# Efficiency and hedging effectiveness in the NYMEX crude oil futures market

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# *Efficiency and hedging effectiveness in the NYMEX crude oil futures market*

*Tarkan Çavuşoğlu*[*] , *Soner Gökten*[**]

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***Abstract .***
*This study aims to investigate the speculative efficiency of the New York Mercantile Exchange (NYMEX) Light Sweet Crude Oil futures market and the effectiveness of these futures contracts in hedging the West Texas Intermediate (WTI) crude oil price risk. The period of interest ranges between October 2001 and August 2006, coinciding with the beginning of an oil price surge following the low-price period of the 1980s and 1990s. Our empirical findings imply that the NYMEX futures market is not an efficient market in the Fama sense for the October 2001-August 2006 period. Moreover, although the time-varying ratios are found to be slightly above the constant one in most of the sample period, the relative hedging effectiveness values based on the portfolio variances of the two hedge ratios are not different from each other in statistical terms.*
***Keywords:*** *Hedging, Hedge Ratio, Crude Oil Futures, Bivariate GARCH*
***JEL Classification:*** *C32, G13, G14*

***Özet.***
***NYMEX Ham Petrol Vadeli İşlem Sözleşmeleri Pazarında Piyasa ve Korunma Etkinlikleri***
*Çalışmada New York Ticaret Borsası (NYMEX) Ham Petrol vadeli işlem sözleşmeleri pazarının spekülatif etkinliği ve Batı Teksas Tipi (WTI) ham petrol vadeli işlem sözleşmeleri korunma etkinliği incelenmiştir. Araştırmada, 1980 ve 1990'ların statik ve düşük fiyat düzeylerinden sonra sert yükselişlerin başlangıcına denk gelen Ekim 2001 – Ağustos 2006 dönemi dikkate alınmıştır. Amprik bulgularımız NYMEX vadeli işlem sözleşmelerinin söz konusu dönem için Fama hipotezi çerçevesinde etkin olmadığını göstermektedir. Bununla birlikte, her ne kadar örneklemin büyük bir bölümünde, zaman-değişimli korunma etkinliği oranlarının sabit korunma etkinliği oranın üzerinde değerler aldıkları tespit edilse de, söz konusu bu iki tip korunma oranı ile hesaplanan portföy varyanslarını temel alan korunma etkinliği oranlarının nisbi değerlerinin istatistiksel anlamda farklı olmadıkları görülmektedir.*
***Anahtar Kelimeler:*** *Korunma, Korunma Etkinliği, Ham Petrol Vadeli İşlem Sözleşmesi, İki Değişkenli GARCH*
***JEL Sınıflaması:*** *C32, G13, G14*

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## 1. Introduction

Investors' need for better risk-hedging strategies in the commodity markets brings out numerous theoretical and empirical works in the finance literature. The characteristics of market expectations and methods of optimal hedging are the leading issues in this literature. The crude oil market is one of the most volatile commodity markets in the world, where sharp price movements generate significant micro- and macro-economic imbalances, even in advanced economies. In this sense, crude oil futures markets are a way for producers, distributors and consumers to eliminate or manage spot oil-price risks of wide fluctuations in the physical crude oil markets. However, the increasing interaction between the physical and futures commodity markets, and their link with assets traded at financial markets results in the complexity of the hedging strategies for investors.

This study aims to investigate the speculative efficiency of the NYMEX light sweet crude oil futures market and the effectiveness of these futures contracts in hedging the WTI crude oil price risk. The period of interest ranges between October 2001 and August 2006, because this period coincides with the beginning of an oil price surge following the low-price period of the 1980s and 1990s. According to the World Oil Outlook 2008, in this low-price period oil producers scaled down their investments, began to implement cost-cutting strategies and reduced their R&D spending. However, when the global economic growth reached above-trend rates, the world was caught unprepared to the increased energy demand and oil prices began to surge (OPEC, 2008: 5).

In the following sections of the study, a brief background is given about the magnitude of hedging activities in commodity and crude oil futures markets in Section 2, followed by the review of the empirical literature on hedging in crude oil markets in Section 3. The methodology used for estimations in the study is described briefly in Section 4. In Section 5, first, the information about the data series is presented, followed by some univariate descriptive statistics. Then, the estimation results about market efficiency, hedge ratios and hedging effectiveness are presented respectively. The conclusion is made in Section 6.

## 2. A Background on Crude Oil Futures Markets

The size of the hedging activity worldwide is reported by various institutions with respect to the measurable amount of derivatives traded. Trading occurs not only in organized futures exchanges but also on exempt commercial and over-the-counter (OTC) markets. The opacity involved in unorganized markets hinders the acquisition of comprehensive information and therefore, the actual size of the derivatives sector cannot be quantified.

Nevertheless, the Bank for International Settlements (BIS) regularly reports statistics about OTC derivatives contracts. According to these statistics, the total notional amount of OTC derivatives contracts was $ 683.8 trillion by June 2008, but fell to $ 548 trillion by December 2008 due to the recent financial crisis. By December 2009, with the latest published figure of $ 614.7 trillion, the amount still had not recovered to the pre-crisis levels. These figures were almost thirteen times the amount reported in the BIS statistics by the end of March 1995, and more than ten times the world gross domestic product in 2009, implying a tremendous growth of hedging through off-exchange trade. Similarly, hedging through commodity contracts showed a sharp surge starting from 2005 until the recent global recession, according to the figures of the BIS. The share of commodity contracts in the total OTC derivatives, which was about 0.60 % in 1998-2004 period, reached 1.9 % by June 2008, but dropped to 0.48 % by December 2009 in the post-crisis period. The notional amounts outstanding for commodity contracts were $ 443.1 billion by June 1998, $ 13.2 trillion by June 2008 and $2.9 trillion by December 2009.

About exchange-traded derivatives, on the other hand, the trading volume surveys of the Futures Industry Association provide contract volume data on global futures, options on futures and stock indexes, interest rates and currency contracts. These surveys show that the volume of futures and options contracts have risen almost eightfold globally since 1998, i.e., from 2.2 to 17.7 billion of contracts by 2008. Contracts on commodities such as agriculturals, precious and industrial metals, and energy products constitute about 10 % of the total on the average. The aggregate volume of these commodity derivatives also rose sharply, from 343.6 million in 1998 to 1.8 billion in 2008. The share of the energy products in the total of these commodity contracts varied over a range between 30-35 %, while about one-fifth of the energy contracts were the crude oil futures traded at the New York Mercantile Exchange (NYMEX). The light sweet crude oil futures and options contracts traded at the NYMEX increased from 37.5 million in 2001 to 134.7 million in 2008, which coincides with the period of higher prices in the physical markets for crude oil in the world. According to the statistics based on the Commitments of Traders Reports published by the U.S. Commodity Futures Trading Commission (CFTC), the total open interest and the number of reportable traders in the NYMEX light sweet crude oil futures rose dramatically in the 2000s. Figure 1 illustrates that the average weekly open interest figures reached to millions in the last years, and the average number of traders that held positions in the NYMEX crude oil futures began to follow a continuous upward trend from 2001 onwards, in parallel with the open interest figures. The increasing open interest together with rising crude oil prices in the 2000s implies that new funds have been flowing into the NYMEX crude oil market for hedging and speculative purposes.

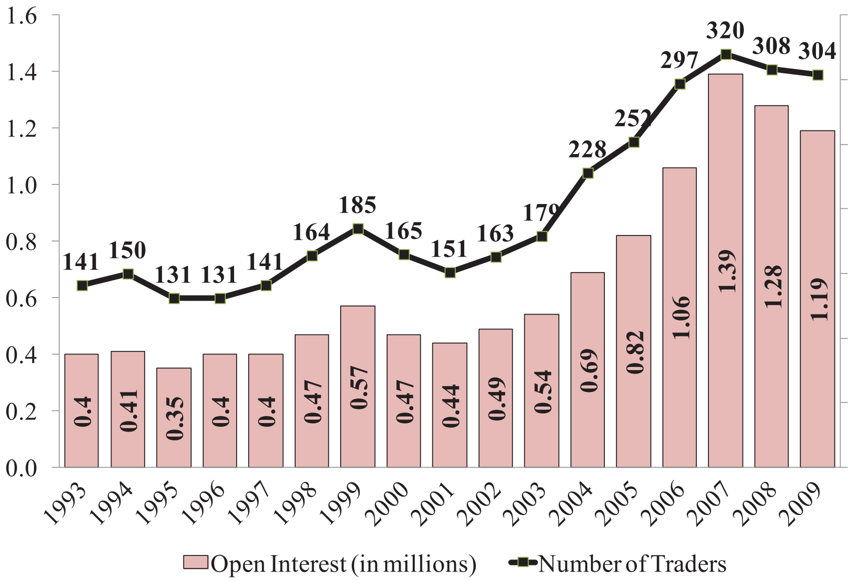


**Figure 1. Statistics on the NYMEX light sweet crude oil futures**

## 3. The Empirical Literature

Some recent studies may help understanding the interrelated economic linkages behind hedging with futures. Although the experimental data analysis of Noussair and Tucker (2006) suggests that futures markets reduce spot market price bubbles, Domanski and Heath (2007) draw attention to the growing resemblance between the commodity and financial markets in terms of the motivations and strategies of participants, resulting from the rapid increase in commodity derivatives. In the latter study, it is argued that the role of financial investors and the speculation in commodity markets has increased in recent years, and hence, the determinants of market liquidity may resemble those in traditional financial markets. On this issue, Roll et al. (2007) show that while illiquidity in financial markets prevents futures prices from converging to spot prices in the New York Stock Exchange, the divergence of futures prices from spot prices also exhibit a predictive power on future shifts in the market liquidity. Concerning the spot and futures linkage, the studies of Coppola (2007) and Bekiros and Diks (2008) analyze the interaction between spot and futures markets for crude oil, i.e., the West Texas Intermediate (WTI) and the NYMEX. However, while according to the former, information provided by the oil futures market is found to explain a sizable proportion of the spot oil-price movements, according to the latter, the causal linkage is found significant only from crude oil spot to futures prices, with respect to the linear and non-linear causality analyses. Similarly, Kasman and Kasman (2008), beside their main finding that the introduction of stock index futures reduces the volatility of stocks traded at the Istanbul Stock Exchange, report the presence of a unidirectional Granger causality running from spot market to futures market both in the short-run and the long-run. Aksoy and Olgun, (2009) tried to estimate the optimal hedge ratio for ISE-30 stock index futures contract, traded in Turkish Derivatives Exchange by using the conventional regression model, the error correction model (ECM) and the GARCH model. The results of their study implied that, the hedge ratio obtained from the GARCH model achieves minimum portfolio variance by outperforming other models' estimates in both horizons. Lee et al. (2009) analyze the spot-futures interaction from a different perspective, by analyzing the influence of the increase in the length of the batching period of the stock closing call in the Taiwan Stock Exchange. The study puts forth that the batching period arrangement not only improves the price efficiency of the stock indices futures but also enhance the hedging performance in terms of hedging risks by reducing the manipulation of stock and futures prices.

Many of the commodity futures may have strong linkages among themselves as complements or substitutes both for the industrial production and for the traders of the commodity options and futures. In this context, Hammoudeh and Yuan (2008) illustrate empirically the influence of changes in the NYMEX crude oil futures prices on the volatility of gold-, silver- and copper-futures prices. Li and Zhang (2009) investigate the cross-market information transmission between the two copper futures markets, i.e., the Shanghai Futures Exchanges (SFE) and the London Metals Exchange (LME), and find a long-run relationship in which the influence of the LME futures on the SFE futures is stronger than that of the SFE futures on the LME futures. Similarly, Chng (2009) argues the presence of cross-market influences between the futures contracts of complementary commodities, such as rubber, palladium and gasoline. The study strikingly demonstrates that multi-commodity hedging based on complementary commodities results in better hedging than a commodity-by-commodity hedging. Moreover, in the study, the trading strategies that include cross-market interaction among related commodities are only found to translate into positive economic profits.

In an environment where hedging with futures is becoming more and more complex due to the interrelations within and between different commodity and asset markets, the issues of market efficiency and hedging effectiveness preserve their priority on the agenda for finance professionals. In this respect, a substantial literature exists on testing the speculative market efficiency hypothesis with different approaches. However, market efficiency in crude oil markets has not been much analyzed because the trade of crude oil futures contracts in NYMEX began in 1983. Moreover, only a few of the existing studies cover recent periods, when both the market was more developed than before and global conditions had changed. Among the studies covering only the 1980s and 1990s for the NYMEX crude oil futures contracts, Bopp and Sitzer (1987), Bopp and Lady (1991), Crowder and Hamid (1993), Gülen (1998), and Peroni and McNown (1998) present some empirical evidence in favour of efficiency. On the other hand, Quan (1992), Deaves and Krinsky (1992), Moosa and Al-Loughani (1994), Fujihra and Mougoue (1997), and Shambora and Rossiter (2007) reach no evidence on market efficiency. With samples extending to the 2000s, the empirical analyses of Abosedra and Baghestani (2004) and Switzer and El-Khoury (2007) provide findings supporting the efficiency hypothesis. Alvarez-Ramirez et al. (2008), through a different technique developed in statistical physics named the 'detrended fluctuation analysis' and the use of only the daily spot prices between 1987-2007, conclude that the crude oil market, represented by the WTI, tends towards an efficiency regime at long time horizons.

Measuring the effectiveness of hedging with futures contracts is the other challenging issue of the relevant literature. However, there are a limited number of studies that focus on crude oil markets. To our knowledge, the first empirical work on crude oil markets is by Chen et al. (1987), who present the finding that a substantial part of the spot price risk can be eliminated through futures contracts. This finding is based on the minimum-risk hedge ratio measure suggested by Ederington (1979) and others. Lindahl (1989) argues that $R^2$ is a consistent measure of hedging effectiveness in crude oil markets only when comparing hedges with different futures price data but the same spot price data. A more recent work by Lien and Schaffer (2002) introduces the multi-period hedging strategy instead of the Ederington-type single-period strategy. In this context, the effectiveness of two types of multi-period approaches, called the 'strip hedge' and the 'stack-and-roll hedge', are compared, and the conclusion is that the two strategies perform equally well. Veld-Merkoulova and de Roon (2003) argue that using multiple futures contracts with different maturities and exploiting the term structure of convenience yields rather than futures prices form a more effective strategy than the naïve hedging strategy. In a similar context, Ripple and Moosa (2007) argue that hedging is more effective with near-month futures contracts than with distant ones. Switzer and El-Khoury (2007) and Alizadeh et al. (2008) exploit the multivariate Generalized Autoregressive Conditional Heteroskedasticity (GARCH) modelling approach to compute time-varying hedge ratios and compare them with the alternative measures of hedge ratio with respect to hedging effectiveness. The former study reveals that the time-varying hedge ratios based on an asymmetric version of the bivariate GARCH model outperform not only the OLS-based and naïve constant hedge ratios but also the ratios based on the symmetric version of the bivariate GARCH model. On the other hand, the latter study allows structural changes in the GARCH process through modelling it with the Markov regime switching approach. In this way, Alizadeh et al. (2008) improve the estimates of the GARCH-based hedge ratios and show that they are more effective than the alternative hedge ratios found in the literature. Another nonlinear approach, based on the spanning polynomial projection, is proposed by Chen and Liu (2008) with an objective to enhance the optimal hedging methods used in the literature. This nonlinear approach is argued to be worthwhile empirically for the crude oil risk hedgers, especially when transaction costs for hedging are low.

Recent studies on hedging effectiveness in crude oil markets appear to focus on rather complex hedging strategies. Bertus et. al (2009) investigate the cross hedging effectiveness in the jet fuel market using crude oil futures contracts, based on the rationale that the market for jet fuel is not liquid enough to support futures contracts. In the study, among models representing six different hedge strategies, the two-factor cross hedging models that allow for stochastic, mean-reverting convenience yield and a mean-reverting spread outperform competing simpler models, especially over longer horizons. Similarly, Yun and Kim (2010) examine the hedging effectiveness of a multiple-risk hedging strategy considering the intercorrelation between the crude oil prices and the exchange rate of the Korean won against the US dollar. Their findings imply a greater effectiveness for the complex hedging relative to the separate one. Moreover, this complex hedging strategy is found to improve hedging effectiveness when crude oil prices become more volatile and the exchange rate fluctuates less, and additionally, when the hedge period increases. Another recent study, in which Chang et. al (2010) compare hedging effectiveness with respect to eight different hedging models, demonstrates that bullish- versus bearish-market structure matter for the hedging performances of the energy futures because investors switch their hedging strategies as the market type changes. Accordingly, the hedging performance is found to be significantly better in the bull market for both crude oil and gasoline than in the bear market.

## 4. The Methodology of the Empirical Analysis

As a preliminary step in the hedging effectiveness analyses, speculative market efficiency is tested following Fama's (1984) two-regression approach. The relevant regressions to test whether or not the future-spot price differential (or the basis) has predictive power on future changes in spot prices and on the risk premium are given below respectively:

$$(S_{t+1} - S_t) = \alpha_1 + \beta_1(F_t - S_t) + \varepsilon_{1,t+1} \tag{1}$$
$$(F_t - S_{t+1}) = \alpha_2 + \beta_2(F_t - S_t) + \varepsilon_{2,t+1} \tag{2}$$

Note that these two regressions are complementary in the sense that $\alpha_1+\alpha_2=0$, $\beta_1+\beta_2=1$, $\varepsilon_{1,t+1}+\varepsilon_{2,t+1}=0$ and therefore, include some identical information. If the estimated $\beta_1$ and $\beta_2$ are found statistically significant in the regression estimates of (1) and (2), it implies that the basis contains information about future changes in spot prices and about the risk premium. However, for the efficient-market hypothesis to be valid, parameter restrictions $\alpha_1=0$ and $\beta_1=1$ (therefore, $\alpha_2=0$ and $\beta_2=0$) should hold, indicating the presence of risk neutral and rational agents. On the other hand, a significant $\alpha$ and a significant deviation of $\beta$ from unity imply the presence of both a significant constant and time-varying risk premiums, respectively.

Hedge ratios have been measured with various methods in the literature as econometric estimation techniques evolved over time. A simple constant measure of the hedge ratio is the ordinary least squares (OLS) estimate, which is based on regressing the change in futures prices ($\Delta F_t$) on the change in spot prices ($\Delta S_t$). Hence, the slope coefficient of this regression estimate is nothing but the ratio of the conditional variances,

$$\hat{\beta} = \frac{Cov(\Delta S_t, \Delta F_t / \Omega_{t-1})}{Var(\Delta F_t / \Omega_{t-1})} \tag{3}$$

where $\Omega_{t-1}$ represents the set of past information available. This coefficient is assumed the optimal or *minimum variance* constant hedge ratio. However, when the conditional variance is time-dependent due to volatility-clusters characterizing the financial data, the hedge ratio given by (3) becomes time-varying. To estimate time-varying hedge ratios, the recent literature on finance suggests using a bivariate GARCH (q, p) system as suggested by Bollerslev et al. (1988):

$$y_t = \nu + u_t \tag{4}$$

$$\mathrm{vech}(H_t) = \gamma + \sum_{i=1}^{q} A_i\ \mathrm{vech}(u_{t-i}u'_{t-i}) + \sum_{j=1}^{p} B_j\ \mathrm{vech}(H_{t-j}) \tag{5}$$

where $y_t=(\Delta S_t, \Delta F_t)'$ is a 2-dimensional variable vector, $\nu=(\nu_1, \nu_2)'$ is a 2×1 intercept vector and $u_t=(u_{1t}, u_{2t})'$ is a white noise residual vector with non-singular 2×2 conditional covariance matrix $H_t$ and with a conditional distribution of the formul $u_t|\Omega_{t-1}\sim N(0, H_t)$. In the specification for the conditional covariance matrix given by (5), vech denotes the half-vectorization operator. Hence, $\mathrm{vech}(H_t)$ is a 3×1 vector of conditional covariances and $\gamma=(\gamma_1, \gamma_2, \gamma_3)'$ is a 3×1 vector of intercepts. When q=p=1 in the GARCH (1, 1) case, the 3×3 coefficient matrices $A_i$ and $B_j$ can be denoted by $A_1$ and $B_1$. In order to restrict $H_t$ to be positive semi-definite and to reduce the number of parameters to be estimated, coefficient matrices are re-specified with all but the first column of coefficients equal to zero, as suggested by Ding and Engle (2001), and conditional covariances of the bivariate GARCH (1, 1) model become

$$\mathbf{H}_{11,t} = \mathbf{\Gamma}^{11} + \mathbf{A}_1^{11}(u_{1,t-1})^2 + \mathbf{B}_1^{11}(\mathbf{H}_{11,t-1}) \tag{6}$$

$$H_{12,t} = \Gamma^{12} + A_1^{12}(u_{1,t-1}\ u_{2,t-1}) + B_1^{12}(H_{12,t-1}) \tag{7}$$

$$\mathbf{H}_{22,t} = \mathbf{\Gamma}^{22} + \mathbf{A}_1^{22}(u_{2,t-1})^2 + \mathbf{B}_1^{22}(\mathbf{H}_{22,t-1}) \tag{8}$$

$\Gamma^{mn}$, $A_1^{mn}$ and $B_1^{mn}$ are the transformed coefficients in the above covariance equations for m, n = 1, 2 and m ≤ n. The process will be stationary if and only if $A_1^1 + B_1^1 < 1$ and $A_1^2 + B_1^2 < 1$. Now that the conditional covariance between the changes in spot and futures prices are estimated by $H_{12,t}$ and the conditional variance of the changes in futures prices are estimated by $H_{22,t}$, the time-varying hedge ratio will be

$$\hat{\beta}_t = \frac{\hat{H}_{12,t}}{\hat{H}_{22,t}} \tag{9}$$

The two βs, that is, $\hat{\beta}$ from OLS and $\hat{\beta}_t$ from GARCH estimations, are used to construct two different portfolios, such as $(\Delta S_t - \hat{\beta}\Delta F_t)$ and $(\Delta S_t - \hat{\beta}_t \Delta F_t)$, in order to perform a comparison between their variances. The relative hedging effectiveness of a portfolio is quantified by calculating the percentage variance reduction achieved with the alternative portfolio as

$$HE = \frac{\sigma^2_{OLS} - \sigma^2_{GARCH}}{\sigma^2_{OLS}} \tag{10}$$

However, we argue that any numerical difference between two portfolio variances does not necessarily mean a statistically significant difference. Therefore, a well-known variance ratio test is suggested to test the significance of the difference in variances:

$$F = \frac{\hat{\sigma}^2_L}{\hat{\sigma}^2_S} \sim F(n-1, n-1) \tag{11}$$

Here $\hat{\sigma}^2_L$ and $\hat{\sigma}^2_S$ are the portfolio variances where the one with the larger (L) value is used in the numerator and the smaller (S) one in the denominator. n is the total number of observations. If the F-statistic is found greater than the corresponding critical value, the null hypothesis of 'equal variances' is rejected, implying that the portfolio with the small variance outperforms the other in terms of hedging effectiveness.

## 5. Estimation Results

### 5.1. Data and Univariate Analysis

The crude-oil spot market subject to this study is the West Texas Intermediate (WTI), of which the futures contracts are traded in the New York Mercantile Exchange (NYMEX). The sample period covers daily closing prices ranging between October 2001 and August 2006, a period which is characterised by a rising trend beginning with the recovery from the burst of the 'dot-com' bubble and a global increase in the energy demand. The extreme volatility observed in the study period is the result of both the production decisions of the OPEC and the geo-political and natural events influencing the world oil production. The weakness of the U.S. economy and the increases in non-

OPEC production put downward pressure on prices in 2001. However, due to the September 11 terrorist attacks crude oil prices fall drastically, in spite of the quota reductions of the OPEC by September 1. In response, both OPEC and non-OPEC oil producers cut their quota again in the beginning of 2002, which results in the desired effect, moving oil prices up. Afterwards, the non-OPEC members restore their production cuts but prices continue to rise and U.S. inventories reach a 20-year low later in the year. This oversupply does not constitute a problem since the unrest in Venezuela in 2003 causes Venezuelan production to fall sharply. OPEC increased quotas in response. However, while the Venezuelan production begins to recover, the burst of the Gulf War causes a long sharp upward swing in oil prices. Meanwhile, the recovery in the U.S. economy increases the demand and the Asian demand for crude oil grows at a rapid pace. The hurricanes of 2005 and the U.S. refinery problems contribute to higher prices. The rush of investors or speculators to buy more oil futures contracts in this increasing price environment adds to demand by driving a self-fulfilling prophecy. In this respect, higher daily volatility than the estimated annual volatility may be one of the attributes of an inefficient market, where daily price swings significantly create mispriced stocks.

The period of the study coincides with a market that starts to show contango characteristics. Thus, hedging effectiveness becomes more challenging in an environment where futures prices are expected to be higher than the spot price, that is, in a contango market.

The spot- and futures-price series in the study are obtained from the Energy Information Administration (EIA) and the Datastream database, respectively. The observations on NYMEX crude oil futures contracts are the daily nearby contract prices while the spot prices are WTI Cushion crude oil prices. To investigate the univariate characteristics of the time series in question (see in Figure 2), first, spot and futures prices ($S_t$ and $F_t$) are tested for the unit root (see in Table 1). Additionally, the integration characteristics of the changes in spot prices ($S_{t+1}$-$S_t$), the basis ($F_t$-$S_t$), and the premium ($F_t$-$S_{t+1}$) are investigated. In order to provide a confirmatory unit root analysis, the ADF (Dickey and Fuller, 1979) and the KPSS (Kwiatkowski, Phillips, Schmidt, and Shin, 1992) tests are used together, where the null hypothesis represents a unit root (non-stationarity) in the former and stationarity in the latter. According to both the ADF and KPSS test statistics, $S_t$ and $F_t$ series are stationary in first-differences; however, they are trend stationary in levels at 5 % significance level according to the ADF test. The series of changes in spot prices, the basis, and the premium are found to be stationary, except for the KPSS test result indicating non-stationarity for the basis. Since the plot of this series exhibits visually stationary characteristics despite significant outliers (not shown for brevity), non-stationarity is thought to be the outcome of the extreme values in the series.

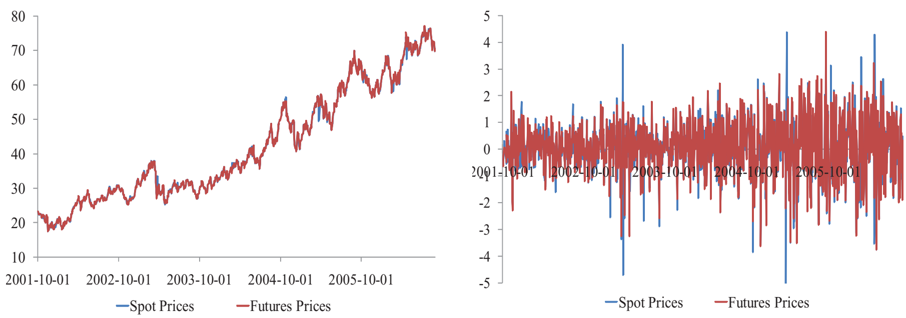


**Figure 2. Plots of spot and futures prices in levels and differences**

**Table 1. Unit root tests**

| | ADF | | KPSS | |
|---|---|---|---|---|
| | (I) | (II) | (I) | (II) |
| $S_t$ | -0.6302 | -3.7686** | 3.9585*** | 0.6692*** |
| $F_t$ | -0.5465 | -3.5877** | 3.9609*** | 0.6828*** |
| $\Delta S_t$ | -34.498*** | -34.489*** | 0.0570 | 0.0193 |
| $\Delta F_t$ | -33.264*** | | 0.0581 | 0.0199 |
| $S_{t+1}-S_t$ | -34.498*** | -34.489*** | 0.0570 | 0.0193 |
| $F_t-S_t$ | -20.525*** | -20.866*** | 1.5898*** | 0.2546*** |
| $F_t-S_{t+1}$ | -30.365*** | -30.429*** | 0.3911 | 0.1031 |

Testing equation includes (I) only an intercept or (II) an intercept and a trend. $S_t$ and $F_t$ are spot and futures crude oil prices. $\Delta$ is the first-difference operator. $S_{t+1}-S_t$ is the change in spot prices; $F_t-S_t$ and $F_t-S_{t+1}$ denote the basis and the premium, respectively. Asterisks ** and *** denote statistical significance at 5 % and 1 % levels, respectively.

The descriptive statistics given in Table 2 exhibit almost similar values for the means and the variances of the differenced $S_t$, $F_t$ and $S_{t+1}$-$S_t$ series. Skewness and kurtosis coefficients show that the distributions of the series are leftward skewed for $\Delta S_t$, $\Delta F_t$ and $S_{t+1}$-$S_t$ series, rightward skewed for $F_t$-$S_t$ and $F_t$-$S_{t+1}$ and leptokurtic for all series in question. Moreover, the normality hypothesis is strongly rejected for all of the series with respect to the Jarque-Bera test statistics. The residuals of the regression models employing series with leptokurtic distributions are usually dealt with using the autoregressive conditional heteroskedastic (ARCH) modelling approach.

**Table 2. Descriptive statistics**

| | $\Delta F_t$ | $\Delta S_t$ | $S_{t+1}$-$S_t$ | $F_t$-$S_t$ | $F_t$-$S_{t+1}$ |
|---|---|---|---|---|---|
| Mean | 0.0437 | 0.0439 | 0.0439 | 0.0124 | -0.0315 |
| Maximum | 4.3900 | 4.3600 | 4.3600 | 5.4500 | 6.6000 |
| Minimum | -3.7600 | -6.5200 | -6.5200 | -5.4500 | -4.7600 |
| St. Deviation | 0.9561 | 1.0714 | 1.0351 | 0.4525 | 1.0488 |
| Skewness | -0.2268 | -0.3998 | -0.3998 | 2.6285 | 0.6316 |
| Kurtosis | 4.3109 | 6.2938 | 6.2938 | 68.982 | 6.7947 |
| Jarque-Bera | 86.187*** | 514.60*** | 514.60*** | 196424*** | 716.44*** |
| Observations | 1075 | 1075 | 1075 | 1076 | 1075 |

Asterisks ** and *** denote statistical significance at 5 % and 1 % levels, respectively.

### 5.2. Estimations on Market Efficiency

In order to investigate the speculative market efficiency in the Fama sense, two regressions are run to test (i) whether the basis ($F_t$-$S_t$) contains information on future spot prices ($S_{t+1}$) and (ii) whether a time-varying risk premium exists. The corresponding estimates of the regressions are given in the first two columns of Table 3. Regression coefficients of the basis are statistically significant in both of the specifications and different from unity as well. The null hypothesis of 'no break points' is not rejected by the Andrews-Quandt test statistics, which justifies the stability of the regression coefficients. However, with respect to the distributional properties of $\Delta S_t$ and $\Delta F_t$ reported in Table 2, one may expect significant ARCH effects in the residuals of the estimated regressions. Accordingly, the Lagrange Multiplier (LM) tests given in Table 3 indicate the presence of such effects for the first- and higher-order cases. Therefore, a re-estimation of the Fama-type regressions will be necessary in order to eliminate any possible efficiency losses by modelling their conditional variances through an ARCH or Generalized ARCH (GARCH) process. The re-estimations and their ARCH-LM tests are given in the third and fourth columns of the table. The insignificant ARCH-LM statistics of the re-estimated regressions imply that ARCH effects are modelled successfully through the GARCH(1,1) process, resulting in also a lower value of the Akaike Information Criterion (AIC) as an indicator of the better fit. However, the normality assumption is still violated, most probably due to the outliers in the data series, as can be seen in the second plot of Figure 2.

**Table 3. Estimates of the Fama-type regressions**

| | $S_{t+1}-S_t$ | $F_t-S_{t+1}$ | $S_{t+1}-S_t$ | $F_t-S_{t+1}$ |
|---|---|---|---|---|
| | (I) | (II) | (III) | (IV) |
| *Mean equation* | | | | |
| Intercept | 0.0386<br>(0.0310) | -0.0386<br>(0.0310) | 0.0422*<br>(0.0251) | -0.0422*<br>(0.0239) |
| $F_t-S_t$ | 0.4305***<br>(0.0686) | 0.5695***<br>(0.0686) | 0.4088***<br>(0.0529) | 0.5913***<br>(0.0529) |
| *Variance equation* | | | | |
| Intercept | | | 0.0084***<br>(0.0022) | 0.0090***<br>(0.0043) |
| $\hat{u}^2_{t-1}$ | | | 0.0447***<br>(0.0066) | 0.0565***<br>(0.0139) |
| $\hat{h}_{t-1}$ | | | 0.9489***<br>(0.0059) | 0.9373***<br>(0.0143) |
| $R^2$ | 0.0354 | 0.0604 | 0.0353 | 0.0603 |
| F-stat. | 39.429*** | 68.983*** | 9.8007*** | 17.168*** |
| *Andrews-Quandt test* | | | | |
| (i) SupF | 1.4519 | | | |
| (ii) ExpF | 0.3058 | | | |
| (iii) AveF | 0.5889 | | | |
| ARCH-LM(1) | 19.170*** | | 0.8065 | |
| ARCH-LM(10) | 56.927*** | | 5.6379 | |
| ARCH-LM(15) | 65.711*** | | 8.7866 | |
| Jarque-Bera | 445.49*** | | 384.11*** | |
| AIC | 2.8736 | | 2.7213 | |
| $H_0$: $\alpha$=0 | 1.5452 | 1.5452 | 2.8331* | 2.8331* |
| $H_0$: $\beta$=1 | 68.983*** | 39.429*** | 124.70*** | 59.601*** |
| $H_0$: $\alpha$=$\beta$-1=0 | 35.006*** | 41.438*** | 140.15*** | 59.636*** |

Columns I and II include OLS estimates, while columns III and IV show the GARCH (1, 1)-conditional-variance estimates, additionally. GARCH estimations are maximum likelihood estimations based on the iterative Marquardt algorithm. Data is in daily frequency ranging from 10/01/2001-8/30/2006. F-stat represents the joint significance of the regression coefficients. SupF, ExpF and AveF are Andrew-Quandt statistics to test the null of 'no break points' with the asymptotic Hansen critical values. ARCH-LM denotes the Lagrange multiplier (LM) test for autoregressive conditional heteroskedasticity (ARCH) in the residuals, with the $\chi^2$ distribution. The Jarque-Bera statistic tests the null of whether the standardized residuals are normally distributed. Asterisks ** and *** denote statistical significance at 5 % and 1 % levels, respectively. The null hypotheses of $\alpha$=0 and $\beta$=1 are tested by a Wald statistic with the $\chi^2$ distribution.

The statistically significant regression coefficients in the mean equations in the third and fourth columns of Table 3 imply that the basis contains information about both the future spot prices and the risk premium. The significant constant terms indicate a constant risk premium while the significant slope coefficients, which are also significantly different from 1, indicate a time-varying risk premium. Hence, this outcome does not provide any support for the speculative market efficiency hypothesis. Futures prices are not unbiased predictors of the future spot prices. Then, determining the optimal hedging strategy essentially constitutes the primary concern of risk-averse investors.

A robustness check of the lack of unbiasedness is performed, following the approach used by Abosedra and Baghestani (2004). The predictive accuracy of futures prices are examined based on the 1-, 2-, 3-, 4-, 5-, 10-, 15-, 20- and 25-day ahead futures prices. The results of the corresponding unbiasedness tests are given in Table 4. Again, the findings show that the futures prices are unbiased at neither of the forecast horizons.

**Table 4. Unbiasedness of Forecasts**

| f-day ahead forecasts | $\alpha$ | $\beta$ | St. Err. | $\chi^2$ |
|---|---|---|---|---|
| f=1 | -0.0084 | 0.9609*** | 0.4497 | 16.9*** |
| | (0.0141) | (0.0111) | | |
| f=2 | -0.0076 | 0.9678*** | 0.4499 | 16.56*** |
| | (0.0141) | (0.0087) | | |
| f=3 | -0.0072 | 0.9740*** | 0.4506 | 14.49*** |
| | (0.0142) | (0.0074) | | |
| f=4 | -0.0066 | 0.9766**** | 0.4508 | 14.36*** |
| | (0.0143) | (0.0074) | | |
| f=5 | -0.0055 | 0.9762** | 0.4503 | 17.6*** |
| | (0.0144) | (0.0085) | | |
| f=10 | -0.0072 | 0.9899*** | 0.4539 | 5.53* |
| | (0.0148) | (0.0074) | | |
| f=15 | -0.0043 | 0.9905*** | 0.4542 | 6.69** |
| | (0.0145) | (0.0062) | | |
| f=20 | -0.0014 | 0.9900*** | 0.4547 | 9.03** |
| | (0.0142) | (0.0056) | | |
| f=25 | 0.0010 | 0.9900*** | 0.4555 | 10.45** |
| | (0.0139) | (0.0056) | | |

Note: The unbiasedness of forecasts is examined by the OLS estimate of the regression

$$S_{t+f} - S_{t-1} = \alpha + \beta (P_{t+f} - S_{t-1}) + v_{t+f} \qquad f=1, 2, 3, 4, 5, 10, 15, 20, 25$$

$S_{t+f}$ is the logarithm of the crude oil spot price in day $t + f$ while $P_{t+f}$ is the forecast of $S_{t+f}$ made in day $t$ with $f$ defined as the forecast horizon. Data is in daily frequency ranging from 10/01/2001-8/30/2006; effective sample changes for different forecast horizon. St. Err. is the standard error of the estimate. $\chi^2$ is the chi-square statistic for testing the joint null hypothesis of unbiasedness $a_0$=0 and $a_1$=1. The numbers in parentheses are standard errors. Asterisks ** and *** denote statistical significance at 5 % and 1 % levels, respectively.



### 5.3. Estimations on Hedge Ratios and Hedging Effectiveness

In this study, the two measures of the hedge-ratio estimated are (i) the constant minimum variance hedge ratio, and (ii) the time-varying hedge ratio. The first is simply the slope coefficient of the regression of $\Delta F_t$ on $\Delta S_t$, while the second is a varying coefficient based on the conditional variance estimates of the bivariate GARCH model. The relevant regression estimate of the former is given in Table 5, where the estimated constant hedge ratio is 0.9395. The Quandt-Andrews statistics reported in the lower panel of Table 5 allow testing for whether or not one or more unknown structural breakpoints exist in $\hat{\beta}$ within trimmed data in the sample.

**Table 5. OLS regression estimate**

$$\Delta\hat{S}_t = \underset{(0.0146)}{0.0029} + \underset{(0.0149)}{0.9395}\Delta F_t \qquad R^2 = 0.7876$$

Quandt-Andrews test:

$$\text{(i) SupF} = 6.660;\ \text{(ii) ExpF} = 1.494;\ \text{(iii) AveF} = 2.019$$

Ljung-Box test (squared standardized residuals):

$$Q(1) = 59.789^{***};\ Q(5) = 96.026^{***};\ Q(15) = 99.103^{***}$$

ARCH-LM test:

$$LM(1) = 64.449^{***};\ LM(5) = 88.285^{***};\ LM(15) = 87.838^{***}$$

Note: Figures in parentheses in the regression are standard errors. Data is in daily frequency ranging from 10/04/2001-8/31/2006. SupF, ExpF and AveF are Andrew-Quandt statistics to test the parameter stability in the OLS regression. Ljung-Box (Q) is the Lagrange multiplier statistic to test for serial correlation in the residuals. ARCH-LM indicates the test for autoregressive conditional heteroskedasticity (ARCH) in the residuals. Asterisks *** denote statistical significance at 1 % level.

The null hypothesis of 'no break points' is not rejected, which justifies the stability of the parameter. However, although such an estimate is accepted and used as a measure of the hedge-ratio conventionally, it should be noted that the residuals of the regression in question exhibit significant ARCH effects for first- and higher-order cases, reflected in the significant statistics computed both by the Ljung-Box tests performed with squared standardized residuals and by the ARCH-LM tests. This implies that the OLS hedge ratio is time dependent due to the time-varying nature of the conditional variances and the covariance of $\Delta S_t$ and $\Delta F_t$.

$\Delta S_t$ and $\Delta F_t$ are modelled and estimated in a bivariate GARCH(1,1) setting with the VECH specification of the conditional variance to obtain a time-varying estimate of the hedge ratio. Estimations are based on the BHHH iterative optimization algorithm (Berndt, Hall, Hall and Hausman, 1974) and on the assumption that errors have the Student's t-type conditional distribution which has more weights in the tails. Moreover, the coefficient matrices of the conditional variance are restricted with all but the first columns of coefficients equal to zero (Ding and Engle, 2001) so that the number of parameters estimated is reduced and the positive semi-definiteness of the conditional covariance matrix is guaranteed. The resulting estimates given in Table 6 show that the bivariate GARCH(1,1) specification fits well to/with the conditional variances and co-variance of $\Delta S_t$ and $\Delta F_t$. The transformed coefficients are statistically significant and satisfy the stationarity condition of the GARCH process. Moreover, the Ljung-Box test statistics based on the squared standardized residuals of the model give no evidence on ARCH effects left in the residuals. Hence, the time series estimates of the conditional covariance between $\Delta S_t$ and $\Delta F_t$ and the variance of $\Delta F_t$ can be used to compute time-varying hedge ratios.

Figure 3 shows the plot of the time-varying ratios together with the constant ratio computed from the OLS estimation over the whole sample. It shows that time-varying hedge ratios move slightly above the constant hedge ratio in most of the sample period, except for the outliers. However, by the end of March 2005, where the extreme ratios are observed, the hedge ratios begin to move closer to the constant hedge ratio with more volatile values. Data series on spot and futures prices reveal that this period is characterized by relatively higher basis risks compared to the previous sample period.

Finally, the relative effectiveness of the constant and time-varying hedge ratios are analyzed by comparing the variances of the two portfolios computed with these ratios. In the literature, hedging effectiveness is measured by the percentage reduction in the variance of a portfolio relative to that of another. However, the numerical difference in the variances does not necessarily mean a statistical difference. Therefore, an inference on effectiveness should be justified by a statistical test of variance-equality. Table 7 shows the portfolio variances calculated by the constant and time-varying hedge ratios.



**Table 6. Bivariate GARCH(1, 1) estimate**

| | GARCH(1, 1) Model | | |
|---|---|---|---|
| *Mean equation* | | | |
| Intercept | 0.0892*** | | |
| | (0.0240) | | |
| Intercept | 0.0852*** | | |
| | (0.0239) | | |
| *Variance equation* | | | |
| Intercept | 1.0661*** | | |
| | (0.2295) | | |
| Intercept | 0.9407*** | | |
| | (0.2077) | | |
| Intercept | 0.8300*** | | |
| | (0.1886) | | |
| $\left(\hat{u}_{1,t-1}\right)^2$ | 0.2998*** | | |
| | (0.0672) | | |
| $\left(\hat{u}_{1,t-1}\hat{u}_{2,t-1}\right)$ | 0.2715*** | | |
| | (0.0608) | | |
| $\left(\hat{u}_{2,t-1}\right)^2$ | 0.2459*** | | |
| | (0.0562) | | |
| $\hat{H}_{11,t-1}$ | 0.3073*** | | |
| | (0.0557) | | |
| $\hat{H}_{12,t-1}$ | 0.3722*** | | |
| | (0.0554) | | |
| $\hat{H}_{22,t-1}$ | 0.4507*** | | |
| | (0.0527) | | |
| *Ljung-Box (Q)* | $Q^2(1)$ | $Q^2(5)$ | $Q^2(15)$ |
| $\Delta S_t$ equation | 0.1022 | 3.5787 | 4.5255 |
| $\Delta F_t$ equation | 0.0204 | 2.2392 | 2.8049 |

Note: Dependent variables of the bivariate system are $\Delta S_t$ and $\Delta F_t$. The bivariate GARCH(1,1) estimation is based on the VECH specification of the conditional variance and on the Berndt-Hall-Hall-Hausman (BHHH) iterative optimization algorithm. $H_{11,t-1}$, $H_{22,t-1}$ and $H_{12,t-1}$ are conditional variances and covariances of $\Delta S_t$ and $\Delta F_t$ Errors are assumed to have the Student's t-type conditional distribution. Ljung-Box (Q) is the Lagrange multiplier statistic to test for serial correlation in the residuals.Asterisks *** denote statistical significance at 1 % level.

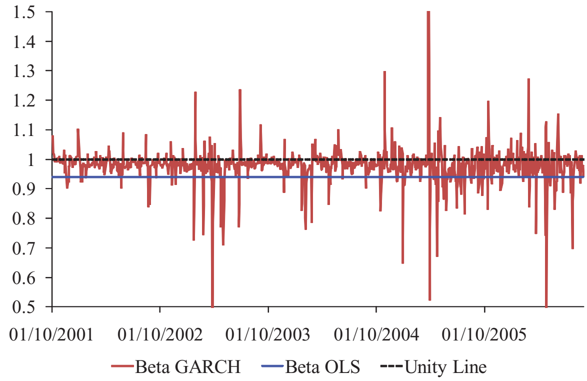


**Figure 3. Plots of the constant and time-varying hedge ratios**

**Table 7. Hedging effectiveness**

| | Constant | Time-varying |
|---|---|---|
| Portfolio variance | 0.2277 | 0.2462 |
| Percentage reduction in variance | -7.5193 | |
| Variance ratio test (F-test) | 1.0813 | |

Note: *Portfolio Variance* is the variance of the residuals $[(S_{t+1}-S_t)-\beta_i(F_{t+1}-F_t)]$, in which $\beta_i$'s are the simple OLS estimate $\beta$=0.939513 for all *i* for the constant case, whereas $\beta_i$'s are the GARCH-based hedge ratios for the time-varying case. *Percentage reduction in variance* denotes the rate of change between the computed constant and time-varying hedge ratios. *Variance ratio test (F-test)* statistic is computed as the ratio of the time-varying portfolio variance to the constant one, and the corresponding probability of non-rejection is found to be 0.196 for F(1074, 1073) critical value.

The variance reduction of about -8 % indicates that time-varying hedge ratios do not outperform the constant hedge ratio in terms of the hedging effectiveness. However, this numerical difference is not found statistically significant with respect to the variance ratio test statistic computed. Thus, it can be concluded that constant and time-varying hedge ratios provide equal hedging effectiveness.

## 6. Conclusion

The objective of this study is to analyze the speculative efficiency and the hedging effectiveness of the NYMEX crude oil futures market. According to our empirical findings, the NYMEX futures market is not an efficient market in the Fama sense between October 2001 and August 2006. In other words, NYMEX futures prices cannot be accepted as unbiased estimators of the WTI spot prices in the period in question. In order to analyze the hedging effectiveness of the NYMEX futures contracts, the time-varying hedge ratios are estimated by using the bivariate GARCH approach, beside the OLS-based constant hedge ratio. These time-varying ratios are found to be slightly above the constant one in most of the sample period, indicating a better hedge. However, when the relative effectiveness of the two hedge ratios is compared with respect to their corresponding portfolio variances, they are found equally effective in statistical terms. The statistically insignificant difference between the portfolio variances based on the constant and time-varying hedge ratios may be a reflection of the static hedging strategy that characterizes investments in crude oil futures markets.

Our inefficiency finding implies that no unpredictable variation exists between the spot and futures prices in the upward trending crude oil market in the October 2001-August 2006 period. Both the market inefficiency evidence and the high hedge ratios estimated may be an indication of the hedging intensive structure of the market. However, such a market may as well be favourable to the risk neutral speculators taking positions to exploit the above normal profit opportunities. Thus, risk neutral speculators could make consistent profits or effective hedging portfolios on long or short positions through time in the crude oil market for the observed period. Accordingly, it may be argued that it is relevant to replace the most commonly used hedging theory, which emphasizes the importance of the utilization of the portfolio approach, with the traditional hedging theories. That is, taking an equal opposite position and being interested in relative prices rather than absolute prices appear to be the main concern of the hedgers in the crude oil futures market for the observed period. This implies that the period of interest, which shows a strong upward trend with extreme volatility, is characterized mainly by the speculative hedging behaviour. Hence, rather than the basis risk, market depth risk and the commission risk, which are beyond the scope of this study, appear to be significant possible issues for future research.



**References**

Abosedra, S. and Baghestani, H. (2004) “On the predictive accuracy of crude oil futures prices,” *Energy Policy*, 32(12), 1389–1393. http://dx.doi.org/10.1016/S0301-4215(03)00104-6

Aksoy, G. and Olgun, O. (2009) “ Optimal Hedge oranı tahminlemesi üzerine ampirik bir çalışma: VOB örneği,” *İktisat, İşletme ve Finans*, Cilt 24, Sayı 274, 33-53.

Alizadeh, A. H., Nomikos, N. K. And P. K. Pouliasis (2008) “A Markov regime switching approach for hedging energy commodities,” *Journal of Banking & Finance*, 32(9), 1970–1983. http://dx.doi.org/10.1016/j.jbankfin.2007.12.020

Alvarez-Ramirez, J., Alvarez, J. and E. Rodriguez (2008) "Short-term predictability of crude oil markets: A detrended fluctuation analysis approach," *Energy Economics*, 30(5), 2645–2656. http://dx.doi.org/10.1016/j.eneco.2008.05.006

Bekiros, S. D. and Diks, C. G. H. (2008) "The relationship between crude oil spot and futures prices: Cointegration, linear and nonlinear causality," *Energy Economics*, 30(5), 2673–2685. http://dx.doi.org/10.1016/j.eneco.2008.03.006

Berndt, E. K., Hall, B. H., Hall, R. E. and J. A. Hausman (1974) "Estimation and inference in nonlinear structural models," *Annals of Economic and Social Measurement,* 3(4), 653–665.

Bertus, M., Godbey, J. and J. E. Hilliard (2009) "Minimum variance cross hedging under mean-reverting spreads, stochastic convenience yields, and jumps: Application to the airline industry," *The Journal of Futures Markets*, 29(8), 736–756. http://dx.doi.org/10.1002/fut.20379.

Bollerslev, T., Engle, R.F. and J.M. Wooldridge (1988) "A capital asset pricing model with time-varying covariance," *Journal of Political Economy*, 96(1), 116-131. http://dx.doi.org/10.1086/261527

Bopp, A. E. and Lady, G. M. (1991) "A comparison of petroleum futures versus spot prices as predictors of prices in the future," *Energy Economics*, 13(4), 274-282. http://dx.doi.org/10.1016/0140-9883(91)90007-M

Bopp, A. E. and Sitzer, S. (1987) "Are petroleum futures prices good predictors of cash value?," *The Journal of Futures Markets*", 7(6), 705-719. http://dx.doi.org/10.1002/fut.3990070609

Chang, C. Y., Lai, J. Y. and I.Y. Chuang (2010) "Futures hedging effectiveness under the segmentation of bear/bull energy markets," Energy Economics, 32(2), 442-449. http://dx.doi.org/10.1016/j.eneco.2009.09.012

Chen, A. S. and Liu, Y. Z. (2008) "Enhancing hedging performance with the spanning polynomial projection," *Quantitative Finance*, 8(6), 605-617. http://dx.doi.org/10.1080/14697680701570101

Chen, K. C., Sears, R. S. and D. N. Tzang (1987) "Oil prices and energy futures," *The Joumal of Futures Markets*, 7(5), 501-518. http://dx.doi.org/10.1002/fut.3990070505

Chng, M. T. (2009) "Economic linkages across commodity futures: Hedging and trading implications," *Journal of Banking and Finance*, 33(5), 958–970. http://dx.doi.org/10.1016/j.jbankfin.2008.10.006

Coppola, A. (2007) "Forecasting oil price movements: Exploiting the information in the futures market," *The Journal of Futures Markets*, 28(1), 34–56. http://dx.doi.org/10.1002/fut.20277

Crowder, W. J. and Hamed, A. (1993) "A cointegration test for oil futures market efficiency," *The Journal of Futures Markets*, 13(8), 933-941.http://dx.doi.org/10.1002/fut.3990130810

Deaves R. and Krinsky, I. (1992) "Risk premiums and efficiency in the market for crude oil futures," *The Energy Journal*, 13(2), 93-117.

Dickey, D. and W. A. Fuller (1979), "Distribution of the estimates for autoregressive time series with a unit root," *Journal of the American Statistical Association*, 74(366), 427-431. http://dx.doi.org/10.2307/2286348

Ding, Z. and Engle, R. (2001) "Large scale conditional covariance matrix modeling estimation and testing," *Academia Economic Papers*, 29(2), 157-184.

Domanski, D. and Heath, A. (2007) "Financial investors and commodity markets," *BIS Quarterly Review*, March 2007, 53-67.

Ederington, L. H. (1979) "The hedging performance of the new futures markets," *The Journal of Finance*, 34(1), 157-170. http://dx.doi.org/10.2307/2327150

Fama, E. F. (1984) "Forward and spot exchange rates," *Journal of Monetary Economics*, 14(3), 319-338. http://dx.doi.org/10.1016/0304-3932(84)90046-1

Fujihra, R. A. and Mougouè, M. (1997) "Linear dependence, nonlinear dependence and petroleum futures market efficiency," *The Journal of Futures Markets*, 17(1), 75–99. http://dx.doi.org/10.1002/(SICI)1096-9934(199702)17:1<75::AID-FUT4>3.0.CO;2-D

Gülen, S. G. (1998) "Efficiency in the crude oil futures market," *Journal of Energy Finance & Development*, 3(1), 13-21. http://dx.doi.org/10.1016/S1085-7443(99)80065-9

Hammoudeh, S. and Yuan, Y. (2008) "Metal volatility in presence of oil and interest rate shocks," *Energy Economics*, 30(2), 606–620. http://dx.doi.org/10.1016/j.eneco.2007.09.004

Kasman, A. and Kasman, S. (2008) "The impact of futures trading on volatility of the underlying asset in the Turkish stock market," *Physica A*, 387(12), 2837–2845.

Kwiatkowski, D., Phillips, P. C. B., Schmidt, P. and Y. Shin (1992) "Testing the null hypothesis of stationarity against the alternative of a unit root: How sure are we that economic time series have a unit root?," *Journal of Econometrics*, 54(1-3), 159-178. http://dx.doi.org/10.1016/0304-4076(92)90104-Y

Lee, H. C., Chien, C. Y. and T. H. Liao (2009) "Determination of stock closing prices and hedging performance with stock indices futures," *Accounting and Finance*, 49(4), 827–847.http://dx.doi.org/10.1111/j.1467-629X.2009.00309.x

Li, X. and Zhang, B. (2009) "Price discovery for copper futures in informationally linked markets," *Applied Economics Letters*, 16(15), 1555-1558. http://dx.doi.org/10.1080/13504850701578801

Lien, D. and Schaffer, D. R. (2002) "Multiperiod strip hedging of forward commitments," *Review of Quantitative Finance and Accounting*, 18(4), 345–358.http://dx.doi.org/10.1023/A:1015401719441

Lindahl, M. (1989) "Measuring hedging effectiveness with $R^2$*: A n*ote," *The Joumal of Futures Markets*, 9(5), 469-475. http://dx.doi.org/10.1002/fut.3990090509

Moosa, I.A. and Al-Loughani N.E. (1994) "Unbiasedness and time varying risk premia in the crude oil futures market," *Energy Economics*, 16(2), 99–105. http://dx.doi.org/10.1016/0140-9883(94)90003-5

Noussair, C. and Tucker, S. (2006) " Futures markets and bubble formation in experimental asset markets," *Pacific Economic Review*, 11(2), 167-184. http://dx.doi.org/10.1111/j.1468-0106.2006.00308.x

OPEC (2008) *World Economic Outlook 2008*, Organization of the Petroleum Exporting Countries, Vienna, Austria.

Peroni, E. and McNown, R. (1998) "Noninformative and informative tests of efficiency in three energy futures markets," *The Journal of Futures Markets*, 18(8), 939-964. http://dx.doi.org/10.1002/(SICI)1096-9934(199812)18:8<939::AID-FUT4>3.0.CO;2-4

Quan, J. (1992) "Two-step testing procedure for price discovery role of futures prices," *The Journal of Futures Markets*, 12(2), 139-149. http://dx.doi.org/10.1002/fut.3990120203

Ripple, R. D. and Moosa, I. A. (2007) "Hedging effectiveness and futures contract maturity: The case of NYMEX crude oil futures," *Applied Financial Economics*, 17(9), 683–689. http://dx.doi.org/10.1080/09603100600722177

Roll, R., Schwartz, E. and A. Subrahmanyam (2007) "Liquidity and the Law of One Price: The Case of the Futures-Cash Basis," *The Journal of Finance*, 62(5), 2201-2234. http://dx.doi.org/10.1111/j.1540-6261.2007.01273.x

Shambora, W. E. and Rossiter. R., (2007) "Are there exploitable inefficiencies in the futures market for oil?," *Energy Economics*, 29(1), 18–27. http://dx.doi.org/10.1016/j.eneco.2005.09.004

Switzer, L. N. and El-Khoury, M. (2007) "Extreme volatility, speculative efficiency, and the hedging effectiveness of the oil futures markets," *The Journal of Futures Markets*, 27(1), 61–84. http://dx.doi.org/10.1002/fut.20235

Veld-Merkoulova, Y. V. and de Roon, F. A. (2003) "Hedging long-term commodity risk," *The Journal of Futures Markets*, 23(2), 109–133. http://dx.doi.org/10.1002/fut.10060

Yun, W. C. and Kim, H. J. (2010) "Hedging strategy for crude oil trading and the factors influencing hedging effectiveness," *Energy Policy*, 38(5), 2404-2408. http://dx.doi.org/10.1016/j.enpol.2009.12.032