contracts with more ease and thus enable traders to isolate more easily discrepancies between the

[15]The actual volume of forward contracts traded by ship-brokers is confidential, and is therefore not presented in this paper. However, interested readers may contact the author for further specific information regarding the growth of the forward market.

[16]I am grateful to an anonymous reviewer for questioning the link between market unbiasedness and lack of volume. I also am grateful to Ian Dudden of LIFFE and Andy James of Clarkson Research Services for useful discussions on this issue.

forward and futures markets (which is nothing more than a futures-based value of the BPI), thus invoking more arbitrage activity between the forward and futures markets. Moreover, forward traders focusing on a major route will have the ability to offset *their* risk in the futures market if necessary, as the correlation between a particular route and the futures price is likely to further strengthen with the narrowing of the index.[17]

INCORPORATING LONG-RUN COINTEGRATING INFORMATION IN FORECASTING

To date, there have been several studies that have attempted to forecast freight prices. For instance, Cullinane (1992) used a Box–Jenkins model of the BFI to forecast freight prices. Results indicated that such models fail at long-range forecasting. Chang and Chang (1996) used time-series models to predict the spot market. They concluded that the BIFFEX futures market can predict future fright prices successfully at one-month, but forecasting performance significantly decreases beyond this horizon. Denning et al. (1994) also used time-series analysis to predict future index values, and presented a predictive nonlinear equation that appears to be able to predict future index values. Veenstra and Franses (1997) utilized a VAR to help forecast monthly future freight rates, incorporating the cointegrating relationship found between six freight prices for different routes. They discovered that such a model did not improve the accuracy of short- or long-term forecasts of future freight values and attributed the failure of their forecasting models to the general stochastic nature of freight prices. Finally, Kavussanos and Nomikos (1999) used daily data to forecast future freight spot prices, using time-series models that incorporated cointegrating information. Their study found that futures prices outperform all time-series models with the exception of a one-month forecast, which is outperformed by an ECM. This study follows the approach of other non-BIFFEX-related studies by Wahab and Lashgari (1993), Bessler and Covey (1991), and Ghosh (1993), who also used daily data and incorporated cointegrating relationships to help forecast future spot prices.

While most of these studies found that error-correction models outperform other time-series models, they all used daily futures prices that had a decreasing time to maturity. Since futures prices and cash prices

[17]I am grateful to the journal editor for suggesting this, and Phillipe Van den Abeele of Clarkson Research Services, Ian Dudden of LIFFE, and Neill Kennedy of the Baltic Exchange for confirming this possibility.

should converge at expiration of the contract, the differential between the two price series also must be decreasing to zero. As such, Brenner and Kroner (1995) suggested that the differential is expected to be nonstationary, and so an appropriate cointegrating relationship captured in the ECM would involve not only the spot and futures prices, but also the differential between the two. It is for this reason that they suggested modeling the spot and futures price series that have a *fixed* time to maturity (as was done throughout this current paper). Therefore to maintain consistency throughout this paper, the forecasts here are based on a monthly horizon.

For each contract, a model was estimated over the appropriate in-sample time frame, and each model then was used to forecast the next out-of-sample data point, starting with October 1997. Each model then was moved ahead one data point, reestimated and then reforecasted. This procedure was repeated 20 times, making available a total of twenty out-of-sample forecasts for the current, one-month, and two-month contracts, ending with a forecast of the May 1999 spot price.

In order to compare the performance of the ECM, a comparison is made to the forecasting performance of a VAR (diff), which is equivalent to using an ECM but imposing $r = 0$ (the number of cointegrating vectors set equal to zero). Also, as a further means of comparison, a VAR (levels) is estimated. The forecastability of the current futures price also was included in the analysis. For the futures market forecast, using the most recent futures price implies a random-walk model, whereas including the futures price as a means of predicting the future spot rate enables one to isolate the effectiveness of simply using the futures price as a forecaster compared to other models. All time-series models were estimated with the inclusion of 11 seasonal dummy variables to account for seasonality, and in the case of the one-month contract, lagged spot and futures prices were included in the model to account for possible short-run inefficiencies.

Forecast performance is evaluated using conventional methods, including Mean Squared Error (ME), Mean Absolute Error (MAE) (measuring the absolute deviation of the forecasted price from the actual value), and Root Mean Squared Error (RMSE) (which takes the average deviation of the forecasted value to the actual while penalizing larger errors more than smaller ones).

Table V reports the forecast statistics for forecasting futures prices for the three contracts (upper panel) and the forecast statistics for forecasting future spot prices (the average of the last 5 days of the index) in the lower panel. The upper panel (forecasting futures prices) illustrates

TABLE V

Forcasting Evaluation Statistics

Model	*Mean Error (ME)*	*Mean Absolute Error (MAE)*	*Root-Mean-Square Error (RMSE)*	*N*
		Forecasting Futures Prices Current-Month Contract		
Futures	−0.017	0.088	0.112	20
VAR (levels)	−0.014	0.045	0.050	20
VAR (diff)	−0.021	0.084	0.103	20
ECM	0.009	0.044	0.049	20
		One-Month Contract		
Futures	−0.018	0.092	0.106	20
VAR (levels)	0.005	0.044	0.058	20
VAR (diff)	−0.006	0.047	0.062	20
ECM	−0.000	0.045	0.058	20
		Two-Month Contract		
Futures	−0.015	0.073	0.084	20
VAR (levels)	0.003	0.050	0.060	20
VAR (diff)	−0.021	0.063	0.073	20
ECM	0.000	0.050	0.060	20
		Forecasting Spot Prices (Average of Last 5 Days of Index) Current-Month Contract		
Futures	−0.010	0.063	0.077	20
VAR (levels)	−0.040	0.078	0.092	20
VAR (diff)	−0.020	0.066	0.084	20
ECM	−0.017	0.070	0.086	20
		One-Month Contract		
Futures	−0.050	0.098	0.115	20
VAR (levels)	−0.050	0.074	0.099	20
VAR (diff)	−0.018	0.070	0.092	20
ECM	−0.017	0.070	0.092	20
		Two-Month Contract		
Futures	−0.085	0.131	0.151	20
VAR (levels)	−0.033	0.074	0.092	20
VAR (diff)	−0.020	0.066	0.086	20
ECM	−0.014	0.068	0.087	20

Futures denotes the futures price associated with the particular contract month. VAR (levels) denotes a Vector Autoregressive Model estimated in the levels, VAR (diff) denotes a Vector Autoregressive Model estimated in first differences, and ECM denotes Error Correction Model (incorporating the long run cointegrating information).

that for each contract, the ECM and the VAR (levels) outperform the random-walk model and the VAR (diff) model. In fact, performance of the ECM and VAR (levels) model are almost identical. This is noted, for example, by seeing that in all cases, the ME, MAE, and RMSE are lower for the ECM for the current contract, and identical to the VAR (levels) for the one- and two-month contract horizons. The result that the VAR (levels) and ECM perform similarly is consistent with the theory of cointegration, and is a result consistent with that found by Bessler and Fuller (1993) and Lin and Tsay (1996). That is, accounting for the cointegrating relationship in the forecasting equation does not necessarily outperform a VAR (levels). These results, and the results found in this paper, are consistent with Engle and Granger (1987), who suggested that a VAR (levels) contains all cointegration information, at least asymptotically. Futures prices do respond to temporary deviations in the long-run equilibrium relationship (to perturbations in β, because $\alpha_F \neq 0$) and so, forecasts of futures prices accounting for $\alpha_F \neq 0$ will be better than any forecast from a VAR (diff) (which sets $\alpha_F = 0$).

For the current contract, the futures price appears to be the best forecaster of the current-month spot price compared to all models. However, the ability of the futures price to predict accurately future spot rates beyond a month horizon decreases. This result is entirely consistent with results of Chang and Chang (1996), who also found that the forecasting ability of the BIFFEX freight futures contract dramatically declines beyond a one-month horizon. Similar results were found by Ma (1989), who found that the forecasting ability of the futures price tended to decrease the longer the time horizon. Beyond the current-month horizon, all time-series models appear to outperform the futures price as more accurate predictors of future spot rates. In particular, the one- and two-month contracts are shown, in general, to be explained best (or at least equally well) by a VAR in first differences rather than using either the futures price, a VAR (levels), or an ECM model. Such a result confirms the previous cointegration results that suggest that the spot freight rate does not appear to adjust to perturbations in the cointegrating relationships. Therefore, a model incorporating such information [like an ECM or a VAR (levels)] likely is to be outperformed by a VAR in first differences in explaining future spot prices, based on the RMSE criterion.

While the superior forecasting performance of time-series models may sound surprising on the surface, it is not entirely inconsistent with results previously found in the literature. For instance, Wahab and Lashgari (1993) used cointegration analysis to examine the long-run cointegrating relationship between stock index and stock index futures prices for the S&P-500 and the FT-100 share index. They illustrated that an error-correction model of the prices results in substantially lower mean absolute forecast errors in many cases compared to other models. Ghosh (1993) undertook a similar approach by looking at S&P-500 data and found that forecasts from an error-correction model were more accurate based on RMSE comparisons. Bessler and Covey (1991) also found a similar result when they incorporated cointegrating relationships between slaughter-cattle cash and futures prices.

CONCLUSION

This article applies the theory of cointegration to the BIFFEX freight futures market using the most complete data set available. Results indicate that the current, one, two, and quarterly contract futures prices are unbiased and hence efficient estimators of future spot freight rates. The implication is that traders can be assured that, in general, the current futures price is likely to predict accurately future spot rates, without the trader having to pay a risk premium for the privilege of trading the contract. Results suggest that futures prices, rather than spot prices, respond to perturbations in the long-run equilibrium. Such a result is intuitively pleasing given that it is much more difficult to trade each of the routes comprising the index than it is to trade the actual BIFFEX freight futures contract.

Several changes in the index over its history suggest that the relationship between the index and the futures price has evolved over time. Therefore, to test the stability of the cointegrating vector between the current-month futures contract price and the spot price (the average of the index for the last five days of trading), an Error-Correction Model was estimated recursively using a 50-month estimation horizon. The results suggest that the long-run relationship between the futures and spot prices indeed has *strengthened* over time—a likely result of the narrowing of the index, thus making it easier to predict. The decrease in the volume of contracts traded is not to be attributed to inefficiencies within the contract—but possibly to the growth of the freight forward market.

Forecasting the futures price with an Error-Correction Model is shown to outperform other models for all contract horizons, while forecasting the future spot freight rate using the actual futures price is shown to outperform other time-series models for the current-month contract. The futures contract is found to be an inferior predictor of more distant spot prices compared to time-series models. More distant contracts are shown to be explained best by a VAR in first differences rather than using either the futures price or models incorporating long-run cointegration information. Such a result confirms the cointegration results found in this article. That is, the spot rate does not appear to adjust to perturbations in the cointegrating relationships, and as such, a model incorporating such information [like an ECM or a VAR (levels)] likely is to be outperformed by a VAR in first differences.

APPENDIX

Baltic Freight Index (BFI) Component Routes: May 1998

Route	*Commodity*	*Weighting (%)*
1. U.S. Gulf–Rotterdam		10%
1A. Transatlantic (Time charter)	Grain, ore or coal	10%
2. U.S. Gulf–South Japan	Grain	10%
2A. Skaw Passero–Japan/Taiwan (Time charter)	Grain, ore or coal	10%
3. U.S. North Pacific–South Japan	Grain	10%
3A. Transpacific (Time charter)	Grain, ore or coal	10%
6. Hampton Roads–Rotterdam	Coal	7.5%
7. U.S. West Coast–Japan (Time charter)	Grain, coal, or petroleum coke	10%
8. Tubarao–Rotterdam	Iron ore	7.5%
9. Tubarao–Beilun	Iron ore	7.5%
10. Richards Bay–Rotterdam	Coal	7.5%
		100%

Baltic Panamax Index (BPI) Component Routes: Dec 1998
(The BPI succeeded the BFI with effect from the November 1999 delivery month)

Route	*Commodity*	*Weighting (%)*
1. U.S. Gulf–Rotterdam	Grain	10%
1A. Transatlantic (Time charter)	Grain, ore or coal	20%
2. U.S. Gulf–South Japan	Grain	12.5%
2A. Skaw Passero–Japan/Taiwan (Time charter)	Grain, ore or coal	12.5%
3. U.S. North Pacific–South Japan	Grain	10%
3A. Transpacific (Time charter)	Grain, ore or coal	20%
7. U.S. West Coast- Japan (Time charter)	Grain, coal, or petroleum coke	15%
		100%

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