Futures Hedging when the Structure of the Underlying Asset Changes: The Case of the BIFFEX Contract

MANOLIS G. KAVUSSANOS*
NIKOS K. NOMIKOS

This article is concerned with the hedging effectiveness of futures contracts whose underlying asset is an index, when the structure of this index is changing. The case of the freight futures (BIFFEX) contract is examined here. Investigation of this issue is particularly interesting as the composition of its underlying asset, the Baltic Freight Index (BFI), has been revised on a number of occasions in order to improve the hedging performance of the market; previous empirical evidence on the market indicates substantially lower variance reduction (4–19%), compared to other markets (up to 98%). The BFI is a weighted average dry-cargo freight rate index, compiled from actual freight rates on 11 shipping routes that are dissimilar in terms of

*Correspondence author, Manolis G. Kavussanos, Department of Shipping Trade and Finance, City University Business School, Frobisher Crescent, Barbican Centre, London EC2Y 8HB, UK. E-mail: m.kavussanos@city.ac.uk. We thank the Editor of the *Journal of Futures Markets*, Professor Robert Webb, and two anonymous referees for their constructive comments and suggestions. All remaining errors are the responsibility of the authors.

Received August, 1999; Accepted February, 2000

- *Manolis G. Kavussanos is a Senior Lecturer in Applied Economics and Director, MSc Trade, Transport and Finance in the Department of Shipping, Trade and Finance at City University Business School in London, United Kingdom.*
- *Nikos K. Nomikos is a researcher in the Department of Shipping Trade and Finance at City University Business School in London, United Kingdom.*

The Journal of Futures Markets, Vol. 20, No. 8, 775–801 (2000)

vessel sizes and transported commodities. The hedging effectiveness of the market is investigated using both constant and time-varying hedge ratios, estimated through bivariate error correction GARCH models. Our results indicate that the effectiveness of the BIFFEX contract as a centre for risk management has strengthened over the recent years as a result of the more homogeneous composition of the index. This by itself indicates that the latest restructuring of the index, in November 1999, which is aimed at increasing its homogeneity even further, is likely to have a beneficial impact on the market. Jrl Fut Mark 20:775–801, 2000.

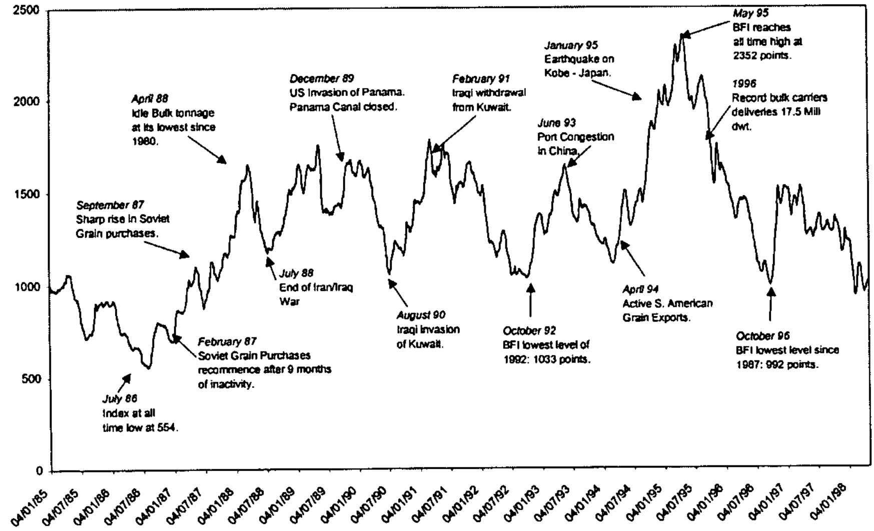
INTRODUCTION

One of the major benefits futures markets provide to economic agents is the ability to control the risk of adverse price changes in the spot market through hedging.[1] Hedging involves taking a position in the futures market, which is opposite to the position that one already has in the spot market. To achieve this, the hedger determines a hedge ratio, i.e., the number of futures contracts to buy or sell for each unit of the spot asset on which he bears price risk. The issue of estimating hedge ratios and assessing the effectiveness of futures hedges has been the focus of considerable empirical research in the literature. Initially, hedge ratios were calculated as the slope coefficient from a regression of futures price changes on spot price changes (Ederington, 1979). However, time-varying distributions of asset returns, imply that optimal, risk-minimising hedge ratios should be time-varying (Kroner & Sultan, 1993). Such time-varying hedge ratios are estimated for currencies by Kroner and Sultan (1993), for interest rate futures by Gagnon and Lypny (1995), for commodity futures by Bera et al. (1997), and for stock indices by Park and Switzer (1995) and Gagnon and Lypny (1997).

The aim of the creation, in 1985, of the Baltic International Freight Futures Exchange (BIFFEX) was to provide a mechanism for hedging freight rate risk in the dry-bulk sector of the shipping industry. On the basis of their exposure to the risk of adverse movements in freight rates, shipowners or charterers can sell or buy futures contracts representing the expected future value of the Baltic Freight Index (BFI). The BFI, is a weighted average index of 11 component routes, each reflecting regularly traded movements in international dry-cargo shipping.

Besides its use as the underlying asset of the BIFFEX contract, the BFI is also recognised by market practitioners in the shipping industry as

[1]The other benefit futures markets provide to economic agents is that of price discovery; see, e.g., Garbade and Silber (1983). For an investigation of the price discovery function of the BIFFEX market, see Kavussanos and Nomikos (1999).



Source: SSY Futures Brokers

FIGURE 1
Developments in the dry bulk market and the Baltic Freight Index (BFI), January 4, 1985–April 30, 1998.

the most reliable indicator of the condition in the dry-bulk shipping markets. Figure 1 presents the history of the BFI over the period January 1985–April 1998. Starting from an initial value of 1,000 on January 4, 1985, the index drifted down over a 2-year period as a result of the prolonged recession in the dry-bulk market. The BFI reached its all-time low of 554 on July 31, 1986. Gray (1990) indicates that this represents the lowest level to which the freight market can possibly go; in other words, the level below which the freight rate revenue earned from a particular voyage is not sufficient to cover the variable costs from the operation of the vessel. From its low point in 1986, the BFI then increased steadily, as the shipping markets were recovering from the recession, to a peak of 1,650 in April 1988. Thereafter, the index fluctuated between 1,780 points and 1,033 points for a period of almost 6 years, until April 1994. From that point, freight rates strengthened significantly, primarily due to increased grain shipments from South America, and the BFI reached its all-time high of 2,352 in May 1995. This high level of freight rates, however, could not be sustained for long as increased deliveries of newly built vessels—17.5 m dwt, the greatest ever level of deliveries, suppressed the level of freight rates. As a result, the BFI dropped to 992 points in October 1996, its lowest point since 1987. The freight market has not recovered since then and the BFI has fluctuated around the 1,000-point level.

Unlike other markets, on which futures contracts are used as a direct hedge against price fluctuations in the underlying asset, in the BIFFEX market futures contracts are employed as a cross-hedge against freight rate fluctuations on the individual shipping routes constituting the BFI. As such, it carries the risk that fluctuations on these routes will not be accurately tracked by the futures prices, thus reducing the effectiveness of the contract as a hedging instrument. For instance, futures contracts in different commodity and financial futures markets are effective in reducing the variability of the spot position by as much as 98%; see Ederington (1979) and Franckle (1980) for interest rate futures, Figlewski (1984) and Lindahl (1992) for stock indices and Malliaris and Urrutia (1991) for currencies. In contrast, Kavussanos and Nomikos (2000a) find that the hedging effectiveness of the BIFFEX contract varies from 19.2% to 4.0% across the different shipping routes that constitute the BFI.

The poor hedging performance of the contract is thought to be the primary reason for the low trading activity evidenced in the market; the daily trading volume averages only 150 contracts per day and open interest is negligible for contracts beyond 6 months ahead. The London International Financial Futures Exchange (LIFFE), the authority responsible for regulating the BIFFEX contract, has attempted to address this issue by altering the structure of the BFI on a number of occasions; first by including time charter routes, in addition to spot freight routes,[2] then by dropping from the index the handysize routes and, eventually, in November 1999, by removing the capesize routes too, leaving only the panamax routes in the renamed, Baltic Panamax Index (BPI).[3]

The success of these structural changes in improving the hedging effectiveness of the BIFFEX market is investigated in this article. Constant and time-varying hedge ratios are estimated for different periods, corresponding to revisions in the composition of the BFI, and their per-

[2]Spot charters and time charters are the two major vessel employment contracts in the shipping industry. In a spot charter, a shipowner undertakes the responsibility to transport a cargo from the loading port to the destination port. The freight paid by the charterers (cargo owners) to the shipowner is expressed as USD ($) per ton of cargo. In a time charter, the shipowner agrees to hire out his vessel to a chatterer for a specified time period. The freight rate paid by the charterer in this case is calculated as $ per day of hire.

[3]There are three major classes of vessels used for the transportation of different dry-bulk commodities across different parts of the world. Capesize vessels (around 120,000 dead-weight tons [dwt]) transport iron ore mainly from South America and Australia, and coal from North America, Australia, and South Africa. Panamax vessels (around 65,000 dwt) are used primarily to carry grain from North America, Argentina, and Australia, and coal from North America, Australia, and South Africa. Finally, handysize vessels (around 30,000 dwt) transport grain, mainly from North America, Argentina and Australia, and minor bulk products—such as sugar, fertilisers, steel, and scrap, forest products, nonferrous metals and salt—virtually from all over the world. See Table I and the section, Description of the Market and Properties of the Data Series, for further discussion on the composition of the BFI.

formance is compared over subperiods and across routes. Moreover, bootstrapping techniques are employed to investigate whether the differences in the degree of hedging effectiveness across the different periods are statistically significant.

This article contributes to the literature on futures markets in a number of ways. Although there have been numerous studies on the effectiveness of constant and time-varying hedge ratios in different commodity and financial futures markets, there has been no study so far in the literature to identify how the effectiveness of such strategies varies after structural changes in the composition of the underlying index. The revisions in the composition of the BFI provide a unique opportunity to investigate this issue. Comparisons of the variance reduction of the BIFFEX market, with other markets in the literature, are also being made. Moreover, the fact that the BIFFEX contract is the only futures contract whose underlying asset is a service, makes this investigation distinct.

Examination of this issue is also of considerable interest to those involved in regulating the BIFFEX market. Given that all the revisions in the structure of the BFI are driven by the intention to generate an underlying index that promotes the more effective functioning of the BIFFEX contract, it is interesting to investigate their impact on the risk management function of the futures contract; this way, we can examine whether they have achieved their intended objective. Moreover, this investigation is particularly timely, given the exclusion from the BFI of the capesize routes, in November 1999. By investigating the effect of previous revisions in the composition of the BFI on the performance of the BIFFEX contract, we can thus provide some preliminary evidence regarding the possible impact of this new restructuring of the index.

This article presents the minimum-variance hedge ratio methodology under the assumption of constant and time-varying risk. Next, some background is provided on the revisions that have been implemented to the BFI over the period covered by this study. The properties of the data are discussed and empirical results offered on the constant and time-varying hedging strategies. The statistical significance of the differences in the degree of hedging effectiveness observed across the different time periods is then evaluated.

MINIMUM RISK HEDGE RATIO METHODOLOGY

The primary purpose of hedging is to reduce the risk of adverse price changes in the physical market. To achieve this purpose, the hedger has to determine the optimal hedge ratio, i.e., the ratio of his holdings in the futures market such that the net gain or loss from his hedged position will be minimised. Johnson (1960), Stein (1961), and Ederington (1979) apply the principles of portfolio theory to show that the minimum-risk hedge ratio of a futures contract is given by the ratio of the unconditional covariance between spot and futures price changes over the unconditional variance of futures price changes.

To illustrate this, consider the case of a shipowner who wants to secure his freight rate income in the freight futures market. The return on the shipowner's portfolio of spot and futures positions can be denoted by:

$$R_{H,t} = R_{S,t} - \gamma R_{F,t}$$

where $R_{H,t}$ is the return on holding the portfolio between $t - 1$ and t; $R_{S,t} = S_t - S_{t-1} = \Delta S_t$ is the return on holding the spot position between $t - 1$ and t (i.e., the freight rate revenue); $R_{F,t} = F_t - F_{t-1} = \Delta F_t$ is the return on holding the futures position between $t-1$ and t; and γ is the hedge ratio, i.e., the number of futures contracts that the hedger must sell for each unit of spot commodity on which he bears price risk. The variance of the returns of the hedged portfolio, is given by

$$\mathrm{Var}(R_{H,t}) = \mathrm{Var}(\Delta S_t) - 2\gamma \mathrm{Cov}(\Delta S_t, \Delta F_t) + \gamma^2 \mathrm{Var}(\Delta F_t) \quad (1)$$

The hedger must choose the value of γ that minimises the variance of the returns of his hedged portfolio. Taking the partial derivative of equation with respect to γ, setting it equal to zero, and solving for γ, we obtain the optimal *conventional* hedge ratio γ^*

$$\gamma^* = \frac{\mathrm{Cov}(\Delta S_t, \Delta F_t)}{\mathrm{Var}(\Delta F_t)} \quad (2)$$

The conventional hedge is equivalent to the slope coefficient, γ^*, in the following regression:

$$\Delta S_t = \gamma_0 + \gamma^* \Delta F_t + u_t; \; u_t \sim iid(0, \sigma^2) \quad (3)$$

Within this specification, the higher the R^2 of eq. (3) the greater the effectiveness of the minimum-variance hedge. However, this method of calculating hedge ratios is criticised by Myers and Thompson (1989) and Kroner and Sultan (1993), as hedge ratios are derived based on the implicit assumption that the risk in spot and futures markets is constant over time. This assumption contrasts with the empirical evidence presented in Kavussanos and Nomikos (2000b), that the BFI routes are characterised by time-varying distributions and hence, optimal risk-minimising hedge ratios for these routes should be time-varying. The latter are defined as the ratio of the conditional covariance of spot and futures price changes over the conditional variance of futures price changes, as follows:

$$\gamma_t^*|\Omega_{t-1} = \frac{\mathrm{Cov}(\Delta S_t, \Delta F_t|\Omega_{t-1})}{\mathrm{Var}(\Delta F_t|\Omega_{t-1})} \tag{4}$$

where Ω_{t-1} refers to the information set available to market agents at period $t - 1$. This is similar to the conventional hedge ratio of eq. (2), except that time-varying conditional moments replace their time-invariant unconditional counterparts, thus allowing the hedge ratio to change through time as new information arrives in the market. In order to allow for time variation in the hedge ratios, the conditional second moments of spot and futures prices are modelled using the family of autoregressive conditional heteroskedasticity (ARCH) models introduced by Engle (1982). This study employs the following vector error-correction model (VECM) with a GARCH error structure

$$\Delta X_t = \mu + \sum_{i=1}^{p-1} \Gamma_i \Delta X_{t-i} + \Pi X_{t-1} + \varepsilon_t;\ \varepsilon_t = \begin{pmatrix} \varepsilon_{S,t} \\ \varepsilon_{F,t} \end{pmatrix} \Bigg| \Omega_{t-1} \sim IN(0, H_t)$$

$$H_t = \begin{pmatrix} h_{SS,t} & h_{SF,t} \\ h_{SF,t} & h_{FF,t} \end{pmatrix} = C'C + A'\varepsilon_{t-1}\varepsilon'_{t-1}A + B'H_{t-1}B \tag{5}$$

where $X_t = (S_t\ F_t)'$ is the vector of spot and futures prices, Γ_i and Π are 2×2 coefficient matrices measuring the short- and long-run adjustment of the system to changes in X_t and ε_t is the vector of residuals $(\varepsilon_{S,t}\ \varepsilon_{F,t})'$, which follow a bivariate normal distribution with mean zero and time-varying covariance matrix, H_t. In the specification of the conditional covariance matrix, C is a 2×2 lower triangular matrix and A and B are 2×2 diagonal coefficient matrices, with $\alpha_{ii}^2 + \beta_{ii}^2 < 1$, $i = 1, 2$ for stationarity.[4]

The following steps may be identified in estimating this model. First, the existence of a long-run relationship between spot and futures prices is investigated in the VECM of eq. (5) through the λ_{max} and λ_{trace} statistics

[4]The A and B matrices are restricted to be diagonal because this results in a more parsimonious representation of the conditional variance (see Bollerslev et al., 1994). Moreover, a GARCH(1,1) model is used because of the substantial empirical evidence that this model adequately characterises the dynamics in the second moments of spot and futures prices; see Kroner and Sultan (1993), Gagnon and Lypny (1995, 1997), Tong (1996), and Bera et al. (1997) for evidence on this.

(Johansen, 1988), which test for the rank of Π.[5] If rank(Π) = 0, then Π is a 2×2 zero matrix and the VECM is reduced to a VAR model in first differences. If rank(Π) = 2, then all the variables in X_{t-1} are $I(0)$, and the appropriate strategy is to estimate a VAR model in levels. If rank(Π) = 1, then there is a single co-integrating vector and Π can be factored as $\Pi = \alpha\beta'$, where α and β' are 2×1 vectors. Using this factorisation, β' represents the vector of co-integrating parameters and α is the vector of error correction coefficients measuring the speed of convergence to the long-run steady state. The significance of incorporating the co-integrating relationship into the statistical modelling of spot and futures prices is emphasised in studies such as Kroner and Sultan (1993), Ghosh (1993), Chou et al. (1996), and Lien (1996); hedge ratios and measures of hedging performance may change sharply when this relationship is unduly ignored from the model specification.

The estimates of the co-integrating vector, representing the long-run relationship between spot routes and futures prices, are then used in the joint estimation of the conditional mean and the conditional variance of the model in eq. (5). This takes place by maximising the bivariate conditional normal log-likelihood function. The standard errors of the estimated coefficients are computed using the Quasi Maximum Likelihood Estimation (QMLE) procedure of Bollerslev and Wooldridge (1992). This yields standard errors that are robust to departures from the maintained assumption of conditional normality.

DESCRIPTION OF THE MARKET AND PROPERTIES OF THE DATA SERIES

The BIFFEX contract is traded at the London International Financial Futures Exchange (LIFFE). Trading takes place every business day from 10:15 to 12:30 and from 14:30 to 16:40 London time. Eight contract months are always traded in the market; the current month, the following 2 months, and January, April, July, and October, up to 18 months ahead. All the contracts that remain open on the last trading day of a particular contract month, which is the last business day of the trading month, or December 20 for the December contract, are settled in cash. The settlement price is computed as the average of the BFI over the last five trading days of the contract month; the monetary value of the settlement price being $10 per index point.

[5]The Johansen (1988) procedure provides more efficient estimates of the co-integrating vector compared with the Engle and Granger (1987) two step approach (see Gonzalo, 1994). Moreover, Johansen's tests are shown to be fairly robust to the presence of non-normal innovations (Cheung and Lai, 1993) and heteroskedastic disturbances (Lee and Tse, 1995).

As mentioned in the Introduction, the BFI is a weighted average index of dry-cargo freight rates on 11 seaborne trade routes. The underlying trade routes and their respective weightings in the composition of the BFI have been revised substantially on a number of occasions to ensure that the index provides a basis for the effective functioning of the BIFFEX market. Since its inception in January 1985, these revisions in the composition of the BFI, are presented in Table I; the notes in Table I describe some minor amendments to the composition of the index.

We can broadly identify four different periods corresponding to differing compositions of the underlying index. During the first period (January 4, 1985 to August 3, 1990), the BFI consisted of capesize, panamax, and handysize spot freight rates. For the period up to November 3, 1988, the BFI consisted of 13 routes, of which 3 were capesize routes (routes 6, 8, and 10 representing 15% of the index composition), 5 were panamax routes (routes 1, 2, 3, 7, and 9, which made up 65% of the index); the remaining 5 were handysize routes (routes 4, 5, 11, 12, and 13, which accounted for the remaining 20%). After November 4,1988, route 13 was deleted, and the number of the BFI constituent routes was reduced to 12. The composition of the BFI was altered again on August 6, 1990 with the introduction of three time charter routes (routes 1A, 3A, and 5). The index was revised once more on February 5, 1991, when the panamax route 7 was replaced with a capesize route and a new time charter route (route 2A) was introduced. An additional time charter route was introduced on February 5, 1993 (route 9). The four handysize routes (i.e., routes 4, 5, 11, and 12) were eventually excluded from the composition of the index in November 1993 and the number of the BFI routes was reduced to 11; since then, the BFI represents only panamax and capesize spot and time charter rates. Finally, in December 1998, LIFFE announced that a new shipping index, the Baltic Panamax Index (BPI), will replace the BFI as the underlying asset of the futures contract from November 1999 onward. The BPI, as its name implies, will consist of the seven panamax spot and time charter routes (routes 1, 1A, 2, 2A, 3, 3A, and 9) that currently compose the BFI. The proposed weights of these routes in the new index are presented in Table I.

These revisions are driven by the intention to generate an underlying index that promotes the effective functioning of the BIFFEX contract. For instance, Gray (1990) indicates that time charter routes were introduced in order to facilitate market participants who wanted to hedge their freight rate risk on these routes. Similarly, Cullinane et al. (1999) indicate that the exclusion of the handysize routes was implemented in response to pressure from market agents, operating on panamax and capesize trade

TABLE I

Baltic Freight Index: Changes in its Composition Since Its Inception*

	Vessel Size (dwt)	Cargo	Route	1/1/85–11/3/88	11/4/88–8/3/90	8/6/90–2/4/91	2/5/91–2/4/93	2/5/93–11/2/93	11/3/93–5/5/98	5/6/98–10/29/99	From 1/11/99
1	55,000	Light grain	U.S. Gulf to ARA	20%	20%	10%	10%	10%	10%	10%	10%
1A	70,000	T/C	Trans-Atlantic round (duration 45–60 days)			10%	10%	10%	10%	10%	20%
2	52,000	HSS	U.S. Gulf to South Japan	20%	20%	20%	10%	10%	10%	10%	12.5%
2A	70,000	T/C	Skaw Passero[a] to Taiwan–Japan (50–60 days)				10%	10%	10%	10%	12.5%
3	52,000	HSS	U.S. Pacific Coast to South Japan	15%	15%	7.50%	7.50%	7.50%	10%	10%	10%
3A	70,000	T/C	Trans-Pacific Round (35–50 days)			7.50%	7.50%	7.50%	10%	10%	20%
4	21,000	HSS	U.S. Gulf to Venezuela	5%	5%	5%	5%	5%			
5	35,000	Barley	Antwerp to Jeddah (Saudi Arabia)	5%	5%						
	38,000	T/C	South America to Far East			5%	5%	5%			
6	120,000	Coal	Hampton Roads (U.S.) to South Japan	5%	7.50%	7.50%	7.50%	7.50%	7.50%		
7	65,000	Coal	Hampton Roads (U.S.) to ARA	5%	5%	5%					
	110,000	Coal	Hampton Roads (U.S.) to ARA				5%	5%	7.50%	7.50%	
8	130,000	Coal	Queensland (Australia) to Rotterdam	5%	5%	5%	5%	5%	7.50%		
9	55,000	Coke	Vancouver (Canada) to Rotterdam	5%	5%	5%	5%				
	70,000	T/C	Japan–Korea to Skaw Passero (50–60 days)					5%	10%	10%	15%
10	90,000	Iron ore	Monrovia (Liberia) to Rotterdam	5%	5%	5%					
	150,000	Iron ore	Tubarao (Brazil) to Rotterdam				5%	5%	7.50%	7.50%	

TABLE I (Continued)

Baltic Freight Index: Changes in its Composition Since Its Inception*

	Vessel Size (dwt)	*Cargo*	*Route*	*1/1/85–11/3/88*	*11/4/88–8/3/90*	*8/6/90–2/4/91*	*2/5/91–2/4/93*	*2/5/93–11/2/93*	*11/3/93–5/5/98*	*5/6/98–10/29/99*	*From 1/11/99*
11	25,000	Pig iron	Vitoria (Brazil) to China	5%							
	25,000	Phosphate	Casablanca (Morocco) to West Coast India		2.50%	2.50%	2.50%	2.50%			
12	20,000	Potash	Hamburg (Germany) to West Coast India	2.50%							
	14,000	Phosphate	Aqaba (Jordan) to West Coast India		5%	5%	5%	5%			
13	14,000	Phosphate	Aqaba (Jordan) to West Coast India	2.50%							
14	140,000	Iron ore	Tubarao (Brazil) to Beilun and Baoshan (China)							7.50%	
15	140,000	Coal	Richards Bay (U.S.) to Rotterdam							7.50%	

Source: The Baltic Exchange.

ARA, Amsterdam, Rotterdam (The Netherlands), and Antwerp (Belgium) area; HSS, heavy grain, soya, and sorghum; T/C, time-charter routes.

[a]Skaw Passero is the range that extends from Cape Skaw in Denmark to Cape Passero in Sicily (Italy). The countries of the remaining ports are in parentheses.

The following minor amendments of the index are not presented: as of May 6, 1998, routes 2 and 3 refer to a 54,000-dwt panamax vessel; routes 1A, 2A, 3A, and 9 were based on a 64,000-dwt panamax vessel for the period up to February 2, 1996; route 5 was 20,000 dwt vessel Barley from Antwerp to Red Sea for the period January 4, 1985 to February 4, 1986; route 7 was based on a 100,000-dwt vessel for the period February 5, 1991 to February 4, 1993; route 8 was based on a 110,000-dwt vessel for the period January 4, 1985 to February 5, 1992; route 10 was based on a 135,000-dwt vessel for the period February 5, 1991 to August 2, 1995; route 11 was 20,000 dwt sugar from Recife (Brazil) to U.S. east coast for the period January 4, 1985 to May 8, 1986.

routes, who wanted to increase the panamax and capesize representation on the index so as to enhance the performance of their hedges. Finally, the exclusion of the capesize routes from the BFI, follows after an extensive review and consultation of LIFFE with BIFFEX market participants, who "put a Panamax index at the top of their list of requirements," as this is expected to increase the performance of hedges on the panamax routes.[6]

Our data set consists of weekly spot and futures prices from August 1, 1988 to October 31, 1997.[7] The spot price data are Wednesday closing prices of all the BFI constituent routes and the futures prices are Wednesday closing prices of the futures contract, which is nearest to maturity; when a holiday occurs on Wednesday, Tuesday's observation is used in its place. It is assumed that a hedger rolls over to the next nearest contract 1 week before the expiration of the current contract. Spot price data are from LIFFE. Futures prices for the period August 1988 to December 1989 are from Knight Ridder and the Financial Times; for the period January 1990 to December 1997, futures price data are collected from LIFFE.

In order to investigate the effect of BFI revisions on the hedging effectiveness of the futures contract, we consider different estimation intervals for each route; these, along with the contribution of each route to the BFI, are presented in Figure 2. For routes 1, 2, 3, 6, and 8, estimation is carried out over three periods. The first period runs from August 1, 1988 (from November 4, 1988 for route 6) to August 3, 1990 (or to February 2, 1991 for route 2), corresponding to the period before the introduction of the time charter routes. The second period runs from August 6, 1990 (or February 5, 19 91 for route 2) to November 2, 1993 and represents the period after the introduction of the time charter routes in the BFI. The third period runs from November 3, 1993 up to October 31, 1997, corresponding to the period after the exclusion of the handysize routes from the BFI. Routes 1A, 2A, 3A, 7, 9, and 10, are estimated over two periods. The starting observation for these routes is, in the case of routes 1A, 2A, and 3A, the day at which the routes were introduced, (August 6, 1990 or February 5, 1991 for route 2A) and in the case of routes 7, 9, and 10, the day at which the routes were revised to their current definitions (February 5, 1991 or February 5, 1993 for route 9). The second estimation period for these routes runs from November 3, 1993 up to October 31, 1997.

[6]Source: LIFFE to Introduce new BIFFEX Futures and Options Contracts. LIFFE news, LIFFE Internet Site, Friday December 11, 1998.

[7]The choice of a weekly hedging horizon is consistent with the empirical studies in other futures markets; see Kroner and Sultan (1993) and Gagnon and Lypny (1995, 1997).

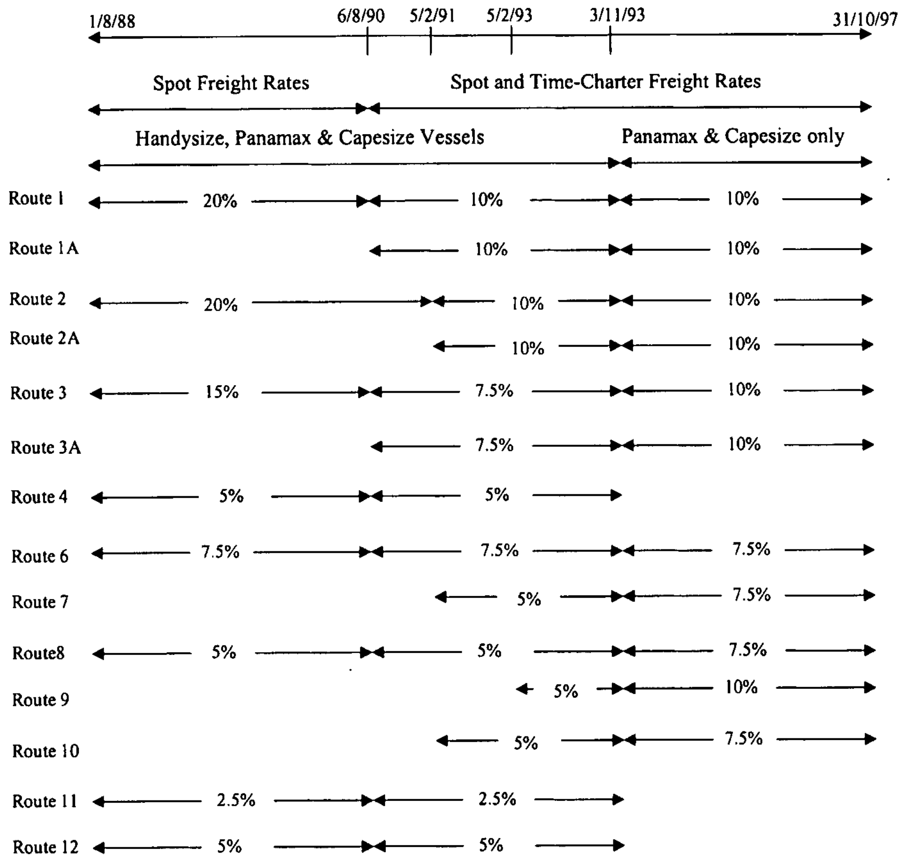


FIGURE 2
Estimation intervals for the Baltic Freight Index (BFI) constituent routes.

The handysize routes (routes 4, 11, and 12) are also estimated over two periods; the first period runs from August 1, 1988 (from November 4, 1988 for route 12) to August 3, 1990 and the second period is from August 6, 1990 to November 2, 1993, when the handysize routes were eventually excluded from the BFI.[8]

EMPIRICAL RESULTS

Summary statistics on logarithmic spot and futures price differences for each route across the different subperiods, indicate the following: first,

[8]Notice that the estimation subperiods are different for each route. These discrepancies in the definitions of the subperiods for each route arise because the revisions in the weights of these routes, or the introduction of new routes, take place at different points in time over the estimation period. However, this helps us identify in greater detail such factors as the revisions in the weights of the routes, the introduction of the time charter routes, or the exclusion of the handysize routes, which affect the hedging performance of each route.

there are significant departures from normality for all return series with the exception of route 3 in the first period, route 9 in the second period, and the futures return in the third period; second, ARCH effects are present in spot returns for some of the subperiods with the exception of routes 2, 2A, and 11, where there is no evidence of ARCH effects in any of the subperiods; finally, Augmented Dickey Fuller (ADF) (Dickey & Fuller, 1981) and PP (Phillips & Perron, 1988) tests on the levels and first differences indicate that the series are first-difference stationary, i.e., $I(1)$ variables.

Having identified that spot and futures prices are $I(1)$ variables, co-integration techniques are used next to examine the existence of a long-run relationship between these series. The results indicate that panamax routes stand in a long-run relationship with futures prices across all the subperiods, with the exception of route 1 in the second period and route 9 for all the periods. On the other hand, the evidence presented for capesize and handysize routes is less strong; capesize routes are co-integrated with BIFFEX prices only in the cases of route 7, route 8 for the third period, and route 10. Moreover, with the exception of route 4 in the first period, none of the handysize routes is co-integrated with the BIFFEX prices.[9] These discrepancies on the existence of a co-integrating relationship between BFI routes and BIFFEX prices across the different classes of vessels can be attributed to the fact that these routes represent different segments of the dry-bulk shipping markets. Since the panamax routes are more heavily represented on the BFI, compared with the capesize and handysize routes, their association with the BIFFEX prices is stronger than in the case of the other routes.

For the routes that are not co-integrated with BIFFEX prices, the estimated statistics indicate that rank(Π) = 0 in eq. (5), implying that a VAR model in first differences is appropriate. For the routes that are co-integrated with the BIFFEX prices, the normalised coefficient estimates of the co-integrating vector, i.e., $\beta' X_{t-1}$ from eq. (5), are used in the joint estimation of the conditional mean and the conditional variance. This takes place by maximising the bivariate conditional normal log-likelihood function. The standard errors of the estimated coefficients are computed using the QMLE procedure of Bollerslev and Wooldridge (1992).[10]

[9]Results for normality, ARCH, unit roots and co-integration tests are available from the authors upon request.

[10]The estimated models are subjected to diagnostics testing to ensure that they are well specified. We employ Ljung-Box Q tests on the standardised residuals and squared standardised residuals, Engle's (1982) tests for ARCH effects and Engle and Ng (1993) tests for asymmetries in the conditional variance equations. These tests do not indicate any misspecification.

Table II presents likelihood ratio tests, comparing the VECM-GARCH time-varying model in eq. (5) and a VECM with constant conditional variance matrix, for each route across the different subperiods.[11] The time-varying GARCH model is preferred over the constant VECM in 18 out of 32 cases (56.25%). The selected model is then used to generate hedge ratios. In the case of the GARCH model, measures of the time-varying variances and covariances are extracted and used to compute the time-varying hedge ratios, γ_t^*, of eq. (4). For the constant VECM, hedge ratios are calculated from the estimated residual series of the model, as the ratio of the residual covariance of the spot and futures equations over the variance of the residuals from the futures equation. Finally, the conventional hedge ratio, γ^* in eq. (2), is also estimated and used for comparison.

To compare formally the performance of each type of hedge, we construct portfolios implied by the computed hedge ratios each week and calculate the variance of the returns to these portfolios over the sample, i.e., we evaluate

$$\mathrm{Var}(\Delta S_t - \gamma_t^* \Delta F_t) \tag{6}$$

where γ_t^* are the computed hedge ratios. The variance of the hedged portfolios is then compared with the variance of the unhedged position, i.e., $\mathrm{Var}(\Delta S_t)$, and the variance reduction achieved through hedging is calculated as follows:

$$VR = 1 - \frac{\mathrm{VAR}(\Delta S_t - \gamma_t^* \Delta F_t)}{\mathrm{Var}(\Delta S_t)} \tag{7}$$

The larger the reduction in the unhedged variance, the higher the degree of hedging effectiveness. When γ_t^* in eq. (7) is the OLS hedge ratio of eq. (2), γ^*, then this measure of hedging effectiveness is the same as the R^2 of eq. (3) (Ederington, 1979). The estimated OLS hedge ratios and their corresponding measures of hedging effectiveness, VR(OLS), as well as the degree of hedging effectiveness, VR in eq. (7), from the selected VECM-GARCH or VECM models in Table II, are shown in Table III.[12]

We can see that the OLS hedge ratios outperform the other hedges in 24 cases out of 33 (72.73%); for the remaining 9 cases, (27.27%), the VECM-GARCH hedges provide higher variance reduction. Therefore,

[11]Owing to the small number of observations for route 9 in the second period ($N = 39$), a GARCH model is not estimated for this route.

[12]To account for the serial correlation and heteroskedasticity appearing in the spot returns, the standard errors of the OLS hedge ratios are computed using the Newey and West (1987) estimator.

TABLE II

Likelihood Ratio Tests of the Estimated VECM-GARCH versus VECM Models on BFI Routes and BIFFEX Prices

Route[a]:	*1*			*1A*		*2*			*2A*		*3*			*3A*		*4*	
Period:	*1*	2	3	2	3	*1*	2	3	2	3	*1*	2	3	2	3	*1*	2
LL (GARCH)[b]	623.1	929.9	1224.7	969.5	1183.3	798.1	821.2	1226.3	834.5	1192.8	640.6	1046.4	1289.7	950.9	1207.4	703.1	1027.7
LL (VECM)[c]	618.9	921.4	1215.8	967.6	1179.1	789.7	820.9	1225.9	834.5	1192.8	640.4	1038.4	1277.9	945.4	1203.3	702.9	1015.6
LR[d]	8.5‡	16.9†	17.8†	4.5	8.5‡	16.8*	0.6	0.8	0.1	0.1	0.4	16.0*	23.5*	11.1†	8.4‡	0.3	24.1*

Route:	*6*			*7*		*8*			*9*		*10*		*11*		*12*	
Period:	*1*	2	3	2	3	*1*	2	3	2	3	2	3	*1*	2	*1*	2
LL (GARCH)	608.9	1096.4	1330.5	819.3	1245.9	703.0	1132.0	1271.4	—	1205.3	915.4	1272.6	602.5	1062.2	586.4	1045.8
LL (VECM)	591.8	1091.3	1320.3	818.6	1240.1	701.9	1119.5	1267.7	237.8	1172.1	909.5	1264.8	600.5	1059.8	585.9	1044.3
LR	34.2*	10.1†	20.3†	1.2	11.5†	2.1	24.9*	7.3	—	66.6*	11.8†	15.5†	4.1	4.8	1.2	2.9

[a] See Figure 2 for the definition of the estimation periods across the different routes.

[b] LL(GARCH) is the maximised value of the log-likelihood function of the GARCH(1, 1) model in equation (5).

[c] LL(VECM) is the maximised value of the log-likelihood function of the model in equation (5) with constant conditional covariance matrix.

[d] LR is the likelihood ratio test for the restriction A = B = O in the conditional variance specification of the VECM-GARCH model of eq. (5). It is computed as two times the difference of the log-likelihood of the GARCH and the VECM; the statistic is $\chi^2(4)$ distributed.

*, †, and ‡ denote that the LR statistic is significantt at the 1%, 5%, and 10% levels, respectively.

TABLE III

OLS Hedge Ratios and Measures of Hedging Effectiveness for the OLS and the GARCH Hedges

Route[a]	*Period*	*1*	2	3	*Route*	*Period*	*1*	2	3
1	γ^* (OLS)	0.8549[a]	0.3604[a]	0.4131[a]	6	γ^* (OLS)	0.1622[a]	0.1066[a]	0.1121[a]
	VR(OLS)	33.70%*	7.17%*	18.64%		VR(OLS)	7.72%	4.48%*	4.73%
	VR(VECM or VECM-GARCH)	31.47%	7.08%	18.69%*		VR(VECM or VECM-GARCH)	10.39%*	4.02%	6.07%*
1A	γ^* (OLS)		0.4281[a]	0.4866[a]	7	γ^* (OLS)		0.0813	0.1565[b]
	VR(OLS)		13.51%*	16.49%		VR(OLS)		1.01%	4.62%
	VR(VECM or VECM-GARCH)		13.33%	17.30%*		VR(VECM or VECM-GARCH)		1.09%*	5.47%*
2	γ^* (OLS)	0.7040[a]	0.3568[a]	0.3839[a]	8	γ^* (OLS)	0.1433[a]	0.0414	0.1471[a]
	VR(OLS)	31.61%*	10.77%*	19.54%*		VR(OLS)	11.20%*	0.88%*	5.65%*
	VR(VECM or VECM-GARCH)	30.70%	10.77%	19.21%		VR(VECM or VECM-GARCH)	10.86%	0.29%	5.61%
2A	γ^* (OLS)		0.3804[a]	0.4438[a]	9	γ^* (OLS)		0.2093[a]	0.4571[a]
	VR(OLS)		13.05%*	17.37%*		VR(OLS)		8.48%*	16.01%
	VR(VECM or VECM-GARCH)		13.01%	17.02%		VR(VECM or VECM-GARCH)		8.41%	16.38%*
3	γ^* (OLS)	0.3907[a]	0.2667[a]	0.2488[a]	10	γ^* (OLS)		0.0965	0.1632[b]
	VR(OLS)	16.05%*	12.38%	13.15%*		VR(OLS)	3.00%	5.65%*	
	VR(VECM or VECM-GARCH)	15.64%	12.60%*	11.10%		VR(VECM or VECM-GARCH)	4.71%*	4.24%	
3A	γ^* (OLS)		0.5151[a]	0.3971[a]	11	γ^* (OLS)	0.0084	0.0439	
	VR(OLS)		17.38%*	14.71%*		VR(OLS)	0.08%*	0.93%*	
	VR(VECM or VECM-GARCH)		16.03%	14.41%		VR(VECM or VECM-GARCH)	0.04%	0.90%	
4	γ^* (OLS)	0.0888[a]	0.1702[a]		12	γ^* (OLS)	0.0626[b]	0.0327	
	VR(OLS)	3.94%*	6.34%*			VR(OLS)	2.65%*	0.46%*	
	VR(VECM or VECM-GARCH)	3.47%	6.10%			VR(VECM or VECM-GARCH)	2.62%	0.45%	

[a]See Figure 2 for the definition of the estimation periods across the different routes.

γ^* is the minimum variance hedge ratio of equation (2). Standard errors for the significance of the estimated coefficient are calculated using the Newey–West (1987) correction for heteroskedasticity and serial correlation, with a truncation lag of 24.

[a], [b], and [c] denote significance at the 1%, 5%, and 10% levels, respectively. An asterisk (*) denotes the model with the largest variance reduction.

VR(OLS) is the variance reduction of the spot position achieved using the conventional hedge ratio, γ^*. It is equivalent to the R^2 of eq. (2).

VR(VECM or VECM-GARCH) is the variance reduction of the spot position of eq. (7) achieved using the hedge ratio generated by the selected model in Table II.

the use of time-varying hedge ratios, while statistically appropriate, provides little gains in reducing the variability of the hedged returns, compared to the OLS hedges. A possible explanation for this is that the time variability in the second moments of spot and futures returns is not large to the extent of justifying the use of time-varying hedging strategies. In line with this article, Myers (1991) and Garcia et al. (1995) also find that there are no gains in variance reduction by using time-varying hedge ratios, in the wheat and soybean futures markets, respectively. This suggests that the additional complexity of specifying and estimating GARCH models may be justified for some commodities but not for others.

The degree of hedging effectiveness seems to differ substantially across subperiods and shipping routes. Consider first the periods before and after the introduction of time charter routes to the index (periods 1 and 2). The hedging effectiveness decreases in all cases, with the exception of routes 4 and 11, where the hedging effectiveness increases from 3.94% to 6.34% and from 0.08% to 0.93%, respectively. The reduction in hedging effectiveness is more striking for routes 1, 2 and 8. More specifically, for routes 1 and 2 the degree of hedging effectiveness is reduced by almost 3 times (from 33.7% to 7.17% and from 31.61% to 10.77%, respectively), and for route 8, the degree of hedging effectiveness is reduced from 11.20% to 0.88%; notice also that the OLS hedge ratio for route 8 in the second period becomes statistically insignificant, which implies that the futures contract has no power in reducing the riskiness of the spot position.

Consider next the period after the exclusion of the handysize routes. The hedging effectiveness increases in all cases, with the exception of route 3A, where the hedging effectiveness decreases from 17.38% to 14.71%. The highest increase is evidenced in routes 1, 2, and 9, where the degree of hedging effectiveness increased almost twofold with respect to period 2; notice that the weight of route 9, increased from 5% to 10% while, the weights of routes 1 and 2 remained the same. Notice also that the hedge ratios for routes 7, 8, and 10 in the third period become statistically significant when compared to the second period.

Another feature of our results is that, on average, the variance reduction for the panamax routes is higher than that for the handysize and capesize routes. For instance, in the first period the lowest hedging effectiveness for a panamax route is 16.05% in route 3, while the highest hedging effectiveness for a handysize route is 3.94% in route 4 and for a capesize route is 11.20% in route 8. The capesize routes seem to perform better than the handysize routes in the first period, although this is reversed in the second period. On average, however, their performance is poor, compared to that of the panamax routes. This follows from the fact that these three categories of vessels respond differently to fluctuations in the demand for the transportation of different commodities. For instance, an increase in the demand for transporting iron ore and coal will affect the freight rates of capesize vessels. Similarly, an increase in the demand for the transportation of grain will primarily affect the freight rates of panamax vessels. Hence, these categories of vessels represent different segments of the dry-bulk market which are weakly correlated with each other. In addition, the heavier representation of the panamax routes on the BFI can explain the superior hedging performance on these routes.

Finally, and in line with the results of Kavussanos and Nomikos (2000a), the futures contract fails to eliminate a large proportion of the variability of the unhedged portfolio, the greatest variance reduction achieved in route 1 for the first period (33.70%). This is well below the variance reduction over the unhedged position evidenced in other markets: 57.06% for the Canadian Interest rate futures (Gagnon & Lypny, 1995), 69.61% and 85.69% for the corn and soybean futures (Bera et al., 1997), and 97.91% and 77.47% for the SP500 and the Canadian Stock Index futures contract (Park & Switzer, 1995), and is a reflection of the heterogeneous composition of the underlying index.[13]

COMPARISON OF THE HEDGING PERFORMANCE ACROSS SUBPERIODS

The preceding analysis indicates that there are nominal differences in the effectiveness of the hedging mechanism, following the revisions in the composition of the BFI. To consider whether the observed nominal differences across the subperiods are statistically significant, we construct empirical confidence intervals for the differences in measures of hedging effectiveness across subperiods, using bootstrapping techniques. Bootstrapping is a data-based simulation method that uses the empirical distribution of the statistic of interest, rather than the theoretical distribution implied by statistical theory, to conduct statistical inference. It is particularly useful in cases where the standard error of the statistic is difficult to estimate analytically; this is the case when the statistic of interest is the difference in the degree of hedging effectiveness, as indicated by Li and Vukina (1998).[14]

[13]Owing to the small number of observations for each route, particularly in the first two subperiods, out-of-sample tests for hedging effectiveness are not pursued here.

[14]The difference between bootstrapping and other simulation methods, most notably Monte Carlo,

We use the following procedure. For each period, we select the model that provides the highest degree of hedging effectiveness (either the GARCH, the VECM or the OLS) and compute the variance reduction achieved through hedging, using eq. (7). This is denoted VR_i, $i = 1,2,3$, for the different periods. The statistic of interest is the difference of the VR_i across two subperiods, $VR_i - VR_{i-1}$ (i.e., $VR_2 - VR_1$ and $VR_3 - VR_2$). The selected model for each period, along with the value of the observed statistic, are presented in Table IV. To investigate whether the observed differences are statistically significant, we draw independent bootstrap samples with replacement from the four sample series of interest, namely the returns of the hedged ($\Delta S_t - \gamma_t^* \Delta F_t$, where γ_t^* is the selected hedge ratio) and the unhedged position (i.e., the returns in the spot market, ΔS_t) across two adjacent periods.

Based on these bootstrap samples, we calculate the variances of the series and compute the variance reduction from hedging for each subperiod, i.e., VR_i using eq. (7); these are then used to compute the bootstrap replication statistic $VR_i - VR_{i-1}$. This procedure is repeated 10,000 times and generates a series of 10,000 observations of the bootstrap statistic. Visual inspection of the histograms from the bootstrap replications suggests that they are asymmetrical. Hence, the confidence intervals are calculated using the bias-corrected and accelerated (BCa) method suggested by Efron and Tibshirani (1993). This is because the BCa method, by accounting for the asymmetrical nature of the distribution of our bootstrap statistics, generates more robust confidence intervals, compared to the conventional percentile method. If the value of 0 falls within the limits of the empirical confidence interval, then the difference in the degree of hedging effectiveness across the subperiods is not significant.[15]

Our results, presented in Table IV, indicate that the differences in the degree of hedging effectiveness between periods 1 and 2 are significant only for routes 1 and 2. Notice that during the first period, routes 1 and 2 are the most heavily represented routes on the BFI, with a weight of 20% each, and the BIFFEX contract offers the greatest variance reduction for hedges on these routes, as evidenced in Table III, whereas in the second period the weights of these routes are halved to 10% each.

This indicates that the reduction in the weights of the BFI routes, affects significantly the hedging performance for these routes which are

is that the former is based on the actual empirical data to obtain a description of the sample properties of the empirical estimators. In contrast, Monte Carlo methods are used to identify whether the theoretical properties of estimators, implied by statistical theory, conform to a particular data set. For an overview of the applications of bootstrapping, see Efron and Tibshirani (1993).

[15]Our results remain qualitatively the same when we consider the percentile confidence intervals and across the 10% and 1% levels of significance.

TABLE IV

Empirical Confidence Intervals for the Differences in the Degree of Hedging Effectiveness across Subperiods

	Selected Model for Period[b]				*Periods 1–2*	*Periods 2–3*
Route[a]	*1*	2	3		VR_2–VR_1	VR_3–VR_2
1	OLS	OLS	GARCH	Obs[c]	−0.2653†	0.1152†
				95% CI[d]	−0.4449 −0.0681	0.0234 0.2153
1A	—	OLS	GARCH	Obs	—	0.0374
				95% CI		−0.0819 0.1693
2	OLS	OLS	OLS	Obs	−0.2084†	0.0877
				95% CI	−0.3392 −0.0619	−0.0476 0.2046
2A	—	OLS	OLS	Obs	—	0.0432
				95% CI		−0.1089 0.0752
3	OLS	GARCH	OLS	Obs	−0.0345	0.0055
				95% CI	−0.2235 0.1452	−0.1501 0.1691
3A	—	OLS	OLS	Obs	—	−0.0267
				95% CI		−0.2007 0.1331
4	OLS	OLS	—	Obs	0.0239	—
				95% CI	−0.0693 0.1129	
6	GARCH	OLS	GARCH	Obs	−0.0604	0.0159
				95% CI	−0.0545 0.0901	−0.0789 0.0864
7	—	GARCH	GARCH	Obs	—	0.0436
				95% CI		−0.0367 0.1061
8	OLS	OLS	OLS	Obs	−0.1031	0.0476
				95% CI	−0.2661 0.0938	−0.0214 0.1117
9	—	OLS	GARCH	Obs	—	0.0791
				95% CI		−0.1024 0.3103
10	—	GARCH	OLS	Obs	—	0.0095
				95% CI		−0.0980 0.0973
11	OLS	OLS	—	Obs	0.0085	—
				95% CI	−0.0248 0.0412	
12	OLS	OLS	—	Obs	−0.0220	—
				95% CI	−0.1267 0.0372	

OLS, Ordinary Least Squares; GARCH, generalized autoregressive heteroskedasticity; Obs, observed; 95% CI, 95% confidence interval.

[a]See Figure 2 for the definition of the estimation periods across the different routes.

[b]The selected model for each period is the model, in Table II, that provides the highest degrees of hedging effectiveness.

[c]Observed is the observed difference in the degree of hedging effectiveness from the selected models between two adjacent periods.

[d]95% CI is the 95% BCa confidence interval for the differences in the degree of hedging effectiveness.

*, †, ‡ denote significance at the 1%, 5%, and 10% levels, respectively.

more heavily represented on the Index. Two reasons may be put forward in support of this. First, for route 3, whose weight in the BFI in the first period was less than that of routes 1 and 2 (15%, cf. 20%), the reduction in the weight of this route to 7.5% in the second period did not result in a significant reduction in its hedging performance. Second, the hedging performance of the routes whose weights remained unchanged between periods 1 and 2 (i.e., the handysize routes 4, 11, and 12 and the capesize routes 6 and 8) has not been altered significantly between these periods. This also suggests that the introduction of the time charter routes to the index did not affect the hedging performance of these routes and, as a consequence, does not account for the observed decrease in the hedging performance of routes 1 and 2.

Therefore, the decrease in the hedging performance for routes 1 and 2 can be attributed to the reduction of their weights in the BFI. This finding contrasts with the evidence presented by Glen and Rogers (1997), who argue that changing the weights in the composition of the constituent routes in the Simpson, Spencer, and Young (SSY) capesize index does not alter significantly the correlation of these routes with the index. It appears that the opposite is true for the BIFFEX contract, since futures prices are more strongly correlated with the shipping routes that have a heavier representation on the underlying index.

Consider next the period after the exclusion of the handysize routes. We can see that there is a nominal increase in the hedging effectiveness for all the routes, with the exception of route 3A. However, this increase is statistically significant only in the case of route 1. Therefore, the exclusion of the handysize routes from the index did not affect the hedging performance across all the BFI routes, as it was anticipated by the regulatory authorities; this may be attributed to the fact that the handysize routes represented a small proportion, only 17.5%, of the index.

Despite this, it seems that the hedging performance of the BIFFEX contract will strengthen even further after the forthcoming exclusion of the capesize routes and the introduction of the BPI as the underlying asset of the contract. Three reasons may be put forward for this. First, the weights of five out of the seven routes of the new index will increase from their current level and the weights of the other two routes will remain unchanged (see Table I). This suggests that the hedging performance, at least for the former routes, will increase since, as our results indicate, freight rate risk can be hedged more effectively for these routes which have a heavier representation on the BFI. A second reason is that the capesize routes currently represent 30% of the index and, as a result, their exclusion from the BFI will have a more beneficial effect on the

TABLE V

Correlation Matrix of the Logarithmic Differences between BFI and BFI Routes: Daily Data*

Route	*1*	*1A*	*2*	*2A*	*3*	*3A*	*6*	*7*	*8*	*9*	*10*	*BFI*
1	—											
1A	0.693	—										
2	0.555	0.624	—									
2A	0.602	0.769	0.719	—								
3	0.384	0.403	0.413	0.445	—							
3A	0.391	0.505	0.432	0.555	0.602	—						
6	0.214	0.261	0.183	0.224	0.201	0.272	—					
7	0.232	0.306	0.203	0.247	0.171	0.236	0.560	—				
8	0.173	0.157	0.110	0.120	0.123	0.193	0.277	0.273	—			
9	0.376	0.486	0.409	0.493	0.585	0.716	0.225	0.235	0.195	—		
10	0.268	0.300	0.207	0.247	0.151	0.218	0.556	0.702	0.293	0.191	—	
BFI	0.729	0.832	0.756	0.841	0.656	0.752	0.430	0.459	0.296	0.715	0.449	—

Arithmetic Mean of the Correlation Coefficients

BFI with panamax routes	=	0.755	Panamax with panamax routes	=	0.531
BFI with capesize	=	0.408	Capesize with capesize routes	=	0.444
			Panamax with capesize routes	=	0.209

BFI, Baltic freight route.*The correlation coefficients are based on logarithmic first differences of the daily prices of the BFI and the BFI routes for the period November 3, 1993 to October 31, 1997.

hedging performance of the market, compared with the exclusion of the handysize routes, which represented only 17.5% of the index.

Finally, the new index will have a more homogeneous structure than the BFI and will consist of shipping routes, which are strongly correlated with each other. Table V presents the correlation matrix of the logarithmic price differences of the BFI with the BFI routes for the period November 3, 1993–October 31, 1997, corresponding to the third period in our analysis (see Fig. 2). The same table, also presents the arithmetic mean of the correlation coefficients across different categories of vessels in the BFI. We can see that the panamax routes are more strongly correlated with the BFI than the capesize routes. More specifically, the average correlation of the panamax routes with the BFI, 75.5%, is higher than that of the capesize routes, 40.8%; this is expected given that the former represent 70% of the index.

We can also note that these two categories of vessels represent two distinct groups of underlying shipping routes. The within-group correlation is strong in both cases although their correlation with the other group is weak; the average correlation for panamax routes is 53.1%, while the average correlation for the capesize routes is 44.4%. On the other hand, the average correlation between panamax and capesize routes is only

20.9%. Gemmill (1985) indicates that the inclusion in the BFI of routes which are weakly correlated with the remaining routes of the index, such as for instance the capesize routes which are weakly correlated with the panamax routes, will not improve much the performance of hedges on the capesize routes and may actually deteriorate the performance of hedges on the panamax routes. For these reasons, it seems that the introduction of the BPI as the new underlying asset of the BIFFEX contract is likely to have a beneficial impact on the risk management function of the market.

CONCLUSIONS

This article discusses the hedging effectiveness of futures contracts whose underlying asset is an index, when the structure of this index is changing. The case of the BIFFEX contract is investigated here. The contract is cash settled against the BFI, a weighted average index of the freight rates on 11 component shipping routes. These routes are diverse, in terms of cargoes, vessel sizes and types of fixture and, as a result, futures contracts are not as effective in eliminating spot market risk to the extent evidenced in other commodity and financial futures markets. For instance, Kavussanos and Nomikos (2000) find that the hedging effectiveness of the BIFFEX contract varies from 19.2% to 4.0% across different shipping routes of the BFI, whereas, the evidence presented in other markets indicates variance reductions by as much as 98%. In view of that poor hedging performance, LIFFE has altered the structure of the BFI on a number of occasions.

Our findings can be summarised as follows. There is no evidence of statistical change in the hedging performance of BIFFEX on any route, following the inclusion of the time charter routes in the BFI, however, there seems to be an increase in the performance of route 1 following the exclusion of the handysize routes. In addition, our analysis has indicated that the reduction in the weight of a route in the BFI (as in the case of the weights of routes 1 and 2) is likely to decrease hedging performance on this route. Furthermore, it may be argued that, the exclusion of the capesize routes from the index, in November 99, with the brief of increasing the homogeneity of the BFI, will have a beneficial impact on the hedging performance of the market; see Kavussanos and Nomikos (2000b) for the latest evidence on this.

Policy implications for the construction of an index, such as the BFI, for hedging point to including (1) routes with high correlation between them; (2) routes which are affected by similar economic factors (e.g., routes within the same subsector of the dry bulk market); and (3) few routes only, resulting in relatively high weight of each. However, hedging performance will be further improved if futures contracts are based on an individual route rather than on an underlying index.

BIBLIOGRAPHY

Bera, A., Garcia, P., & Roh, J. (1997). Estimation of time-varying hedge ratios for corn and soybeans: BGARCH and random coefficients approaches. Office for Futures and Options Research, 97-06.

Bollerslev, T., & Wooldridge, J. M. (1992). Quasi-maximum likelihood estimation of dynamic models with time varying covariances. Econometric Reviews, 11, 143–172.

Bollerslev, T., Engle, R. F., & Nelson, D. B. (1994). ARCH models. In Engle R. F., & McFadden V. (Eds.). Handbook of econometrics. Amsterdam: North-Holland, pp. 2959–3038.

Cheung, Y., & Lai, K. (1993). Finite sample sizes of Johansen's likelihood ratio tests for co-integration. Oxford Bulletin of Economics and Statistics, 55, 313–328.

Chou, W., Denis, K., & Lee, C. (1996). Hedging with the Nikkei index futures: The conventional versus the correction model. Quarterly Review of Economics and Finance, 36, 495–505.

Cullinane, K., Mason, K., & Cape, M. (1999). A comparison of models for forecasting the Baltic Freight Index: Box-Jenkins revisited. International Journal of Maritime Economics, Vol. 1 (special edition)., No 2, 15–39.

Dickey, D., & Fuller, W. (1981). Likelihood ratio statistics for autoregressive time series with a unit root. Econometrica, 49, 1057–1072.

Ederington, L. H., (1979). The hedging performance of the new futures markets. Journal of Finance, 34, 157–170.

Efron, B., & Tibshirani, R. (1993). An introduction to the bootstrap. Monographs on statistics and applied probability. New York: Chapman & Hall.

Engle, R. F. (1982). Autoregressive conditional heteroskedasticity with estimates of the variance of United Kingdom inflation. Econometrica, 50, 987–1008.

Engle, R. F., & Granger, C. W. (1987). Co-integration and error correction: Representation, estimation, and testing. Econometrica, 55, 251–276.

Engle, R. F., & Kroner, K. F. (1995). Multivariate simultaneous generalised ARCH. Econometric Theory, 11, 122–150.

Engle, R. F., & Ng, V. K. (1993). Measuring and testing the impact of news on volatility. Journal of Finance, 48, 1749–1778.

Figlewski, S. (1984). Hedging performance and basis risk in stock index futures. Journal of Finance, 39, 657–669.

Franckle, C. (1980). The hedging performance of the new futures markets: Comment. Journal of Finance, 35, 1273–1279.

Gagnon, L., & Lypny,.G. (1995). Hedging short-term interest risk under time-varying distributions. Journal of Futures Markets, 15, 767–783.

Gagnon, L., & Lypny, G. (1997). The benefits of dynamically hedging the Toronto 35 stock index. Canadian Journal of Administrative Science, 14, 52–68.

Garbade, K., & Silber, W. (1983). Price movements and price discovery in futures and cash markets. Review of Economics and Statistics, 65, 289–297.

Garcia, P., Roh. J., & Leuthold, M. (1995). Simultaneously determined, time-varying hedge ratios in the soybean complex. Applied Economics, 27, 1127–1134.

Gemmill, G. (1985). The behaviour of the Baltic freight index. London, UK: City University Business School. Unpublished manuscript.

Ghosh, A. (1993). Hedging with stock index futures: Estimation and forecasting with error correction model. Journal of Futures Markets, 13, 743–752.

Glen, D., & Rogers, P. (1997). Does weight matter? A statistical analysis of the SSY capesize index. Maritime Policy and Management, 24, 351–364.

Gonzalo, J. (1994). Five alternative methods of estimating long-run equilibrium relationships. Econometrica, 60, 203–233.

Gray, J. (1990). Shipping futures. London, UK: Lloyd's of London Press.

Johansen, S. (1988). Statistical analysis of co-integrating vectors. Journal of Economic Dynamics and Control, 12, 231–254.

Johnson, L. (1960). The theory of hedging and speculation in commodity futures. Review of Economic Studies, 27, 139–151.

Kavussanos, M., & Nomikos, N. (1999). The forward pricing function of the shipping freight futures market. Journal of Futures Markets, 19, 353–376.

Kavussanos, M., & Nomikos, N. (2000a). Constant vs. time-varying hedge ratios and hedging efficiency in the BIFFEX market. Transportation Research, Part E (Logistics and Transportation Review), 36(4), 229–248.

Kavussanos, M., & Nomikos, N. (2000b). Short run deviations volatility and hedging in the freight futures market. Journal of Derivatives, (forthcoming) 2000.

Kroner, K., & Sultan, J. (1993). Time-varying distributions and dynamic hedging with foreign currency futures. Journal of Financial and Quantitative Analysis, 28, 535–551.

Lee, T., & Tse, Y. (1996). Co-integration tests with conditional heteroskedasticity. Journal of Econometrics, 73, 401–410.

Li, D., & Vukina, T. (1998). Effectiveness of dual hedging with price and yield futures. Journal of Futures Markets, 18, 541–562.

Lien, D. (1996). The effect of the co-integration relationship on futures hedging: A note. Journal of Futures Markets, 16, 773–780.

Lindahl, M. (1992). Minimum variance hedge ratios for stock index futures: Duration and expiration effects. Journal of Futures Markets, 12, 33–53.

Ljung, M., & Box, G. (1978), On a measure of lack of fit in time series models. Biometrika, 65, 297-303.

Malliaris, A., & Urrutia, J. (1991). The impact of the lengths of estimation periods and hedging horizons of the effectiveness of a hedge: Evidence from foreign currency futures. Journal of Futures Markets, 11, 271–289.

Myers, R. (1991). Estimating time-varying optimal hedge ratios on futures markets. Journal of Futures Markets, 11, 39–53.

Myers, R., & Thompson, S. (1989). Generalised optimal hedge ratio estimation. American Journal of Agricultural Economics, 71, 858–868.

Newey, W., & West, K. (1987). A simple positive semi-definite, heteroskedasticity and autocorrelation consistent covariance matrix. Econometrica, 55, 703–708.

Park, T., & Switzer, L. (1995). Bivariate GARCH estimation of the optimal hedge ratios for stock index futures: A note. Journal of Futures Markets, 15, 61–67.

Phillips, P., & Perron, P. (1988). Testing for a unit root in time series regressions. Biometrica, 75, 335–346.

Stein, J. (1961). The simultaneous determination of spot and futures prices. American Economic Review, 51, 1012–1025.

Tong, W. (1996). An examination of dynamic hedging. Journal of International Money and Finance, 15, 19–35.