Does Options Trading Lead to Greater Cash Market Volatility?

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INTRODUCTION

There has been much debate over the possible role of option and futures trading in the variability of cash markets. Despite the inconclusive (and even encouraging) nature of the evidence regarding the impact of options and futures trading activity, there has been a move toward increased regulation of these markets [See Bessembinder and Seguin (1992)]. Reconciling the relationship between derivative trading activity and the behavior of the underlying asset remains pivotal to whether or not the current regulatory measures an options and futures markets need to be reexamined.[1]

The purpose of this article is to examine the relationship between options trading activity on the S&P 100 Index and the variability of the

[1]Although it is by no means clear that regulators should be concerned about greater volatility in cash markets per se, some argue that higher volatility in cash prices reflected in higher costs of capital [see Edwards (1988b)].

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underlying cash index. The study contributes to the existing literature on derivative trading–asset behavior relationship in two respects. First, the study focuses on the impact of trading in index options on the underlying assets. While the relationship between index futures trading and the behavior of prices in cash markets has been extensively investigated, the influence of index options trading on cash market volatility remains to be examined. Thus, the study is motivated by the possibility that the dissimilarities across options and futures markets extend to their respective roles in price discovery, and to their impact on the volatility of the underlying cash markets.[2] Second, the study examines whether option trading intensifies the *rate of change in cash markets*. Such an investigation is especially relevant to the arguments for and against increased regulation of derivative trading on the grounds that speculative trading intensifies the onslaught of market calamities.

The evidence in this study indicates that while increased volatility in the cash market is followed by an increase in the level of option trading, an increase in option trading is followed by a decline in cash price variability. Moreover, there is no evidence to indicate that options trading intensifies the rate of change in cash market prices. Thus, the evidence supports the notion that option trading has a stabilizing influence on cash markets.

The second section of this article briefly reviews the past literature on the relationship between speculative trading activity and cash market volatility. The third section presents the data and methodology employed. The fourth section reports the empirical results. The final section summarizes the findings of the study.

EXISTING LITERATURE ON THE IMPACT OF SPECULATIVE TRADING ON CASH VOLATILITY

Opponents of speculative trading generally agree that increased trading in futures and options leads to unnecessary price volatility in the underlying cash markets. Some researchers argue that the participation of speculative traders in systems that allow high degrees of leverage lowers the quality of the information in the market [see for example, Figlewski

[2]There is a growing tendency for institutional investors to employ index options in their portfolios. There preferences for index options over index futures markets may be traced to the differences in the rights and obligations tied to the two securities. First, the option buyer can limit the investment risk to the premium of the option. Second, option buyers are not subject to margin calls and thus have *staying power*, the ability to maintain a market position during adverse market swings.

(1981) and Stein (1987)]. Cox (1976) and others note that uninformed traders could play a destabilizing role in cash markets. Others question the role of futures markets as representing an institutional alternative for accurate price forecasting [e.g., Martin and Garcia (1981)].

On the other hand, it has also been suggested that the futures and options markets have become important mediums of price discovery in cash markets [e.g., Schwarz and Laatsch (1991)]. Several authors have argued that trading in these markets improves the overall market's depth and informativeness [e.g., Powers 1970)], enhances market efficiency [e.g., Stoll and Whaley (1988)], increases market liquidity [e.g., Kwast (1986)], and compresses cash market volatility [e.g., Danthine (1978), Bray (1981), and Kyle (1985)].

Questions pertaining to the impact of derivative trading on cash market volatility have been empirically addressed in two ways. First, researchers have attempted to establish the impact of speculative trading on cash markets by comparing cash market volatilities during the pre- and post-futures/options trading eras. The majority of this class of studies suggests that speculative markets either add to the stability, or do not impact the volatility of cash markets.[3] For instance, Damodaran and Lim (1991) examine the effect of options listing on the return processes of firms and conclude that options listing lowers the variability of daily stock returns [Also see Simpson and Ireland (1985), Edwards (1988a), and Skinner (1989)].

Second, researchers have examined the relationship between speculative trading activity and cash markets by directly evaluating the impact of options and futures trading (generally proxied by trading volume) on the behavior of cash markets. While the evidence in the earlier studies in this set is inconclusive, the evidence in the more recent studies mainly suggests that futures and options trading activity either improves or has no impact on the stability of cash markets.[4] For instance, Edwards (1988b) and Bessembinder and Sequin (1992) provide evidence that futures trading improves the stability in equity indexes. Anthony (1988) finds that trading volume on options leads trading volume on stocks with a one-day lag, and suggests that informed traders are more likely to trade

[3]Exceptions include Harris (1989), who documents an increase in volatility in the S&P 500 Index subsequent to the introduction of futures trading on the Index.

[4]The inconclusiveness of the evidence from the earlier studies is probably best represented by the string of investigations into the impact of GNMA futures trading. For instance, while Froewiss (1978) documents a negative relationship between open interest in GNMA futures contracts and the volatility in the GNMA cash market, Figlewski (1981) provides evidence to suggest that futures trading increases the variability in the GNMA market. In a subsequent study, Simpson and Ireland (1985) find no evidence of an association between GNMA futures trading and GNMA cash prices.

options rather than stocks. Finally, Ely (1991) finds no evidence that would indicate the level of trading activity in interest-rate futures and options has a significant impact on the corresponding cash markets. In sum, the evidence on the impact of derivative on cash market variability has generally been favorable.

DATA AND METHODOLOGY

Data

This investigation employs trading data for put and call options written on the S&P 100 Index (ticker symbol OEX), and price levels of the underlying cash index. Data on the options volume and open interest (the number of contracts currently outstanding) for the OEX, and daily closing, and high and low values of the S&P 100 Index are obtained from the Pinnacle Data Corporation. The data series span the period, February 22, 1984 to November 1, 1993. The options on the S&P 100 Index represent the most actively traded contract on the Chicago Board Options Exchange (CBOE), with volume exceeding the combined volume of all CBOE-listed stock options [also see Harvey and Whaley (1992)].[5] These options have an end-of-day exercise and are cash settled. Since the emphasis of the study is to demonstrate the price–volume relationships in a general context rather than to isolate the nature of the relationships on and around expiration, the closing volume includes contracts of all expirations.

Given the imperfections in the measurement of variability in general, the investigation considers three measures of price volatility in addition to measures for the rate of change of market prices and option trading activity. The first of the volatility measures, also considered in studies by Rutledge (1979) and others, is the daily price range given by $P_t^H - P_t^L$, where P_t^H and P_t^L represent the high and low values of the S&P 100 Index respectively. The second volatility measure is the adjusted price range ($ADJR$) employed in Garcia, Leuthold and Zapata (1986). The series $ADJR_t$ ($t = 1 \,..\, n$) is constructed by adjusting the range if the low price at t exceeds the closing price (P^C) at $t - 1$, or

[5]Trading on the OEX began on March 11, 1983 when only 4575 contracts changed hands. Volume grew dramatically over the next year with May 11, 1984 producing a trading volume of 355,870 contracts. There is no comparable trend in volume for recent years, with high-volume days occurring on the pre-expiration Fridays of November, 1986 (695,000), January, 1987 (620,000), and February, 1993 (639,577). While there was limited trading in futures contracts on the S&P 100 in 1983 (0.4 million contracts) and in 1985 (0.2 million contracts), there has been no futures trading on this index since.

if the high at t is less than the closing price at $t-1$. In the event $P_t^L > P_{t-1}^C$, the difference $(P_t^L - P_{t-1}^C)$ is added to the range. In the event $P_t^H < P_{t-1}^C$, the difference $(P_{t-1}^C - P_t^H)$ is added to the range. This adjustment is aimed at controlling for the patterns of overnight price jumps that will not otherwise be reflected in measures of intraday cash price variability. The third measure of volatility, one that captures the interday variability of cash prices, is the commonly employed proxy for variability in price–volume studies: the absolute change in daily closing prices, $|P_t^C - P_{t-1}^C|$.[6] The measure of the rate of change in market prices is $|P_t^C - P_{t-1}^C| - |P_{t-1}^C - P_{t-2}^C|$. Finally, the proxy for the options trading activity is closing volume, V_t, standardized by the open interest, OT_t.[7]

Methodology

The relationship between cash market volatility and trading activity is investigated by employing the vector autoregressive (VAR) system,

$$\sigma_t = \alpha_{1_t} + \sum_{j=1}^{k} \tau_j \sigma_{t-j} + \sum_{j=1}^{k} \beta_j \left[\frac{V_{t-j}}{OI_{t-j}} \right] + \epsilon_t \tag{1a}$$

$$\frac{V_t}{OI_t} = a_{1_t} + \sum_{j=1}^{k} \pi_j \left[\frac{V_{t-j}}{OI_{t-j}} \right] + \sum_{j=1}^{k} b_j \sigma_{t-j} + u_t \tag{1b}$$

where σ_t represents the various measures of volatilities and trends in the cash market; τ_j and π_j are the coefficients for the lagged regressors of the dependent variable; β_j and b_j are the coefficients for the lagged independent variable; and ϵ_t and u_t are the random error terms. Partial autocorrelation functions are employed in the selection of the optimal distribution lag structure.[8] The VAR characterization of the volatility–activity relationship is especially meaningful since several researchers [including Rogalski (1978)] are not able to identify

[6]For instance, see Ying (1966) and Tauchen and Pitts (1983).

[7]Rutledge (1979) suggests that daily patterns in volume largely reflect speculation, since hedgers' transactions comprise relatively minor proportions of daily volume. This contention is well grounded since hedging orders are generally held for more than one day, whereas there may be many speculative trades that are day-orders. However, since total option volume also varies as a function of time to expiration, the procedure used by Garcia, Leuthold, and Zapata (1986) in examining daily volume relative to the level of open interest is followed here. This standardization procedure also removes the concern of trends in the level of trading volume per se.

[8]An intuitive guide to establish the optimal length in a VAR model is to select k that results in an estimated model without significant autocorrelation. The interested reader is referred to Amemiya and Fuller (1967) for the test with distributed lag structure. Also see Sims (1982) and Litterman (1986).

unidirectional causality in price and volume for which there is no lag effect.[9]

In the framework of Granger causality, σ_t causes V_t/OI_t if the set of estimated coefficients on the lagged σ_t in eq. (1b) is statistically different from zero as a group, and the set of estimated coefficients on the lagged values of V_t/OI_t in eq. (1a) is not statistically different from zero. Two-way causality (or feedback) is said to exist between σ_t and V_t/OI_t when the independent variable regressor coefficients are significant in both the equations. The test for causality is based on an F-statistic that is computed by employing the residual sum of squares from the reduced and full-form solutions of eqs. (1a) and (1b).[10]

It is possible that models that consider only lagged values of the exogenous variables(s), such as the one presented above, will not capture a possible contemporaneous influence across intraday cash market volatility and trading activity [see Koch (1993)]. To address this concern, the study follows Koch (1993) and extends eq. (1a) and (1b) for the $P_t^H - P_t^L$ and V_t/OI_t pairing, employing a simultaneous equation approach that includes contemporaneous instrument values of the exogenous variable.[11] The t-statistic from the two-stage least squares (2SLS) estimation will test the null hypothesis that the sum of the lagged coefficients in each regressor group is zero. Since the framework treats intraday volatility as well as daily option trading as endogenous variables, the coefficients on the contemporaneous regressors are not interpreted.[12] The results from this model represent a test for robustness of the VAR treatment of the intraday volatility–option trading relationship.

EMPIRICAL RESULTS

Table I presents the summary statistics for the measures of S&P 100 cash volatility and S&P 100 options trading. All the series with the exception of P_t^C are found to be stationary, as indicated by the statistically significant and negative augmented Dickey Fuller (ADF)

[9]It should be noted that the VAR model is strictly appropriate only when the variables are stationary. For nonstationary variables, they may be valid only approximately, or not at all [see Charemza and Deadman (1992)].

[10]For instance, the F-statistic pertaining to the null hypothesis that V_t/OI_t does not cause σ_t is given by $[(SSE_{r(1a)} - SSE_{f(1a)})/m] \div [SSE_{f(1a)}/(T - 2m - 1)]$ where $SSE_{r(1a)}$ and $SSE_{f(1a)}$ are residual sum of squares from the reduced and full form solutions of eq. (1a), T is the number of observations, and m is the number of lags.

[11]For an excellent description of the 2SLS methodology, see Kmenta (1986).

[12]For instance, the level of intraday volatility and daily option trading may reflect responses to a third factor (for instance, the arrival of information), rather than to each other.

TABLE I[a,b,c]

S&P 100 Cash Prices and OEX Trading Activity Summary Statistics (2/1984–11/1993)

	Mean	SD	Partial Autocorrelations (Lag) 1	2	3	4	5	ADF	N
P_t^C	287.080	79.783	0.998**	0.006	0.028	0.016	0.008	1.684	2453
$P_t^H - P_t^L$	3.4183	2.479	0.4625**	0.4142**	0.3661**	0.3383**	0.3397**	−9.767**	2453
$ADJR_t$	3.4175	2.481	0.4556**	0.4120**	0.3648**	0.3648**	0.3368**	−9.727**	2453
$\lvert P_t^C - P_{t-1}^C\rvert$	1.9794	2.308	0.226**	0.145*	0.137*	0.060	0.062	−19.364**	2453
$\lvert P_t^C - P_{t-1}^C\rvert - \lvert P_{t-1}^C - P_{t-2}^C\rvert$	0.00034	2.870	0.527**	−0.023	0.108*	0.085*	0.069	−13.492**	2542
V_t/OI_t	0.74×10^{-6}	0.36×10^{-6}	0.4720**	0.3433**	0.2512**	0.2134**	0.1685**	−13.92**	2453

[a] P_t^C represents the day t closing price of the S&P 100 Index; P_t^H and P_t^L represent the high and low price of the S&P 100 Index at day t. $ADJR_t$ represents the adjusted range. V_t and IO_t represent the total volume and open interest on the S&P 100 index options for day t.

[b] SD and N represent the standard deviation and number of observations, respectively.

[c] ADF denotes the augmented Dickey-Fuller test statistic for nonstationarity [Dickey and Fuller (1981) and Engle and Granger (1981)]. In the equation

$$\Delta X_t = a_o + a_1 X_{t-1} + \sum_{j=1}^{m} b_j \Delta X_{t-j} + \epsilon_t$$

the null hypothesis that x_t is nonstationary is rejected if the estimate of a_1 is negative and significantly different from zero. Critical values are from Engel and Yoo (1987). The optimal lag length in the ADF is determined from the Akaike's (1973) final prediction error (FPE) criterion.

*,**Significance levels of 5% and 1%, respectively.

statistics. There is evidence of autocorrelation in these series. The partial autocorrelation coefficients are significant up to 5 lags for the series, V_t/OI_t, $P_t^H - P_t^L$ and $ADJR_t$, and significant up to 3 lags for $|P_t^C - P_{t-1}^C|$, and 4 lags for $|P_t^C - P_{t-1}^C| - |P_{t-1}^C - P_{t-2}^C|$. Subsequently, prior to undertaking the VAR estimation, distributed lag structures for the latter two series are estimated employing eqs. (1a) and (1b). The results from the symmetric lag structure of V_t/OI_t, σ_{t-j} $(i = j = 5)$ and a distributed lag structure, V_{t-i}/OI_{t-i}, σ_{t-j} $(i = 4, j = 5)$, are found to leave the implications unchanged, so that only results from the VAR employing the symmetric lag structure, $(i = j = 5)$, are reported.

Table II presents the results from the VAR specifications. In each of the specifications, the significance of the lagged dependent regressor coefficient confirm the autoregressive patterns in both σ_t and V_t/OI_t. The results pertaining to the two intraday volatility measures ($P_t^H - P_t^L$ and $ADJR_t$) series are similar. The lagged $(t = -1, -5)$ V_tOI_t coefficients are significantly negative, indicating that greater trading on the prior day and prior week compresses the intraday volatility, as proxied by the range and adjusted range variables. On the other hand, there is some evidence that intraday volatility has a positive influence on trading activity. The lagged $(t = -1)$ coefficients for $P_t^H - P_t^L$ and $ADJR_t$ are significantly positive. The Granger causality test statistics further support the notion that there is a bicausal relationship between intraday volatility and trading activity. The null hypotheses that intraday volatility does not cause V_t/OI_t and vice versa are rejected by the F-statistics from the VAR system.[13]

The VAR results pertaining to the interday volatility measure $|P_t^C - P_{t-1}^C|$, also suggest that an increase in options trading dampens the volatility in cash prices. The lagged $(t = -1, -5)$ V_t/OI_t coefficients are significantly negative. Moreover, consistent with the findings for the intraday volatility measures, the coefficient for the lagged $(t = -1)$ $|P_t^C - P_{t-1}^C|$ variable in the equation where V_t/OI_t is the dependent variable, is positive and significant. However, the lagged $(t = -2, -3, -4)$ $|P_t^C - P_{t-1}^C|$ coefficients are negative. Thus, the results only partially support the notion that interday cash market variability is followed by an increase in options trading. Nonetheless, the F-values

[13]Since the S&P 100 index has risen during the sample period, it is possible that trends in the intraday volatility variable may provide biased results. Therefore, the robustness of the VAR results pertaining to this variable are tested separately. An adjusted intraday volatility measure, $(P_t^H - P_t^L)/\frac{1}{2}(P_t^H - P_t^L)$, is formulated and employed in the VAR specification. The results from this specification are consistent with the results from the specification that employs the unadjusted measure for intraday volatility. These results are available from the author.

continue to reject the null hypotheses that $|P_t^C - P_{t-1}^C|$ is not Granger caused by V_t/OI_t and V_t/OI_t is not Granger caused by $|P_t^C - P_{t-1}^C|$.

The final set of VAR results provide evidence to indicate that the rate of change in cash market prices is compressed by options trading. The lagged ($t = -1, -5$) V_t/OI_t coefficients are significantly negative, indicating an inverse relationship between options trading and the rate of change in cash prices, $|P_t^C - P_{t-1}^C| - |P_{t-1}^C - P_{t-2}^C|$. On the other hand, the rate of change in cash market prices is found to positively impact options trading. The lagged ($t = -1, -2, -3$) $|P_t^C - P_{t-1}^C| - |P_{t-1}^C - P_{t-2}^C|$ coefficients are significantly positive. The Granger F-statistics continue to support a strong feedback effect between cash market trends and option trading.

Table III presents the results from the 2SLS estimation of the relationship between intraday cash market volatility and option trading. For the purpose of comparison, coefficients from the ordinary least square (OLS) estimation are also provided. As in the VAR estimation, the models include five lags of the relevant regressors.[14] To test the significance of the lagged impact of options trading (intraday volatility) on the intraday volatility (options trading), t-statistics are computed for the cumulative lagged regressors. These results are presented near the bottom of Table III. The results indicate that option trading has a negative impact on intraday volatility (the cumulative lagged coefficients are negative at the 1% level of significance). A similar conclusion is reached from the OLS results. On the other hand, the evidence on the impact of lagged intraday volatility on trading is mixed. The OLS coefficients are negative and significant, whereas the 2SLS coefficients are not significant. In sum, the results in Table III support the earlier indication that a rise in option trading is followed by a decline in intraday volatility.

Further insight into the cash price versus option trading relationships is provided by simulating the cash market variability and options trading responses to innovations in trading and cash price variability from the estimated VAR system. Figures 1A, 2A, 3A, and 4A present the patterns of dynamic responses of option trading, V_t/OI_t, to 1-standard deviation (SD) shocks in the two measures of intraday volatility, the measure of interday volatility, and the measure of rate of change in cash market prices, respectively. Alternatively, Figures 1B, 2B, 3B, and 4B present the responses of the intraday and interday volatility measures and the measure of the rate of change in cash market prices to 1-SD

[14]The results are robust to different lag lengths.

TABLE II
VAR Results: S&P 100 Cash Price Volatility versus OEX Trading[a]

Dependent Series	Independent Series	Lag	Lagged Volatility	t	Lagged Activity	t	DW
$P_t^H - P_t^L$	V_t/OI_t	−1	0.331***	14.221	−2.686***	−5.550	2.003
		−2	0.164***	6.782	0.315	0.611	
		−3	0.071***	2.904	0.973	0.141	
		−4	0.035	1.409	0.206	0.399	
		−5	0.101***	4.150	−0.968*	−1.887	
V_t/OI_t	$P_t^H - P_t^L$	−1	0.003***	2.988	0.288***	12.324	2.000
		−2	−0.002	−1.493	0.094***	3.794	
		−3	−0.007	−1.411	0.029	1.181	
		−4	−0.001	−0.955	0.052**	2.116	
		−5	−0.001	−0.449	0.046*	1.877	

Ho: $(P_t^H - P_t^L)$ is not Granger caused by V_t/OI_t: F-value = 6.7027***
Ho: V_t/OI_t is not Granger caused by $(P_t^H - P_t^L)$: F-value = 3.0698***

Dependent Series	Independent Series	Lag	Lagged Volatility	t	Lagged Activity	t	DW
$ADJR_t$	V_t/OI_t	−1	0.318***	13.703	−2.514***	−5.179	2.004
		−2	0.169***	7.012	0.169	0.328	
		−3	0.074***	3.058	0.041	0.079	
		−4	0.037	1.506	0.219	0.425	
		−5	0.099***	4.102	−0.982*	−1.910	
V_t/OI_t	ADJR_t	−1	0.003***	2.680	0.292***	12.499	2.000
		−2	−0.002	−1.359	0.091***	3.679	
		−3	−0.002	−1.342	0.029	1.171	
		−4	−0.001	−0.886	0.052**	2.090	
		−5	−0.000	−0.428	0.046*	1.875	

Ho: $ADJR_t$ is not Granger caused by V_t/OI_t: F-value = 6.1723***
Ho: V_t/OI_t is not Granger caused by $ADJR_t$: F-value = 2.7786***

continued

TABLE II (Continued)

Dependent Series	*Independent Series*	*Lag*	*Lagged Volatility*	t	*Lagged Activity*	t	*DW*
$\lvert P_t^C - P_{t-1}^C \rvert$	V_t/OI_t	−1	0.176***	8.075	−1.710***	−3.554	1.996
		−2	0.109***	4.972	−0.360	−0.715	
		−3	0.104***	4.684	0.135	0.267	
		−4	0.024	1.072	−0.005	−0.009	
		−5	0.124***	5.620	−1.083**	−2.154	
V_t/OI_t	$\lvert P_t^C - P_{t-1}^C \rvert$	−1	0.006***	6.000	0.275***	12.582	2.002
		−2	−0.002**	−2.374	0.105***	4.605	
		−3	−0.002**	−2.114	0.033	1.468	
		−4	−0.002**	−2.032	0.052**	2.248	
		−5	−0.000	−0.282	0.044*	1.934	
Ho: $\lvert P_t^C - P_{t-1}^C \rvert$ is not Granger caused by V_t/OI_t: F-value = 4.1707***							
Ho: V_t/OI_t is not Granger caused by $\lvert P_t^C - P_{t-1}^C \rvert$: F-value = 8.4180***							
$\lvert P_t^C - P_{t-1}^C \rvert - \lvert P_{t-1}^C - P_{t-2}^C \rvert$	V_t/OI_t	−1	−0.756***	−34.489	−2.122***	−4.227	2.005
		−2	−0.581***	−20.654	−0.624	−1.205	
		−3	−0.414***	−13.662	−0.120	−0.230	
		−4	−0.330***	−10.882	−0.344	−0.660	
		−5	−0.139***	4.973	−1.509***	−2.918	
V_t/OI_t	$\lvert P_t^C - P_{t-1}^C \rvert - \lvert P_{t-1}^C - P_{t-2}^C \rvert$	−1	0.007***	6.773	0.271***	12.345	1.999
		−2	0.005***	3.855	0.103***	4.499	
		−3	0.003**	2.362	0.031	1.363	
		−4	0.002	1.264	0.048**	2.064	
		−5	0.002	1.643	0.040*	1.760	
Ho: $\lvert P_t^C - P_{t-1}^C \rvert - \lvert P_{t-1}^C - P_{t-2}^C \rvert$ is not Granger caused by V_t/OI_t: F-value = 8.2173***							
Ho: V_t/OI_t is not Granger caused by $\lvert P_t^C - P_{t-1}^C \rvert - \lvert P_{t-1}^C - P_{t-2}^C \rvert$: F-value = 7.8879***							

[a] t and DW represent the t- and Durbin Watson statistics, respectively.

*,**,***Significance levels of 10%, 5%, and 1%, respectively.

TABLE III
Estimation of the Intraday Volatility–Option Trading Relationship with Contemporaneous Variables[a,b]

Independent Variables	Dependent Variable			
	$P^H - P^L$		V_t/OI_t	
	OLS	*2SLS*	*OLS*	*2SLS*
Constant	−0.115	−1.171	0.124***	0.151
$P^H - P^L$				
lag				
0			0.024***	0.007
1	0.296***	0.274***	−0.005***	0.001
2	0.183***	0.193***	−0.006***	−0.003
3	0.089***	0.099***	−0.003***	−0.002
4	0.046**	0.053**	−0.002**	−0.001
5	0.105***	0.109***	−0.003***	−0.001
V_t/OI_t				
lag				
0	10.413***	16.958***		
1	−5.683	−7.574**	0.353***	0.307***
2	−0.667	−1.277	0.086***	0.092***
3	−0.234	−0.432	0.027	0.029
4	−0.342	−0.682	0.047**	0.051**
5	−1.452***	−1.760**	0.069***	0.053**
The sum of lagged coefficients				
$P^H - P^L$			−0.019***	−0.006
V_t/OI_t	−8.378***	−11.725***		
Adjusted R^2	0.484	0.415	0.361	0.250
DW	2.000	2.000	2.000	2.000

[a]DW is the Durbin Watson Statistic.
[b]The significance levels are based on t-statistics for Ho: $\Sigma_i x_i = 0$, where X_i ($i = 1 \ldots 5$) represent the lagged coefficients of the independent variable.
*,**,***Significance levels of 1%, 5%, and 10%, respectively.

shocks in trading. In each of the figures, the middle function represents the estimated response, while the two bounding lines represent the upper and lower limits of the responses with 2 standard errors.[15]

In Figure 1A the response of V_t/OI_t to shocks in $P_t^H - P_t^L$ is found to be positive for days 1–6. This would indicate that options trading rises with an increase in intraday cash market volatility. In Figure 1B, the response of $P_t^H - P_t^L$ to innovations in V_t/OI_t is generally negative for the 16-day time period. This would indicate that intraday cash market volatility drops following innovations in options trading. Similar patterns are found in the impulse responses pertaining to trading and the

[15]The impulse responses, with upper and lower 2-standard error bounds, are obtained via a Monte Carlo simulation program.

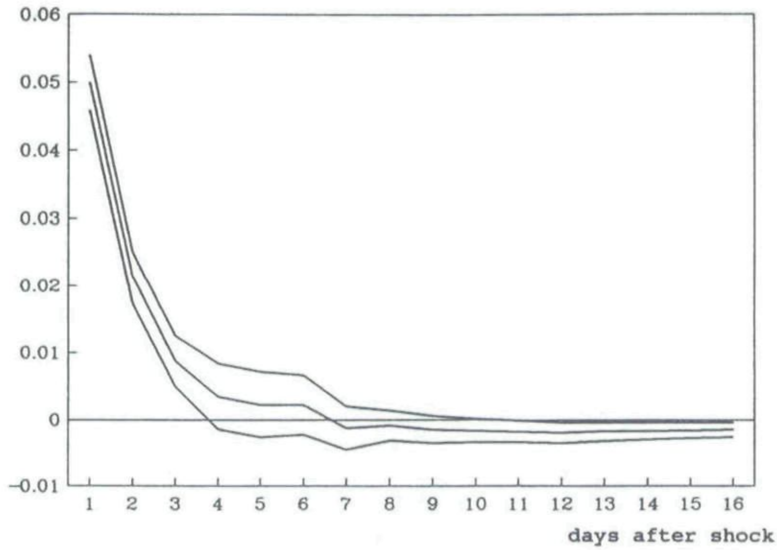

FIGURE 1A
Impulse response of option trading (V_t/OI_t) to 1-standard deviation shock in $P^H - P^L$.

second measure of intraday volatility, $ADJR_t$. The positive responses of trading to innovations in $ADJR_t$ is found to endure until day 6 (Fig. 2A). On the other hand, there is evidence to indicate that $ADJR_t$ drops as a response to innovations in V_t/OI_t. Figures 3A and 3B pertain to the impulse responses of options trading and the measure of interday

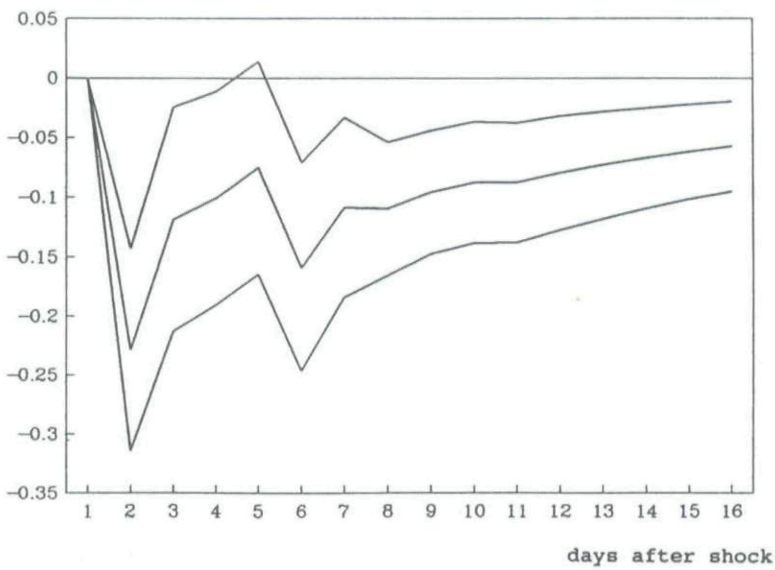

FIGURE 1B
Impulse response of $P^H - P^L$ to 1-standard deviation shock in option trading.

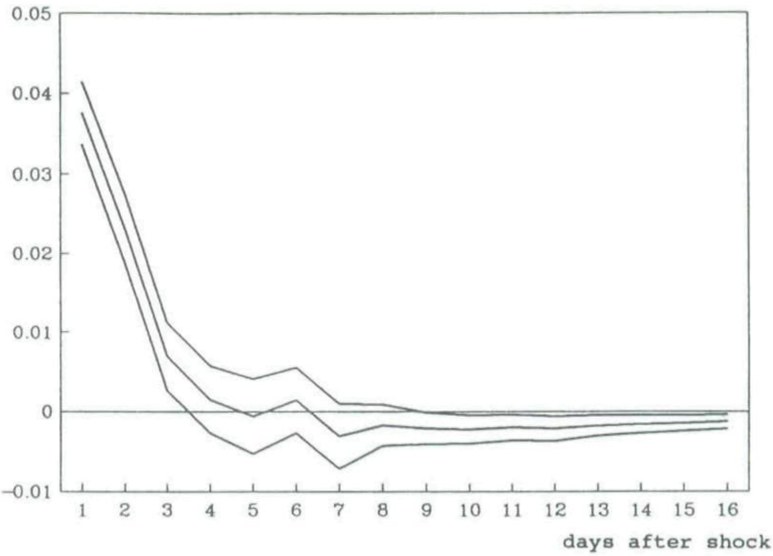

FIGURE 2A
Impulse response of option trading (V_t/OI_t) to 1-standard deviation shock in *ADJR*.

volatility, $|P_t^C - P_{t-1}^C|$. Once more, there is evidence that options trading increases with shocks in price variability (Fig. 3A). Moreover, as before, cash price volatility is found to decline with positive shocks in options trading (Fig. 3B). The three sets of impulse responses thus indicate that while an increase in options trading depresses subsequent

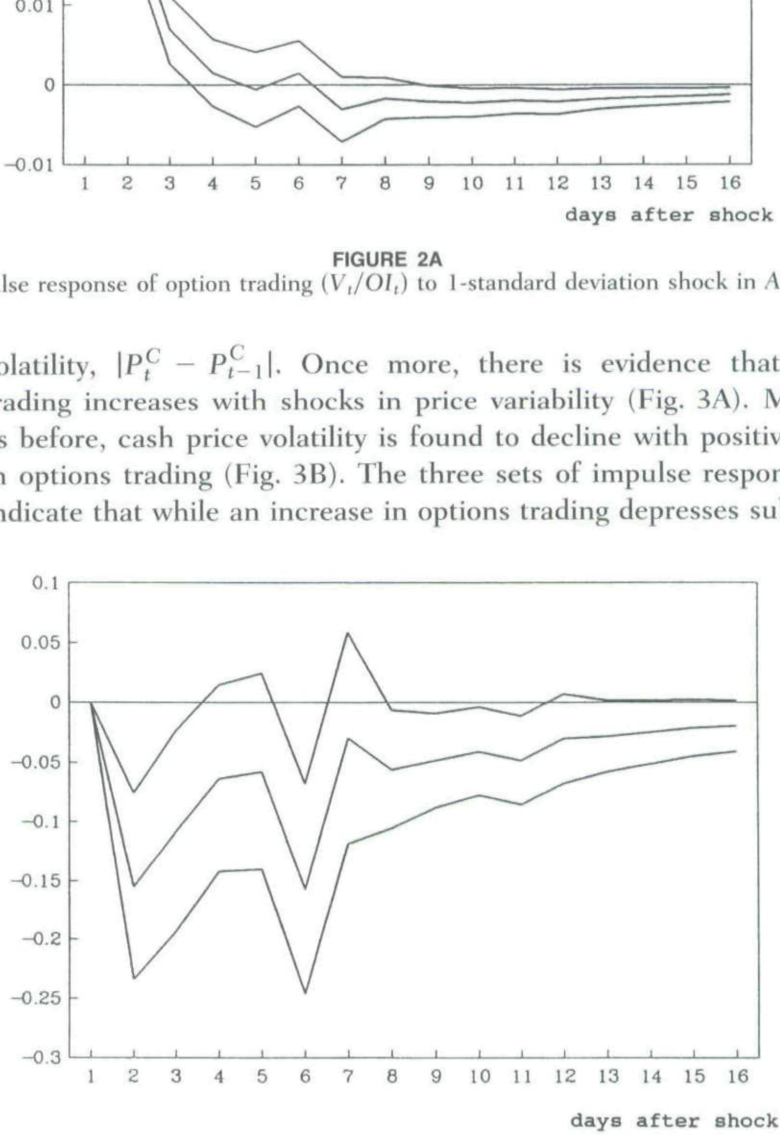

FIGURE 2B
Impulse response of *ADJR* to 1-standard deviation shock in option trading.

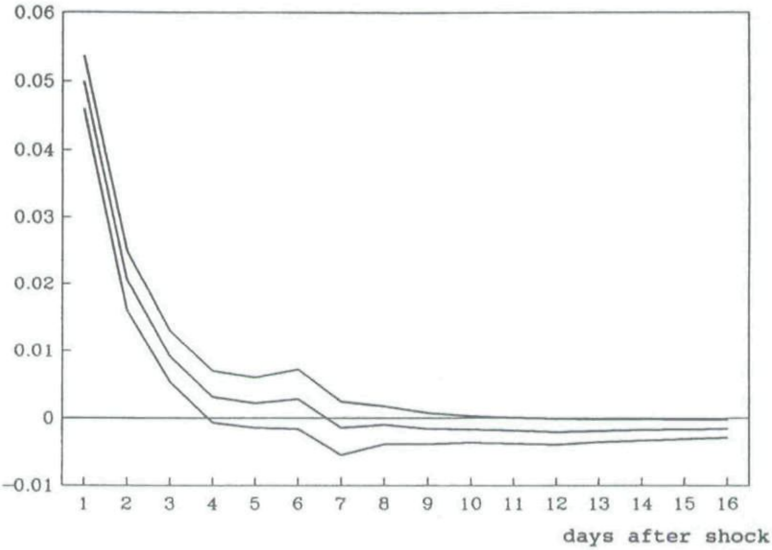

FIGURE 3A
Impulses responses of option trading (V_t/OI_t) to1-standard deviation shock in $|P_t^C - P_{t-1}^C|$.

cash price volatility, an increase in case price volatility has a positive impact on options trading. Finally, the impulse response in Figure 4A suggests that shocks in the rate of change in cash markets have a positive impact on options trading. On the other hand, the impulse

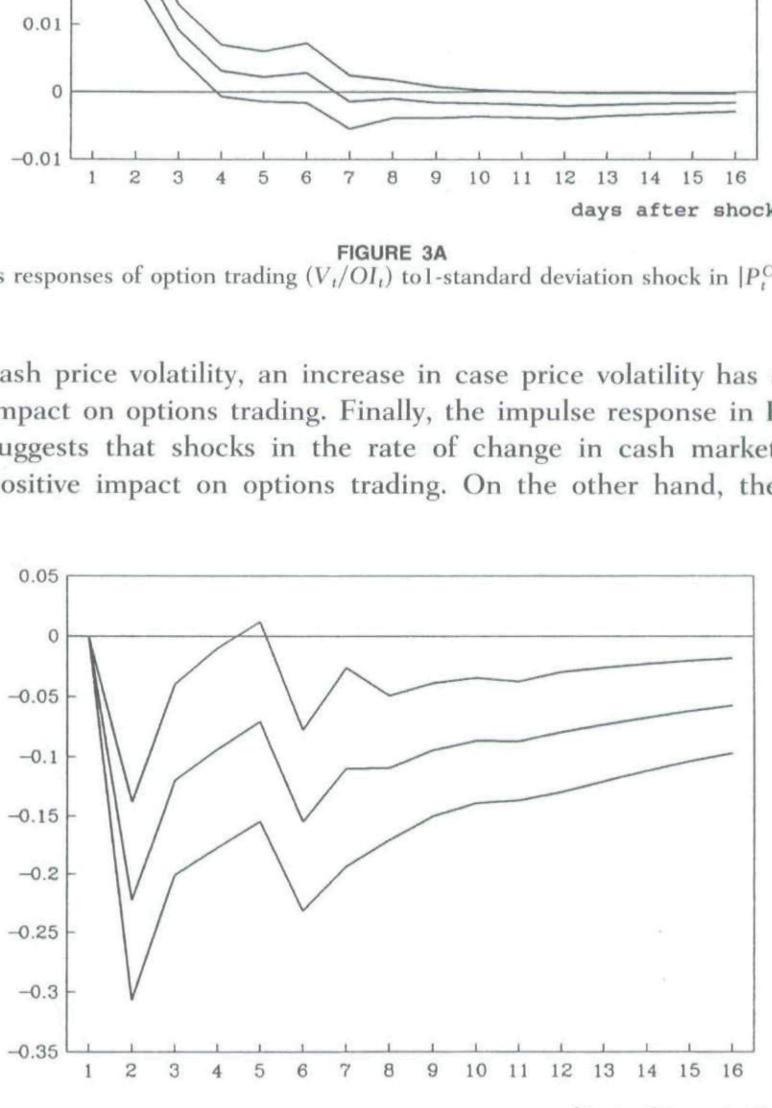

FIGURE 3B
Impulse response of $|P_t^C - P_{t-1}^C|$ to 1-standard deviation shock in option trading.

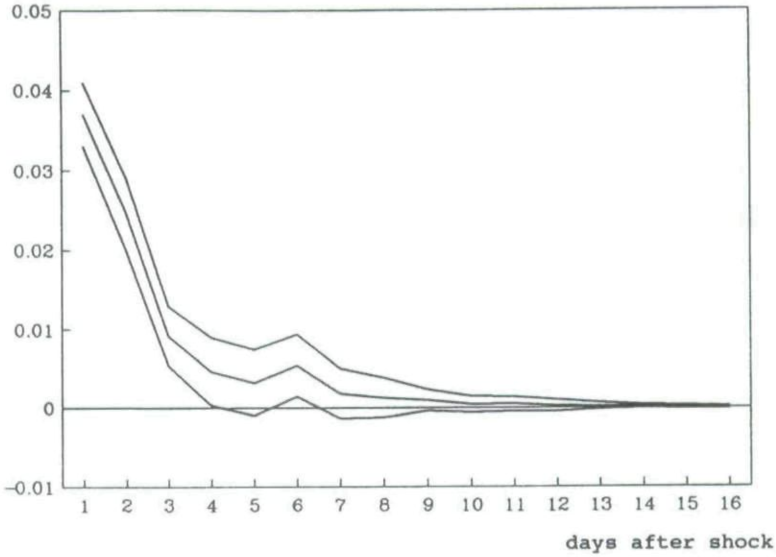

FIGURE 4A
Impulse response of option trading (V_t/IO_t) to 1-standard deviation shock in $|P_t^C - P_{t-1}^C| - |P_{t-1}^C - P_{t-2}^C|$.

response in Figure 4B provides little indication to support the hypothesis that options trading exacerbates the rate in cash market prices. The impulse responses of this variable have mixed signs, with the lower limit remaining negative for the most part.

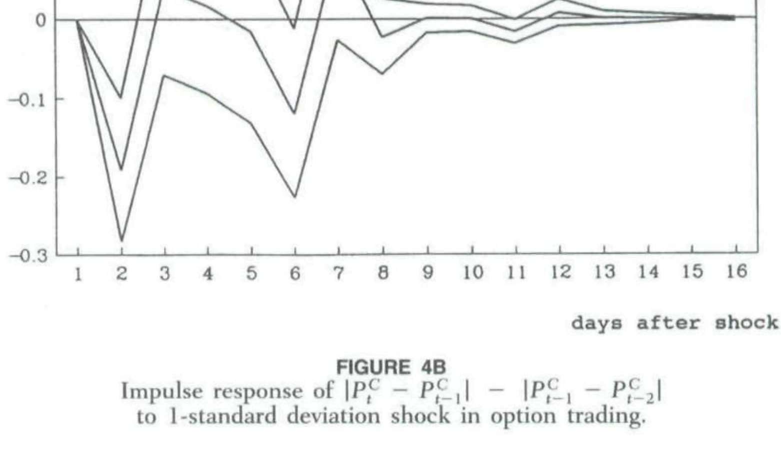

FIGURE 4B
Impulse response of $|P_t^C - P_{t-1}^C| - |P_{t-1}^C - P_{t-2}^C|$ to 1-standard deviation shock in option trading.

SUMMARY AND CONCLUSIONS

This study addresses two questions, namely, whether options trading leads to an increase/decline in interday and intraday cash market volatility, and whether options trading exacerbates/compresses the rate of change in cash market prices. These questions are addressed with respect to daily activity in the S&P 100 option index and the trends in the corresponding cash index over the period, February, 1984 to November, 1993. The results indicate that an increase in the intraday and interday volatility in the case market is followed by an increase in the level of options trading, and an increase in options trading is followed by a decline in cash market variability. This evidence is consistent with prior studies that provide evidence of lower cash market volatility after the introduction of option trading. Furthermore, the study finds no evidence to indicate that options trading intensifies the rate of change of cash market prices. The overall evidence is consistent with the notion that options trading has a stabilizing influence on cash markets.

An extension to the framework employed in this study could further the understanding of the option trading–cash market volatility relationship. The potential impact of options trading may be evaluated alongside (after controlling for) any futures trading in the underlying asset under investigation. This is possible for several indexes (e.g., S&P 500) and commodities (e.g., currencies) for which both futures and options trading are available. An investigation that employs options, futures and cash market data could shed light on the lead–lag relationships between futures and options markets, and differentiate between the impact of these markets on the underlying assets.

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