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Using implied volatility on options to measure the relation between asset returns and variability

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Abstract

Prior research has documented that volatility in financial asset markets is most directly related to trading rather than calendar days, and that there is an inverse asymmetric relation between volatility and returns in both stocks and long-term bonds. We examine these relations in 37 futures options markets representing a wide variety of asset types. Using futures prices and implied volatilities from this extensive array of markets, we confirm that in all of them, save one, market volatility is more directly related to trading days. However, the nature of the association between implied volatility and underlying asset returns varies greatly across asset categories and across exchanges. Thus, we show that findings from equity markets apparently are not generalizable to other asset classes. © 2001 Elsevier Science B.V. All rights reserved.

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1. Introduction

In a recent study, Fleming et al. (1995) found a strong temporal inverse relation between the CBOE Market Volatility Index and stock market returns, whereby negative stock market returns are associated with increases in market volatility and positive market returns result in decreased volatility. Their study further demonstrated that this relation is asymmetric, with negative stock returns being associated with larger changes in volatility than positive returns. We examine this issue using data from 37 options on futures contracts. Previous research has generally focused on equity-contracts. We examine a large variety of other asset categories, to see if results from equity markets can be generalized more broadly. In addition, following Bessembinder and Seguin (1993), we control for information inflows using trading volume, and for market depth using open interest.

This study builds on prior research in several ways. First, we use data from options on futures contracts. Since the options and their underlying futures contracts trade on the same floor, traders of both assets should receive information at the same time. Information asymmetries between the markets could influence the results so the simultaneous receipt of information should reduce or eliminate the asymmetries. In addition, transaction costs are considerably lower in these markets than in equity markets (Fleming et al., 1996) allowing for faster and easier arbitrage. Option valuation models assume instantaneous and costless arbitrage so these markets provide the ability to test the relation between options and underlying assets in a setting that is closer to the theoretical ideal than is the case with options on equity markets. Since the options trade on the same floor as their underlying futures and trading ends at the same time each day, we lessen the non-synchronous data problems that occur when there are separate markets and closing times (Harvey and Whaley, 1992). If the last trades on the option and futures occur at different times in a day, the simultaneity issue is not entirely eliminated, but since pit committees attempt to set a settlement price that reflects the fair value in both the futures and the futures option markets at the same time, the non-synchronous problem is somewhat alleviated.

Most significantly, this study adds to the prior literature by studying 37 separate futures options markets, including interest rate, currency, energy, grain, livestock, and metal futures options, from eight different exchanges, in addition to the S&P 500 index. The options we examine trade on the Chicago Board of Trade (CBOT), Chicago Mercantile Exchange (CME), Coffee, Sugar, and Cocoa Exchange (CSCE), Commodity Exchange (COMEX), New York Cotton Exchange (NYCE), New York Mercantile Exchange (NYME) London Financial Futures Exchange (LIFFE), and Marché à Terme International de France (MATIF). Since we use different categories of contracts from different exchanges, we can examine whether the results of our tests differ according to their category or exchange, and whether earlier findings for equity markets hold for other asset classes as well. This issue is important for two reasons. First, these other markets have extensive trading activity and are important in their own right: in terms of total contract value, open interest in Treasury Bond and Eurodollar futures easily rivals that in the S&P 500 index, and open interest in the Corn, Soybean, Gold, Crude Oil, 5- and 10-year Treasury Note, and some of the currency futures all are in excess of three billion dollars. The other reason for examining the relation between volatility and underlying asset returns in these markets is to determine whether the linkages in the equity options markets are similar to those in other markets. This has important implications for the ultimate cause and generalizability of these linkages. If the asymmetric volatility versus underlying asset return relation turns out to be unique to equities, we must question if the relation is real, reflecting fundamental economic forces at work in the equity market, or if the relation occurred purely by chance in the time interval examined by Fleming et al., (1995) and is unlikely to be repeated in the future.

Our results suggest that on S&P 500 Index futures options, positive returns on the underlying futures result in decreases in implied volatility, while negative returns on the underlying futures produce increases in the implies. In terms of absolute magnitude, implied volatility changes are larger for negative futures contract price changes than for positive price changes. These results are consistent with prior findings of Fleming et al. (1995) on cash-index equity options and with the findings of Simon (1997) on Treasury Bond futures options. They indicate that the implied volatility versus underlying asset return relation holds in the futures options market, when we control for trading volume and open interest, and when the sample period is somewhat different. However, we find that the nature of the volatility–return relation for equities is unique in many ways. In general, these results do not hold for other non-equity categories of contracts, although there are some notable similarities between the S&P 500 and long-term interest rate futures.

2. Implied volatility and underlying asset returns

2.1. General

The relation between volatility and asset returns has been studied numerous times, generally using stock market data. Fama (1965) demonstrated that large stock price changes tend to be followed by large volatility. Black (1976) and Christie (1982) found a negative correlation between stock returns and changes in the ex-post future volatility. French et al. (1987) showed that the ex-ante risk premium on common stock is positively related to expected stock market volatility. Schwert (1989) found that large increases in expected volatility occur in response to negative stock market returns, while positive stock returns lead to moderate decreases in expected volatility.

Option valuation requires an estimate for volatility. The best estimation technique for volatility has been the subject of considerable debate. Prior research has suggested both historical volatility (Canina and Figlewski, 1993; Hull and White, 1987; Scott, 1987; Wiggins, 1987) and implied volatility in which volatility is observed by substituting the option price into an option pricing model (Lamoureux and Lastrapes, 1993; Day and Lewis, 1992; Schmalensee and Trippi, 1978; Chiras and Manaster, 1978; Latané and Rendleman, 1976) as alternate methods for estimating option volatility. Day and Lewis (1994) and Jorion (1995) show that for futures options, which are used in this study, implied volatility dominates other methods for predicting realized volatility. Thus, we focus on implied volatility and its relation to changes in the underlying asset prices.

The use of implied volatility assumes that the market efficiently and correctly prices options. If implied volatility is an unbiased estimate of the underlying asset's future volatility, and if volatility in the immediate past and future volatility are related, then underlying asset returns should have an impact on implied volatility measures in an efficient market.

Fleming et al. (1995) found an asymmetric reaction of the CBOE Market Volatility Index and contemporaneous stock market returns. They hypothesized that a decrease in volatility would be accompanied by a stock market increase while an increase in volatility would be accompanied by a stock market decrease. While they confirm this expectation in their study, the change in volatility accompanying negative stock returns was more than twice the size of those accompanying positive stock returns. Their conclusion is that negative stock price changes have a greater impact on volatility. Our goal is to shed more light on the nature of this relation, and whether it is consistent across different types of asset markets.

2.2. Moderating influences to price and variability

There is a large body of literature that documents a positive contemporaneous correlation between price volatility and trading volume. [1] This relation exists in a wide range of financial markets. Bessembinder and Seguin (1993), for example, find that futures price volatility is positively related to volume and that unexpected volume changes have a bigger impact on volatility than when the market expects the change.

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[1] See Karpoff (1987) for a survey of this literature.

We consider volume changes to be a proxy for inflow of information, which can be called the mixture of distributions hypothesis. In one form of this hypothesis the amount of information that arrives in a market is a mixing variable (Clark, 1973; Tauchen and Pitts, 1983). Since price volatility is considered to be an increasing function of the number of information arrivals, and trading volume is also an increasing function of price changes, price volatility and trading volume are expected to have a positive correlation. Harris (1986) documented results consistent with the predictions of this hypothesis.

An alternate explanation appeared in Copeland (1976, 1977). It is a "sequential arrival of information" model used to explain the positive correlation between price volatility and trading volume. In his model, information is disseminated to a single trader at a time, and price changes occur sequentially as it spreads. This information will result in a clustering of price changes and trading volume.

Regardless of the cause of the positive correlation between trading volume and volatility, the fact that these measures are related has important implications for estimating the relation between underlying asset returns and volatility. It is important to examine whether the return–volatility relation is just an artifact of the links between trading volume and volatility, and trading volume and return, or if it is a separate and distinct effect. Thus, we use trading volume as a moderating (i.e. control) variable in our tests examining linkages between implied volatility and underlying asset returns.

Another factor that may influence the relation between asset returns and volatility is market depth. Market depth can be considered to be the volume of order flows required to move the asset's price by one minimum-price unit (one tick) allowed by that market's rules (Kyle, 1985). Bessembinder and Seguin (1993) use open interest as a proxy for market depth. They argue that open interest is an effective proxy because open interest at the close largely represents hedging activity, as opposed to the positions of speculators; many of the latter are day traders who do not hold overnight positions. They argue that using open interest with volume data "may provide insights into the price effects of market activity generated by informed versus uninformed traders or hedgers versus speculators" (p. 25). In addition, open interest may be a proxy for the willingness of current traders to accept risk. Bessembinder and Seguin (1993) state "Willingness is in part determined by traders' risk aversion while ability is partly determined by existing wealth constraints. If these and other determinants of depth do not change quickly, then a variable constructed from lagged open interest should contain information on current depth" (p. 25). Whether open interest serves as a measure of hedging activity or a measure of investor willingness to accept risk, we use it as an additional control variable in our tests.

3. Data

The data in this study were obtained from Knight–Ridder. [2] They consist of information on 37 futures options traded on eight exchanges and include financial, agricultural, industrial, and equity index futures options. The Knight–Ridder data contain daily time series of the implied volatility of futures options as well as data on the underlying futures such as daily open, high, low, and settlement prices, volume and open interest. Knight–Ridder calculates implied volatility using: (1) the Black (1976) option valuation formula; (2) the settlement prices of the futures and futures option contract; and (3) the risk-free rate obtained from the T-bill futures contract. The choices of the model and the proxy for the risk-free rate introduce error, however. The futures options used in this study are American-style, and the Black model values European-style options. This means that the Knight–Ridder implied volatility estimates are upward biased. Also, the T-bill futures provide an estimate of a forward interest rate while the option formula requires a spot rate. This, too, introduces error. [3] For the following reasons we believe that our results are not significantly affected by these errors.

The implied volatilities used in this study are obtained from the nearest-to-the-money options. Knight–Ridder calculates implied volatility on the two calls and the two puts with the strike prices nearest to the underlying futures price, and reports an unweighted average of these four implied volatilities. Whaley (1986, p. 134) and Barone-Adesi and Whaley (1987, p. 314) show that, for at-the-money options, on futures the difference between theoretical option, prices obtained from the Black (1976) model for European options and an alternative American option valuation model (such as the Barone-Adesi and Whaley, 1987) is typically less than 1% of the option's price. As the underlying futures have a zero cost of carry, the early exercise on an American futures option becomes valuable only if the option is in-the-money by a large amount (Barone-Adesi and Whaley, 1987). As Knight–Ridder uses the two nearest-to-the-money calls and puts, any measurement error that is in-

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[2] The data were purchased from Knight–Ridder for a nominal fee. Knight–Ridder Financial was acquired in 1997 by Bridge, a US financial information provider. The futures and futures options data that are used in this study are now available from Bridge (for more information: www.crbindex.com).

[3] These choices are, unfortunately, beyond our control, as they are imposed on us by Knight–Ridder.

troduced by using Black's (1976) model would be negligible. [4] In addition, because our regressions use the daily difference in implied volatility as the dependent variable, any systematic bias in implied volatility that arises because Knight–Ridder uses Black's (1976) model should wash out when taking the differences.

Another concern is that Knight–Ridder uses the risk-free interest rate obtained from T-bill futures rather than the risk-free interest rate obtained from T-bills themselves. The Black (1976) model calls for a spot rather than a forward interest rate. The term-premium included in the T-bill futures price could potentially bias the implied volatilities calculated by Knight–Ridder. During the period of the study, the risk-free rate (bank discount rate) obtained from spot T-bills is approximately 16 basis points lower than that obtained from the T-bill futures (median difference is approximately 15 basis points lower). As a result of using a forward interest rate that is higher than the spot risk-free rate Knight–Ridder incorporates upward bias of 0.03% on average into the implied

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[4] To see if Whaley (1986, p. 134) and Barone-Adesi and Whaley (1987, p. 314) results for at-the-money options hold for the nearest-to-the-money options as well, we follow Jorion (1995, p. 512) and conduct the following exercise for the S&P 500 Index futures option, one agricultural, one energy, one livestock, one long-term interest rate, one short-term interest rate, one currency, and one metal option contract (the first option for each group of contracts in Table 1). First, we look at the differences in option prices obtained from the Black (1976) model for European futures options and Barone-Adesi and Whaley (1987) model for American futures options. For each type of contract, we use a representative futures price (F), (separately) the two closest-to-the-money exercise prices (X), a representative futures price volatility (σ), the representative time until option maturity (T) of 60 days, and the representative risk-free interest rate (r) of 6%. The Black (1976) model underestimates option price by less than 1% of option's value. For any of the eight representative contracts we conduct the exercise for, the difference between the nearest-to-the-money European and American option prices are not higher than 1% of the option's price. Second, we approximate the measurement error that Knight–Ridder introduces in the implied volatility estimate by relying on the Black (1976) model rather than the Adesi-Barone and Whaley (1987) analytic approximation. We do this by approximating the bias in implied volatility that would arise as a result of the Black (1976) model's underestimating the option price. As the Knight–Ridder dataset does not contain option prices, for this exercise we treat option prices obtained from the Adesi-Barone and Whaley (1987) analytical approximation as the true option price. The results of our exercise indicate that relying on the Black model over-estimates implied volatilities in a systematic way. However, the bias is less than 0.25% of implied volatility for most of the contracts, and less than 0.94% of for all of the contracts that we examine. Since the bias is systematically affecting all observations, most of it is eliminated when we difference the daily implied volatilities in our analysis. We conclude that, for the nearest-to-the-money futures options, the effect of relying on the Black (1976) model rather than the Adesi-Barone and Whaley (1987) analytic model is negligible.

volatility estimate. We do not believe that this measurement error is large enough to materially affect our results. [5]

Knight–Ridder calculates implied volatility based on the number of calendar days until maturity. However, several studies have suggested that implied volatilities are more appropriately calculated using trading days to expiration, because the variance of returns over the Friday close to Monday close (weekend) period is not even close to being three times higher than the variance over a normal trading day, e.g. Monday close to Tuesday close. For example, French and Roll (1986) find that the weekend return variance for all NYSE and AMEX stocks is only 10.7% greater than the normal trading day variance. Ederington and Lee (1996) report that the weekend return variance in T-Bond and Eurodollar Futures is actually less than the normal trading day variance, while in Deutschemark futures the weekend return variance is only slightly greater.

To our knowledge, the relative magnitudes of weekend and weekday return variances have not been examined for many of the markets examined in our study, particularly those in the agriculture, livestock, energy, and metals categories. Intuitively, there are reasons why results might vary across different categories of markets. If short-term interest rates are primarily driven by monetary policy, which is not normally conducted on weekends, we would not expect the weekend variance in T-bill or Eurodollar futures to be greater than the weekday variance. Conversely, if price changes in agricultural futures are strongly linked to the weather, we would expect substantially greater weekend variance because weather-related disturbances are indeed three times likelier to occur over a 72-hour period than over 24 hours.

We report results for normal weekday versus normal Monday (i.e. weekend) variances of underlying futures returns for 37 markets in Table 1. To simplify interpretation of the results, trading days following holidays are excluded from the analysis. If volatility is most strongly associated with calendar days, we would expect the variance ratio of Mondays versus other days to equal three, whereas if volatility is associated with trading days we would expect a variance ratio of one. As expected, our findings do vary across markets. Roughly consistent with previous studies of US equity markets, the variance ratio for S&P 500 index futures is approximately 1.18. Also consistent with previous studies, in US and UK interest rate futures (but, interestingly, not German and

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[5] Another concern is the measurement error that could be introduced if the trading in T-bill futures during our sample period was as low as it has been most recently. During the sample period, the average daily trading volume in T-bill futures is 7118, average daily open interest is 36,502. During the 1995–1998 period that follows our sample period, the average daily volume for T-bill futures fell to 1167, while the average daily open interest became 11,673. As a result, we do not believe that during our sample period infrequent trading, by itself, is a source of measurement error that affects Knight–Ridder's implied volatility calculations.

French interest rate futures), the variance ratio is actually below one. Conversely, variance ratios tend to be higher in many of the agricultural, energy and gold futures, with most other markets not markedly different than the S&P 500. Nevertheless, it is important to note that with one exception (Coffee), the observed variance ratios are all closer to one than to three, in most cases considerably closer.

Another way to test whether implied volatilities need to be adjusted to reflect trading days is to examine the behavior of the unadjusted, calendar day-based Knight–Ridder implies on Fridays versus Mondays. If weekend variances are not appreciably greater than normal weekday variances and market participants price options based on trading days to expiration, then (calendar day-based) implied volatility on Friday should be lower than on Monday. This is because holding an option's price and other factors constant, a decrease in time to expiration results in an increase in implied volatility, and the measured decrease in the time to expiration over the weekend (3 calendar days) is larger than the real change (1 day). Thus, as Ederington and Lee (1996) also argue, measuring implied volatility using calendar days imposes a spurious weekend effect, whereby implied volatilities appear to increase over the weekend. We examine this issue in Table 1, where for each market we report the mean unadjusted, calendar day-based implied volatility on Friday and Monday, and a one-sided *t*-statistic for the Friday mean minus the Monday mean. Consistent with our expectation that market participants measure implied volatility in trading days, the implieds are lower on Friday than on Monday in all 37 markets; the difference is significant in 29 cases at the 10% level or better.

All of the results in Table 1 show that, generally, implied volatilities are more properly associated with trading days left until option maturity. Therefore, following Whaley (1993) and consistent with the treatment in Jorion (1995), Fleming et al. (1995), and Ederington and Lee (1996), we adjust the Knight–Ridder implied volatilities as follows:

$$\mathrm{IV}_t = \mathrm{IVKR}_t * \sqrt{\frac{T_\mathrm{c}}{T_\mathrm{m}}}, \tag{1}$$

where IV_t and IVKR_t are, respectively, the adjusted and the Knight–Ridder implied volatilities on day t, and T_c and T_m are, respectively, the number of calendar days and trading days left until the expiration of the option.

From the Knight–Ridder data, we create a single daily implied volatility series for each contract. In doing so, we attempt to match each option contract as closely as possible to the underlying futures by using the futures option contract with the closest expiration date prior to the maturity date of the underlying futures contract. In most markets, this task is straightforward because futures and options on futures have the same monthly expiration cycle, but in some cases there are differences. For example, S&P 500 Index futures options are available with expirations for every month during the year, but

Table 1
Variance of close-to-close futures returns and calendar day implied volatilities, by day of the week[a]

		Futures return variance Monday versus other days					Futures option implied volatility Monday versus Friday		
		Normal Tuesday, Wednesday, Thursday, or Friday		Normal Monday		Monday versus other days	Mean implied volatility (%) as reported by Knight–Ridder		
		# of Obs.	Variance	# of Obs.	Variance	Variance ratio	Friday	Monday	*t*-test
S&P 500 index		1172	0.0000573	276	0.0000678	1.183	13.1174	13.7681	−2.1362**
Agricultural products									
Corn	(CME)	1140	0.0000657	267	0.0001220	1.857	16.4340	17.1386	−1.7553**
Soybeans	(CME)	1140	0.0000938	267	0.0001530	1.631	15.9764	16.5257	−1.3414*
Soybean meal	(CME)	1140	0.0000981	267	0.0001720	1.753	14.8137	15.6583	−1.9972**
Soybean oil	(CME)	1140	0.0001270	267	0.0001830	1.441	17.4291	18.0299	−1.6713**
Wheat	(CME)	1140	0.0001440	267	0.0001700	1.181	16.2150	16.8979	−2.8785***
Coffee	(CSCE)	939	0.0002900	213	0.0010410	3.590	28.0492	28.9142	−0.9721
Sugar	(CSCE)	939	0.0002840	213	0.0003800	1.338	25.5816	26.3938	−1.7477**
Cocoa	(CSCE)	939	0.0002150	213	0.0003060	1.423	26.6054	27.4993	−2.9288***
Cotton	(NYCE)	954	0.0001490	218	0.0001940	1.302	16.4407	17.0119	−2.7388***
Orange juice	(NYCE)	938	0.0003070	212	0.0003040	0.990	26.0813	27.7321	−1.8672**
Energy products									
Crude oil	(NYME)	1164	0.0004900	267	0.0006950	1.418	25.7914	27.3789	−1.2594
Heating oil #2	(NYME)	1152	0.0003580	264	0.0004900	1.369	24.2320	25.5161	−1.1001
Gasoline-unleaded	(NYME)	1111	0.0003930	255	0.0004810	1.224	24.6671	26.1757	−1.2835*
Natural gas	(NYME)	427	0.0002280	96	0.0003540	1.553	31.2276	33.7938	−2.5290***
Livestock									
Live cattle	(CME)	1998	0.0000753	471	0.0000867	1.151	12.2842	12.7132	−1.6273**
Live hogs	(CME)	1949	0.0001500	458	0.0001820	1.213	18.0005	18.6344	−2.5255***
Feeder cattle	(CME)	1568	0.0001880	366	0.0002750	1.463	8.5460	8.8010	−1.2625*

Long-term interest rates									
Treasury bonds	(CBOT)	1237	0.0000424	292	0.0000324	0.764	8.0037	8.3220	−3.3896***
T-notes, 10 year	(CBOT)	1237	0.0000192	292	0.0000153	0.797	5.7278	5.9764	−3.6951***
T-notes, 5 year	(CBOT)	894	0.0000076	213	0.0000061	0.803	3.6880	3.8542	−2.4080***
Long UK gilt	(LIFFE)	877	0.0000328	210	0.0000272	0.829	7.7034	7.8003	−0.3849
German bond	(LIFFE)	878	0.0000105	210	0.0000133	1.267	4.3170	4.4493	−0.9217
French long bond	(MATIF)	725	0.0000101	171	0.0000121	1.198	4.3310	4.4790	−0.9156
Short-term interest rates									
T-bills	(CME)	1173	0.0000004	277	0.0000003	0.788	15.9827	16.7468	−2.3574***
Eurodollar (Ed)	(CME)	1925	0.0000008	453	0.0000005	0.587	13.3928	14.0899	−2.7025**
Eurodollar (E-)	(LIFFE)	867	0.0000012	213	0.0000007	0.569	14.3530	15.0740	−1.6015*
EuroMark	(LIFFE)	859	0.0000006	205	0.0000009	1.477	9.7306	10.3145	−1.8517**
Short Sterling	(LIFFE)	859	0.0000019	205	0.0000016	0.835	12.1086	12.8296	−1.5206**
Currencies									
Japanese Yen	(CME)	1737	0.0000430	407	0.0000474	1.102	10.3090	10.5325	−1.0758
Deutschemark	(CME)	2153	0.0000648	508	0.0000838	1.293	9.9489	10.2933	−3.0255***
Canadian Dollar	(CME)	1679	0.0000094	394	0.0000122	1.298	4.2051	4.3479	−2.0790**
British Pound	(CME)	1938	0.0001110	456	0.0001490	1.342	9.8504	10.2676	−2.7981**
Swiss Franc	(CME)	1938	0.0000711	456	0.0000825	1.160	10.3319	10.7260	−3.3350***
Metals									
Copper	(COMEX)	909	0.0001610	210	0.0001400	0.870	18.5549	18.8867	−0.6508
Gold	(COMEX)	1170	0.0000461	270	0.0000798	1.731	11.7728	12.1449	−1.2951*
Silver	(COMEX)	1137	0.0001480	263	0.0001510	1.020	18.7304	19.3199	−0.9882

[a] For the futures contract return variances, trading days following holidays are excluded from the analysis. Consequently, for all normal Mondays, the time difference between the current day close and the previous close is 72 hours, and for all other normal trading days the time difference is 24 hours. The implied volatilities of futures options recreated by Knight–Ridder are based on calendar days to expiration. Because we expect Monday implied volatilities to be higher than Friday implied volatilities we use a one-sided t-test. The futures and the option futures contracts trade on Chicago Board of Trade (CBOT), Chicago Mercantile Exchange (CME), Coffee, Sugar, Cocoa Exchange (CSCE), Commodity Echange (COMEX), New York Cotton Exchange (NYCE), New York Mercantile Exchange (NYME), London International Financial Futures Exchange (LIFFE), and Marché á Terme Inernationale de France (MATIF).

[*] Statistical significance at the 10% level.

[**] Statistical significance at the 5% level.

[***] Statistical significance at the 1% level.

S&P 500 Index futures have only March, June, September, and December expirations. For the S&P 500 Index futures options, we use only the March, June, September, and December expiration options contracts because they have the closest expiration date prior to the expiration of the futures contracts. [6]

We use the implied volatility from the nearby options until five trading days prior to option expiration, when we switch to the next deferred option series. This procedure allows us to provide uniformity in volatility estimation across the contracts we examine. To eliminate any effect at the time of rollover that could arise if the change in implied volatility (ΔIV) is calculated as the difference of the on-the-run contract volatility and the first-deferred contract volatility, we measure the daily change in implied volatility only from the identical contract. Therefore, on the day of rollover, we gather implied volatility for both the nearby and first-deferred contracts so that the change in the implied volatility is measured with the same contract month.

4. Statistical procedure

To estimate the relation between implied volatility and the returns on the underlying contracts, we use the first difference of the implied volatility (ΔIV), instead of implied volatility (IV), to compare ΔIV to changes in prices of the futures contracts and also because the change in the implied volatility is of primary interest to practitioners. As previous studies report a significant relation between volatility and volume (Karpoff, 1987) and the effect of market depth on the relation, where the open interest proxies the market depth (Bessembinder and Seguin, 1993), we control for market depth and information

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[6] A special note should also be given to the calculation of implied volatility in short-term financial futures options. The short-term interest rate futures used in this study are the Eurodollar and Treasury Bill from the Chicago Mercantile Exchange, and the Eurodollar, Euromark and Short Sterling from the London International Financial Futures Exchange. For these contracts, we find the continuous change of the applicable discount yield calculated from the reported index levels. For example, the price quotation of the Treasury Bill futures uses the International Money Market (IMM) Index which is a function of the discount yield:

$$\text{IMM Index} = 100.00 - \text{Discount Yield}.$$

Also, the Eurodollar uses the IMM Index for its price quotation and uses the London Interbank Offer Rate (LIBOR) as London dominates the Eurodollar deposit market. The quoted price is:

$$\text{IMM Index} = 100.00 - \text{LIBOR}.$$

Since we use the calculated rates instead of the quoted price and the exercise price is reported with the same terms as the price quotation, the call option behaves as a put option and the put option behaves as a call option.

inflows using the open interest and the volume of the futures contracts, both of which Knight–Ridder provides. Also, we control for volatility persistence (well established in previous studies) by including lagged values of ΔIV.

The regression used in this analysis is

$$\Delta \mathrm{IV}_t = c + \sum_{i=0}^{10} \omega_i \mathrm{VOL}_{t-i} + \sum_{i=0}^{10} \eta_i \mathrm{OI}_{t-i} + \sum_{i=1}^{10} \delta_i \Delta \mathrm{IV}_{t-i} + \sum_{i=0}^{10} \theta_i \mathrm{Prtn}_{t-i} + \sum_{i=0}^{10} \phi_i \mathrm{Nrtn}_{t-i} \tag{2}$$

where $\Delta \mathrm{IV}_t$ is $\mathrm{IV}_t - \mathrm{IV}_{t-1}$, VOL is the trading volume of underlying futures contract, OI the open interest of underlying futures contract, Prtn is $\ln(P_t/P_{t-1})$ of the underlying futures if the return is positive, 0 otherwise and Nrtn is $-\ln(P_t/P_{t-1})$ of the underlying futures if the return is negative, 0 otherwise.

We use a lag length of ten as in Bessembinder and Seguin (1993). By using a large number of lags, we are not necessarily attempting to find the best fit for each series, but instead we are controlling for high order autocorrelation.

From Eq. (2), we test the asymmetry of responses of the implied volatility against the sign of the return with the following null hypothesis:

$$\mathrm{H}_0 : P(0\text{–}10) = N(0\text{–}10),$$

where $P(0\text{–}10) = \sum_{j=0}^{10} \theta_j$ is the combined effect on IV of positive contemporaneous and lagged returns up to 10 lags and $N(0\text{–}10) = \sum_{j=0}^{10} \phi_j$ is the combined effect on IV of negative contemporaneous and lagged returns up to 10 lags.

If futures options markets are weak-form efficient, implied volatilities should reflect all information subsumed in past futures prices. Thus, only contemporaneous futures price changes should have an impact on IV. This reasoning leads to another set of testable null hypotheses:

$\mathrm{H}_0^{\mathrm{P}} : P(1\text{–}10) = 0,$ for the positive returns;
$\mathrm{H}_0^{\mathrm{N}} : N(1\text{–}10) = 0,$ for the negative returns; and
$\mathrm{H}_0^{\mathrm{C}} : P(0) = N(0)$ to test whether the contemporaneous coefficients on positive and negative returns are equal,

where $P(1\text{–}10) = \sum_{j=1}^{10} \theta_j$ is the combined effect on IV of lagged positive returns from one to ten trading day lags, $N(1\text{–}10) = \sum_{j=1}^{10} \phi_j$ the combined effect on IV of lagged negative returns from 1 to 10 trading day lags, $P(0) = \theta_0$, and $N(0) = \phi_0$.

$\mathrm{H}_0^{\mathrm{P}}$ is the null hypothesis to test the lagged effect of positive returns and $\mathrm{H}_0^{\mathrm{N}}$ is for the lagged effect of negative returns. Since the Ljung and Box (1978) *Q*-statistics for up to 30 lags are insignificant for all series tested, we do not adjust for either heteroskedasticity or for serial correlation in the residual term.

Therefore the test statistic for H_0 (the null hypothesis testing the symmetry of impact on IV of positive and negative futures returns) and H_0^C have *F*-distributions, and the test statistics for H_0^P, H_0^N, have *t*-distributions.

It is possible that for at least some of the contracts, estimates of regression (2) could be highly sensitive to outliers in the dependent variable. Infrequent but largely unexpected events can result in very large changes in implied volatilities. Therefore, we test the robustness of our results against outliers. Specifically, we apply filters to drop outliers from our sample and re-estimate regression (2). The filters we use drop observations that fall outside the mean change in implied volatility plus or minus two, three and four standard deviations. The examination of the filtered coefficient estimates for Prtn and Nrtn showed that our results are robust to outliers with the major exception of the Energy contracts (results are not reported to conserve space and are available upon request).

For the energy contracts, the results obtained from the full-sample and those obtained from the filtered subsamples differ significantly. For all of the energy contracts, the coefficient estimates for Prtn(0) are positive and significant irrespective of the sample used. However, for crude oil, heating oil and gasoline, Nrtn(0) is negative and significant for the full sample, but positive and significant for the subsamples obtained after filtering. Thus, clearly, the negative coefficient estimates for the full sample are due to outlier observations for these three contracts. One explanation for the dramatic impact of outliers for the energy contracts is the events surrounding the Gulf war period. To investigate this possibility we examined the dates for those observations that are dropped from our sample due to filtering, and indeed, the largest outliers occur during August 1990 when Iraq invaded Kuwait, and during January 1991 when the allied coalition began air attacks on Iraq. In the subsequent analysis, we estimate model (2) and report results for the energy contracts using only post-Gulf war data (April 1991–November 1994). These postwar results are very similar to the full period results with large outliers excluded. [7]

5. Results

We report regression results for implied volatility as the dependent variable with volume, open interest, lagged values of implied volatility and underlying asset returns as independent variables in Table 2. The full sample results are in Table 2. Results for the periods before and after October 1987, and results for subsamples with outliers in the dependent variable excluded, are reported in

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[7] We are grateful to an anonymous referee for suggesting these robustness checks and for calling our attention to this potential problem in the energy contracts.

Appendix Tables (available from the authors upon request). We report the contemporaneous coefficients and the sum of the lagged coefficients in the tables. Since we multiply the negative returns by negative one, positive (negative) coefficients represent an increase (decrease) in implied volatility for positive returns and for negative returns. Since most previous studies examined stocks, we will pay particular attention in describing the results to the S&P 500 Index futures options and then discuss each category of options in turn.

Overall, we find large and statistically significant contemporaneous responses for the contracts. That is, in general the coefficients for Prtn(0) and Nrtn(0) are statistically significant. The coefficient estimate for Prtn(0) is significant for 27 of the 37 option contract series. Only Treasury-Bill and Short-Sterling contracts have no contemporaneous relation for either Prtn(0) or Nrtn(0) with implied volatility. For most of the contracts implied volatility is affected by price changes in the underlying asset.

The combined lagged effects $P(1\text{–}10)$ and $N(1\text{–}10)$ are statistically significant for sixteen options series and fourteen options series, respectively. Thus, there are no consistent lagged effects across all contracts. The insignificance of the combined lagged effect could be a result of using a long lag interval. We chose 10 trading day lags in an attempt to incorporate all potential lagged effects. The significance of the lagged coefficients vary across individual contracts. The individual coefficients of the lagged returns show significance for at least one of the lagged returns for all contracts. When we examine the individual lagged coefficients, we observe a large number of significant coefficients for the lagged returns. [8]

For all contracts except Short-Sterling and British pound the coefficient for $\Delta \mathrm{IV}_{-1}$ is negative and significant. This finding, which implies mean-reversion in implied volatilities, is consistent with earlier studies of implied volatility in equity markets (Stein, 1989; Harvey and Whaley, 1992), Treasury Bonds (Ederington and Lee, 1996; Simon, 1997), and currencies (Jorion, 1995; Ederington and Lee, 1996). As noted by Stein (1989) and Ederington and Lee (1996), negative correlation between contemporaneous and lagged values of

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[8] At the suggestion of one of the referees we tested whether lags 6–10 for the explanatory variables are warranted in our specifications. Specifically, we tested for the null-hypothesis that lags six through ten are jointly equal to zero. We rejected the null hypothesis at the 5% level for 17 out of 37 contracts that we study. Specifically we rejected the null hypothesis for the S&P 500; all the agricultural commodities except soybean oil, wheat, cotton, and orange juice; all the energy contracts except natural gas; all the livestock contracts; all the short-term interest contracts except T-bills, Eurodollar (E-), and Short-Sterling. We could not reject the null-hypothesis at the 5% level for any of the long-term interest rate contracts; and for any of the currency contracts except the British Pound. Our conclusions are not materially affected if we exclude lags 6–10 in those cases where we failed to reject the null hypothesis. When reporting the results in Table 2, in order to be consistent across specifications for different contracts, we kept all 10 lags for all the contracts.

Table 2
The reaction of implied volatility to market returns on the underlying futures contracts[a]

	Constant	Vol_0	OI_0	$\Delta\mathrm{IV}_1$	Prtn_0	Nrtn_0	$\frac{H_0^C:}{\mathrm{Prtn}_0} = \mathrm{Nrtn}_0$
S&P 500 Index							
(CME)	−0.297	0.000	0.000	−0.090	−18.314	117.464	1,962.84*
	(−4.551)*	(−0.157)	(2.024)**	(−4.892)*	(−6.039)*	(56.046)*	
Agricultural products							
Corn (CBOT)	0.016	0.000	0.000	−0.223	91.000	−9.007	147.95*
	(0.058)	(−1.483)	(1.572)	(−8.228)*	(10.994)*	(−1.184)	
Soybean (CBOT)	0.489	0.000	0.000	−0.161	95.471	−41.834	393.23*
	(2.377)**	(0.075)	(1.294)	(−5.883)*	(13.659)*	(−6.509)*	
Soybean meal (CBOT)	0.118	0.000	0.000	−0.195	105.540	−12.769	311.43*
	(0.406)	(−0.789)	(1.851)***	(−7.211)*	(15.52)*	(−1.927)*	
Soybean oil (CBOT)	0.542	0.000	0.000	−0.311	69.573	−15.096	94.64*
	(1.360)	(−1.196)	(1.557)	(−11.549)*	(7.905)*	(−1.724)***	
Wheat (CBOT)	−0.300	0.000	0.000	−0.115	39.788	7.323	45.99*
	(−1.271)	(2.036)**	(−0.614)	(−4.262)*	(7.780)*	(1.472)	
Coffee (CSCE)	−0.928	0.000	0.000	−0.161	44.401	19.209	22.74*
	(−2.356)**	(0.226)	(0.311)	(−5.332)*	(9.729)*	(3.512)*	
Sugar (CSCE)	−0.110	0.000	0.000	−0.175	51.848	15.272	37.40*
	(−0.328)	(3.686)*	(0.924)	(−5.696)*	(7.638)*	(2.368)*	
Cocoa (CSCE)	0.030	0.000	0.000	−0.475	29.456	4.305	27.47*
	(0.107)	(2.644)*	(0.513)	(−15.797)*	(6.357)*	(0.838)	
Cotton (NYCE)	−0.169	0.000	0.000	−0.176	20.370	18.648	0.17
	(−0.952)	(0.244)	(1.777)[c]	(−5.868)*	(4.667)*	(4.295)*	
Orange juice (NYCE)	0.035	0.000	0.001	−0.242	28.686	27.291	0.02
	(0.116)	(−0.356)	(1.631)	(−7.974)*	(2.479)**	(2.589)**	
Energy contracts							
Crude oil 1/11/1989	−0.003	0.000	0.000	−0.117	26.769	43.752	4.58**
	(−0.007)	(2.969)*	(1.089)	(−3.381)*	(3.182)*	(5.620)*	
Heating oil #2 (NYME)	0.276	0.000	0.000	−0.190	21.845	5.382	4.52**
	(1.029)	(3.860)*	(1.408)	(−5.509)*	(2.686)*	(0.697)	
Gasoline unleaded (NYME)	−0.213	0.000	0.000	−0.220	18.339	23.003	1.58**
	(−0.609)	(1.278)***	(2.538)*	(−6.420)*	(2.349)**	(3.145)*	
Natural gas (NYME)	−1.156	0.000	0.000	−0.230	24.115	52.058	8.45*
	(−1.578)	(3.003)*	(2.465)**	(−4.902)*	(2.500)**	(5.93)*	
Livestock							
Live cattle (CME)	−0.171	0.000	0.000	−0.091	1.592	52.226	220.45*
	(−1.584)	(−0.395)	(−1.15)	(−4.513)*	(0.446)	(14.656)*	
Live hogs (CME)	−0.005	0.000	0.000	−0.269	2.012	15.196	11.16*
	(−0.035)	(−3.279)*	(−3.06)*	(−13.243)*	(0.503)	(3.743)*	
Feeder cattle (CME)	0.034	0.000	0.000	−0.239	−5.197	57.267	132.14*
	(0.518)	(−1.155)	(1.017)	(−10.418)*	(−0.973)	(10.107)*	
Long-term interest rates							
Treasury bonds (CBOT)	0.019	0.000	0.000	−0.165	−18.541	31.624	195.80*
	(0.315)	(−3.536)*	(1.914)[c]	(−6.319)*	(−5.119)*	(8.620)*	
T-notes – 10 yr (CBOT)	−0.064	0.000	0.000	−0.094	−13.287	37.059	191.85*
	(−2.127)**	(−1.769)***	(1.134)	(−3.595)*	(−3.816)*	(10.332)*	
T-notes – 5 yr (CBOT)	0.012	0.000	0.000	−0.126	−10.073	37.661	91.46*
	(0.413)	(−0.116)	(−1.126)	(−4.056)*	(−2.140)**	(7.889)*	
Long UK gilt (LIFFE)	−0.008	0.000	0.000	−0.704	−2.665	33.430	8.60*
	(−0.073)	(1.045)	(0.082)	(−22.627)*	(−0.269)	(3.131)*	
German bonds (LIFFE)	−0.042	0.000	0.000	−0.067	−5.956	45.042	139.58*
	(−1.431)	(4.048)*	(0.537)	(−2.141)*	(−1.408)	(10.336)*	
French long bonds (MATIF)	−0.081	0.000	0.000	−0.175	−28.329	82.223	301.94*
	(−1.892)*	(2.307)**	(2.269)**	(−5.137)*	(−4.494)*	(13.041)*	

$\sum \text{Vol}_{1-10}$	$\sum \text{OI}_{1-10}$	$\sum \Delta\text{IV}_{2-10}$	$\text{H}_0^\text{P}: \sum \text{Prtn}_{1-10} = 0$	$\text{H}_0^\text{N}: \sum \text{Nrtn}_{1-10} = 0$	$\text{H}_0: \sum \text{Prtn}_{0-10} = \sum \text{Nrtn}_{0-10}$	Adj.-R^2	No. of Obs.	Sample period
0.000 (2.100)**	0.000 (−2.040)**	−0.386 (−6.555)*	−9.906 (−1.232)	−50.083 (−6.831)*	48.58*	0.67	2990	1/28/1983 –12/7/1994
0.000 (−0.193)	0.000 (−1.489)	−0.417 (−3.638)*	−47.173 (−1.865)	−43.409 (−2.145)**	8.20*	0.17	1432	2/24/1989 –11/10/1994
0.000 (−0.034)	0.000 (−1.410)	−0.244 (−2.210)**	−52.681 (−2.486)**	−25.314 (−1.436)	13.58*	0.30	1412	2/24/1989 –10/13/1994
0.000 (0.833)	0.000 (−1.854)***	0.009 (0.094)	−90.969 (−4.711)*	−49.512 (−2.784)*	8.10*	0.27	1432	2/24/1989 –11/10/1994
0.000 (0.084)	0.000 (−1.583)	−0.097 (−0.854)	−45.790 (−1.880)***	−31.647 (−1.399)	4.91**	0.15	1432	2/24/1989 –11/10/1994
0.000 (−1.819)^{c}	0.000 (0.733)	−0.157 (−1.601)	−4.919 (−0.344)	16.688 (1.177)	0.46	0.09	1432	2/24/1989 –11/10/1994
0.000 (2.867)**	0.000 (−0.294)	−0.235 (−2.130)**	−47.651 (−4.178)*	−14.480 (−1.078)	0.22	0.18	1160	3/5/1990 –10/27/1994
0.000 (−0.273)	0.000 (−1.085)	−0.370 (−3.240)*	−11.013 (−0.549)	−10.403 (−0.639)	2.07	0.17	1119	3/6/1990 –8/31/1994
0.000 (−0.747)	0.000 (−0.638)	−0.685 (−4.305)*	5.893 (−0.429)	3.392 (0.216)	2.51	0.24	1160	3/6/1990 –10/27/1994
0.000 (0.429)	0.000 (−1.801)^{c}	0.004 (0.037)	18.847 (1.287)	−22.203 (−1.955)^{c}	7.56*	0.07	1189	3/6/1990 –11/4/1994
0.000 (0.582)	−0.001 (−1.646)^{c}	−0.672 (−4.711)*	−60.655 (−2.314)**	−23.623 (−0.909)	1.18	0.09	1140	3/2/1990 –9/29/1994
0.000 (−0.264)	0.000 (−1.184)	−0.515 (−3.691)*	−74.285 (−3.130)*	−15.721 (−0.740)	7.33*	0.14	898	4/1/1991 –11/3/1994
0.000 (−1.570)	0.000 (−1.490)	−0.527* (−3.666)	−57.980 (−2.621)*	9.323 (0.395)	3.67***	0.09	905	4/1/1991 –11/14/1994
0.000 (0.851)	0.000 (−2.622)*	−0.257 (−1.874)***	−57.625 (−2.627)*	10.444 (0.481)	6.77*	0.08	905	4/1/1991 –11/14/1994
0.000 (−1.092)	0.000 (−2.366)**	−0.361 (−1.910)***	28.516 (1.073)	32.411 (1.060)	1.09	0.18	519	10/2/1992 –11/10/1994
0.000 (−0.129)	0.000 (1.288)	−0.338 (−4.768)*	−39.733 (−3.796)*	19.459 (1.889)***	65.45*	0.15	2536	10/30/1984 –11/23/1994
0.000 (−0.517)	0.000 (3.264)*	−0.168 (−2.069)**	1.891 (0.165)	15.223 (1.384)	3.97**	0.09	2471	2/1/1985 –11/23/1994
0.000 (−1.011)	0.000 (−1.128)	−0.005 (−0.056)	−1.053 (−0.072)	−2.719 (−0.182)	11.74*	0.12	1972	1/9/1987 –11/9/1994
0.000 (0.483)**	0.000 (−1.988)	−0.360 (−2.999)*	10.662 (0.899)	29.929 (2.289)**	24.97*	0.17	1475	9/1/1988 –11/10/1994
0.000 (−0.671)	0.000 (−0.989)	−0.313 (−3.287)*	16.950 (1.541)	15.703 (1.373)	13.53*	0.16	1556	9/1/1988 –11/10/1994
0.000 (−0.052)	0.000 (1.017)	−0.313 (−3.287)*	11.650 (0.821)	−15.084 (−1.099)	1.45	0.11	1556	5/24/1990 –11/10/1994
0.000 (−0.175)	0.000 (−0.192)	−1.479 (−6.407)*	−82.801 (−2.564)**	82.825 (2.585)**	17.27*	0.33	1089	7/13/1990 –11/10/1994
0.000 (−2.824)*	0.000 (−0.519)	−0.002 (−0.021)	4.886 (0.432)	−7.699 (−0.697)	6.19*	0.20	1090	7/12/1990 –11/15/1994
0.000 (−1.031)	0.000 (−2.148)	−0.114 (−0.878)	−22.339 (−1.306)	−14.870 (−0.806)	22.34*	0.34	910	3/11/1991 –11/16/1994

Table 2 (*Continued*)

	Constant	Vol_0	OI_0	ΔIV_1	Prtn_0	Nrtn_0	$\frac{\text{H}_0^\text{C}:}{\text{Prtn}_0} = \text{Nrtn}_0$
Short-term interest rates							
T-bills							
(CME)	−0.009	0.000	0.000	−0.294	4.366	5.649	0.03
	(−0.045)	(0.584)	(0.501)	(−11.09)*	(0.595)	(0.834)	
Eurodollar (Ed)							
(CME)	−0.214	0.000	0.000	−0.275	26.963	39.584	5.32**
	(−2.747)*	(−4.987)*	(−5.984)*	(−13.368)*	(4.962)*	(9.106)*	
Eurodollar (E-)							
(LIFFE)	0.005	0.000	0.000	−0.303	0.984	11.318	1.48
	(0.034)	(−0.254)	(−0.247)	(−9.753)*	(0.124)	(1.728)***	
Euromark							
(LIFFE)	−0.040	0.000	0.000	−0.264	−12.288	49.815	46.48*
	(−0.494)	(−1.075)	(−0.802)	(−8.403)*	(−1.372)	(6.139)*	
Short Sterling							
(LIFFE)	−0.093	0.000	0.000	−0.033	1.738	−3.064	0.46
	(−0.801)	(0.249)	(1.089)	(−1.045)	(0.239)	(−0.683)	
Currencies							
Japanese Yen							
(CME)	−0.197	0.000	0.000	−0.107	40.912	26.809	21.63*
	(−3.721)*	(6.602)*	(2.607)*	(−4.975)*	(13.974)*	(8.359)*	
Deutsche Mark							
(CME)	−0.099	0.000	0.000	−0.086	27.128	27.402	0.01
	(−2.457)**	(6.185)*	(−0.325)	(−4.452)*	(10.899)*	(10.398)*	
Canadian Dollar							
(CME)	−0.015	0.000	0.000	−0.094	2.265	44.699	121.94*
	(−0.659)*	(1.801)*	(0.543)	(−4.279)*	(0.578)	(11.893)*	
British Pound							
(CME)	−0.129	0.000	0.000	−0.025	18.810	27.323	14.40*
	(−2.625)*	(4.475)*	(−0.700)	(−1.219)	(8.665)*	(11.915)*	
Swiss Franc							
2/25/1985	−0.211	0.000	0.000	−0.068	22.928	16.849	7.09*
	(−3.233)*	(5.212)*	(1.190)	(−3.322)*	(10.309)*	(7.148)*	
Metals							
Copper							
(COMEX)	−0.631	0.000	0.000	−0.131	17.200	16.503	0.01
	(−3.388)*	(1.795)****	(1.951)***	(−4.303)*	(2.931)*	(2.826)*	
Gold							
(COMEX)	0.039	0.000	0.000	−0.182	73.290	25.048	51.55*
	(0.305)	(0.883)	(3.719)*	(−6.800)*	(10.615)*	(4.570)*	
Silver							
(COMEX)	0.551	0.000	0.000	−0.074	73.275	19.552	95.94*
	(1.727)***	(0.805)	(2.008)**	(−2.742)*	(12.370)*	(3.537)*	

a

$$\Delta\text{IV}_t = \text{Constant} + \sum_{i=0}^{10}\omega_i\text{VOL}_{t-i} + \sum_{i=0}^{10}\eta_i\text{OI}_{t-i} + \sum_{i=1}^{10}\delta\Delta\text{IV}_{t-i} + \sum_{i=0}^{10}\theta_i\text{Prtn}_{t-i} + \sum_{i=0}^{10}\phi_i\text{Nrtn}_{t-i} + \varepsilon_i$$

where IV_t is the option contract's IV on day t, $\Delta\text{IV}_t = \text{IV}_t - \text{IV}_{t-1}$; VOL the total trading volume of the underlying futures contract; $\text{Nrtn} = -(P_t/P_{t-1})$ if futures return is negative, 0 otherwise; and ε_t is an error term. H_0 is the null hypothesis that the sum of the comtemporaneous and all lagged negative returns (Nrtn). H_0 is tested using a F-statistic. H_0^P (H_0^N) is the null hypothesis that the sum the coefficient estimates on contemporaneous positive and negative returns are equal. H_0^C is tested with a F-statistic. Futures and (CSCE), Commodity Exchange (COMEX), New York Cotton Exchange (NYCE), New York Mercantile Exchange (NYME),

* Statistical signifiance at 10% level.

** Statistical signifiance at 5% level.

*** Statistical signifiance at 1% level.

$\sum \text{Vol}_{1-10}$	$\sum \text{OI}_{1-10}$	$\sum \Delta \text{IV}_{2-10}$	$\text{H}_0^\text{P}: \sum \text{Prtn}_{1-10} = 0$	$\text{H}_0^\text{N}: \sum \text{Nrtn}_{1-10} = 0$	$\text{H}_0: \sum \text{Prtn}_{0-10} = \sum \text{Nrtn}_{0-10}$	Adj.-R^2	No. of Obs.	Sample period
0.000 (−0.765)	0.000 (−0.427)	−0.360 (−2.999)*	−5.407 (−0.275)	−21.028 (−1.202)	0.34	0.08	1475	1/3/1989 –11/10/1994
0.000 (2.257)**	0.000 (5.981)*	−0.426 (−4.640)*	1.868 (0.137)	−32.916 (−3.155)*	1.76	0.16	2451	3/20/1985 –12/9/1994
0.000 (0.174)	0.000 (0.460)	−0.068 (−0.562)	8.669 (0.473)	−41.369 (−2.243)**	2.47	0.08	1084	8/15/1990 –12/9/1994
0.000 (1.123)	0.000 (0.710)	−0.283 (−2.104)**	−4.079 (−0.164)	−14.546 (−0.737)	2.57	0.14	1084	8/15/1990 –12/9/1994
0.000 (2.304)**	0.000 (−1.251)	−0.324 (−3.000)*	−20.327 (−1.060)	−9.174 (−0.808)	0.08	0.08	1086	8/15/1990 –12/13/1994
0.000 (−3.299)*	0.000 (−2.310)**	−0.219 (−2.810)*	−3.669 (−0.437)	−30.653 (−3.461)*	14.55*	0.20	2203	3/5/1986 –12/1/1994
0.000 (−2.765)*	0.000 (0.010)	−0.243 (−3.535)*	−15.279 (−2.233)**	−13.434 (−1.922)*	0.07	0.13	2737	1/24/1984 –12/1/1994
0.000 (−0.302)	0.000 (−0.709)	−0.170 (−2.245)**	−33.890 (−3.196)*	−1.631 (−0.171)	29.74*	0.12	2130	6/18/1986 –12/1/1994
0.000 (−0.941)	0.000 (0.280)	−0.183 (−2.656)[a]	−1.191 (−0.214)*	−16.280 (−2.944)*	0.82	0.13	2462	2/25/1985 –12/1/1994
0.000 (−1.406)	0.000 (−1.174)	−0.210 (−2.954)*	2.529 (0.396)	−11.381 (−1.692)[c]	6.71*	0.10	2462	2/25/1985 –12/1/1994
0.000 (−1.432)	0.000 (−1.702)***	−0.078 (−0.769)	33.531 (2.081)**	14.978 (0.972)	1.01	0.09	1140	4/26/1990 –11/15/1994
0.000 (−0.722)	0.000 (−3.782)*	−0.400 (−3.879)*	5.826 (0.294)	−54.221 (−3.311)*	18.86*	0.16	1464	1/3/1989 –11/3/1994
0.000 (−0.337)	0.000 (−2.288)**	−0.290 (−3.207)*	−29.777 (−1.876)***	−19.819 (−1.480)	4.63**	0.19	1421	3/3/1989 –11/3/1994

OI the total open interest of the underlying futures contract; Prtn $= \ln(P_t/P_{t-1})$ if futures return is positive, 0 otherwise; coefficient estimates on contemporaneous and all lagged positive returns (Prtn) is equal to the sum of the coefficient estimates of of all lagged positive returns (negative returns) is equal to zero. H_0^P and H_0^N are tested using t-statistics. H_0^C is the null hypothesis that their options trade on Chicago Board of Trade (CBOT), Chicago Mercantile Exchange (CME), Coffee, Sugar,and Cocoa Exchange London International Financial Futures Exchange (LIFFE), and Marche a Terme Inernationale de France (MATIF).

ΔIV may be due, in part, to measurement error in IV. In addition, Jorion (1995) shows that in currencies, the degree of mean reversion in implied volatility is similar to that observed in future realized volatility over the remaining life of the option.

We next examine whether the Fleming et al. (1995) findings are supported in our sample period. While they use an option on the S&P 100 Stock Index, we use the S&P 500 Futures Option because this is the data available from Knight–Ridder. For the option on the S&P 500 Index futures, the contemporaneous response for Prtn(0) and Nrtn(0) are both significant at 0.01. Notice, however, that the effect of Nrtn(0) is nominally larger (as is its t-statistic) than that of Prtn(0) indicating that the response to negative returns is much larger than to positive returns. There is also a significantly positive relation for OI_0 and a negative relation with the lagged value for implied volatility, IV_{-1}. These results are consistent with Fleming et al. (1995). For this contract, P(1–10) is insignificant, but N(1–10) is significant at the 0.01 level. The F-statistic that directly tests H_0, i.e. asymmetric impact of positive and negative futures returns including lags, is very large and significant at the 0.01 level. Thus, these results for the S&P 500 index clearly indicate a contemporaneous and lagged asymmetric relation between IV and underlying asset price changes, with positive futures returns (on balance) reducing IV and negative futures returns increasing IV.

The implied volatilities on agricultural futures on the Chicago Board of Trade, for Corn, Soybeans, Soybean Meal, Soybean Oil and Wheat contracts, react differently to underlying price changes than in the case of the S&P 500 Index. The contemporaneous response to the positive returns [Prtn(0)] are all positive and significant at the 0.01 level and most of the coefficients for the negative returns [Nrtn(0)] are negative, and in the case of the three soybean contracts, significant. When one adds the contemporaneous and lagged coefficients on the positive and on the negative returns, the same general pattern remains; on balance, positive returns are associated with increases in IV and negative returns (except for Wheat) with decreases in IV. This asymmetric relation, as evidenced by the F-statistics, is significant at the 0.01 level for Corn, Soybeans and Soybean Meal, and at the 0.05 level for Soybean Oil.

The significant increase in implied volatility when there is a positive return could be due to expected shortages of these commodities. For example, an increase in the price of the futures contract may arise from a weather-related shortage in the underlying commodity, and there may be no rebound from the shortage of supply until the following crop year. Thus, a large price increase could impose considerable uncertainty on the behavior of the agricultural commodities' prices in the future. On the other hand, a decrease in futures prices implies an excess supply for those commodities that will be placed in storage, and hence prevail for a considerable period of time. Since the excess supply can be pulled from storage when needed traders may be less uncertain about future price behavior, so implied volatility may decline. In support of this argument, we find that the sum of the lagged coefficients for these options for Prtn(1–10) and Nrtn(1–10) are generally negative.

There are five additional agriculture based options: coffee, sugar and cocoa that trade on the CSCE, and cotton and orange juice that trade on the NYCE. For these contracts the coefficient for Prtn(0) is significantly positive but the coefficient for Nrtn(0) is also positive and significant for 4 of the 5 contracts. In addition, when one also considers the impact of lagged returns, no consistent pattern emerges, and with the possible exception of Cotton there is little evidence of asymmetric futures return impacts on IV. It is interesting to note that the reaction for Prtn(0) is consistent across all the agriculture contracts, but the coefficients for Nrtn(0) vary across exchanges (negative on the CBOT, positive on the CSCE and NYCE).

We examine the energy contracts next. Here the implied volatility initially increases when there are positive returns on the underlying futures: Prtn(0) is positive and significant in all cases. However, for crude oil, heating oil and gasoline, the sum of the lags of Prtn are significantly negative and larger in absolute magnitude than the coefficient on contemporaneous Prtn. Thus, except for natural gas, positive returns in the underlying futures eventually result in reductions in implied volatility. On the other hand, in all cases, negative futures returns result in increased implied volatility, with most of the increase occurring contemporaneously.

The results for the three livestock contracts that trade on the CME are mixed. The *F*-statistics are significant, indicating that there is a disparate impact of positive versus negative futures returns, but all of the "action" appears to be in the negative returns, which result in increased IV. Except in Live Cattle, where the combined lagged positive returns are significantly negative, the positive returns do not have a significant impact. The livestock results bear a greater similarity to the S&P index than to those obtained with agricultural commodities.

The long-term interest rate futures, i.e. Treasury Bonds, 10-year Treasury Notes, and the 5-year Treasury Notes that trade on the CBOT, the Long UK Gilt and the German Bond that trade on LIFFE, and the French Long Bond that trades on the MATIF, have a contemporaneous negative coefficient for positive asset returns and all of them are significant except for the German Bond and the Long UK Gilt. For negative futures returns, all the long-term interest rate implied volatilities show significant contemporaneous positive coefficients. These findings are consistent with the results for the S&P 500 Index futures options but inconsistent with the findings for agricultural commodities. All of these contracts (with the exception of 5-year T-notes) also have a significant *F*-statistic for H_0. This result is consistent with there being an asymmetric response of implied volatility across positive and negative price changes. On the other hand, the combined lagged effect of the long-term interest rate futures does not strongly support our presumption. The French Long Bond is the only contract which shows a consistent sign for the contemporaneous and the combined lagged effects that we expect, but its coefficients are insignificant for Prtn(1–10) and Nrtn(1–10).

The positive returns in the long-term interest rate futures are associated with a decrease in interest rates. Since Knight–Ridder calculates implied volatility on short-term interest futures differently from long-term interest rate futures, the calculated positive return on short-term interest rate futures is associated with an increase in the interest rate. The coefficient for the positive returns on the short-term financial futures should be interpreted the same way as a coefficient of the negative returns in the long term interest rate futures. Out of the five short-term interest rate futures, none have *F*-statistics that are significant at the 0.10 level. In addition, the contemporaneous and the combined lagged effects of futures returns on implied volatility on these contracts are not consistent with our expectation based on earlier findings for stock index and long-term interest rate markets. The contemporaneous coefficient on positive futures returns, Prtn(0), should be negative based on our reasoning above, and we do not observe any significant negative coefficients [indeed, Prtn(0) is significantly positive for the CME-traded Eurodollars (Ed)]. To be consistent with the S&P 500, Nrtn(0) should be significantly negative for these contracts, but we generally observe the opposite. Thus, the short-term interest rate results are markedly different from those obtained for the S&P 500 or for long-term interest rates.

We examine the currency futures contracts next. Contrary to the S&P 500 Index futures, all of the currency futures have a significantly positive coefficient for the contemporaneous returns for both Prtn(0) and Nrtn(0) except for the coefficient for Prtn on the Canadian Dollar. All the combined lagged effects are also normally negative except for the combined lagged effect for positive returns on the Swiss Franc but several coefficients are insignificant. When we observe the contemporaneous response and the combined lagged response, most of the results are consistent with each other but not consistent with our a priori expectations.

The metals contracts on the COMEX have positive coefficients for Prtn(0) and Nrtn(0). However, the lagged sums are difficult to generalize; the only noteworthy finding is that the lagged negative returns for Gold have a significantly negative impact on IV. Considering the contemporaneous and lagged returns together, we find that positive Gold futures returns are associated with sharp increases in IV, while negative returns are associated with moderate decreases in IV; the *F*-statistic testing impact asymmetry is significant at the 0.01 level. For Silver, the findings are not as strong but similar to those for Gold, while for Copper there does not appear to be any asymmetry in the impact of positive and negative returns on IV. [9]

6. Conclusions

Implied volatility is the market's expectation for future volatility and can be inferred from option prices. This study examines the relation between implied volatility on futures options and the returns on the underlying futures, while controlling for most of the known factors that affect the volatility of the quoted price. We use 37 futures options from several exchanges, including foreign exchanges. Different from most previous studies using only stock market-related options, we use options on a broad variety of asset categories, including agriculture, interest rate, livestock, metals, energy, currencies and the S&P 500 Index futures. This set of data allows us to examine (1) whether findings from prior research on stock index futures options can be generalized to other contracts and (2) whether these findings are robust.

Prior research has documented that volatility in many financial asset markets is most directly related to trading rather than calendar days. We show that, with one exception, this finding can be extended to agricultural, livestock, energy, and metal futures, although (as expected) weekend variances in some of these markets are greater than in financial futures. We also find that implied volatility on S&P 500 Index futures and on long-term interest rate futures show a similar pattern of behavior to that of simple volatility measures used in prior research. This prior research has generally found an asymmetric reaction between positive and negative price changes and volatility (Schwert, 1989, 1990; Fleming et al., 1995). There is a much larger reaction for negative price changes than for positive price changes.

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[9] The stability of estimation for this section can be observed in Table 1-A and Table 2-A of Appendix. These tables are available from the authors on request. These tables contain results for the before and after market crash time periods. There are only ten contracts with sufficient data before and after the crash to permit this breakdown. The test results seem to vary as we use different time periods for the regression. One special caution should be given in interpreting the stability of estimation results here. The regression has 56 coefficients to estimate. This large number of coefficients will greatly reduce the power of the regression results by reducing degrees of freedom. Since the number of observations for the before market crash range from 120 for the Feeder Cattle to 1117 for the S&P 500 Index futures, some caution is required in interpreting them, especially for those contracts with a relatively small number of observations. However, we still observe an unstable estimation result for the S&P 500 Index futures which has a similar number of observations for the before and the after market crash subperiods. Therefore, the fact that we find inconsistent results from different subperiods supports the explanation that results of an empirical test cannot be generalized to other contracts and/or sample periods.

Our most significant finding is that results obtained for equities and long-term bond contracts cannot be generalized to other categories of contracts. While some markets, e.g. livestock and petroleum, do behave somewhat like stock and bond markets, most do not. In many of the agricultural and precious metal markets, for example, while we do find an asymmetric reaction in implied volatility to positive versus negative price changes in the underlying futures, the direction is exactly opposite to what is observed in equities: In these markets, increases in prices are associated with increases in implied volatility, while underlying price reductions are associated with decreases in implied volatility. In still other markets, e.g. in most currencies and short-term interest rate contracts, we find little evidence of any asymmetric reaction in implied volatility to positive versus negative returns on the underlying futures. Thus, the only sure conclusion to emerge from this analysis is that the relation between implied volatility and underlying asset returns varies greatly across different categories of underlying assets, and to a somewhat lesser extent across specific markets within a given asset category. Future research in delineating this relation is likely to prove fruitful.

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