An Analysis of the Linkage between Japanese Stock Market and the Derivatives Markets : Estimates of the Implied Stock Prices*

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1 INTRODUCTION

Since January 4, 1990, the *Nikkei 225* stock index which peaked at the end of 1989 has continued to have a declining trend. Despite outside factors including the qualitative change in the state of the Japanese economy and the effects of the exchange and interest rate market, there is another important reason for this dramatic change in Japanese stock market, that is, 1990 was the first year that new derivative goods (for example, futures and options) began to operate. Can it be said that the derivatives prices influence the cash stock index ?

In general, the investors in the new derivative markets can bring some informational contents of prices in the markets in which their operate, in addition to providing increased risk sharing. It has been shown that the entry of the new information has a positive or negative effect : the spot price can become more or less informative to those traders already in the market (see

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Grossman, 1988 and Stein, 1987). This effect suggests that the spot index is influenced by the expected information of the derivative investors and that the index can become an assessment value of the investors if the markets are efficient.

Motivated by these considerations, this paper examines the long-run relationship between the spot price and the equilibrium prices of the underlying assets in the Japanese derivatives markets using three variables: an actual stock index, an implied stock index from a futures instrument, and an implied stock index from an option instrument. The implied stock index (*SC*) is directly calculated from futures prices. The implied stock index from option prices (*ISP*) is estimated by combining with the implied standard deviations (*ISD*) in accordance with the Black-Scholes model (1973) (B-S model). Given observed option prices and values for several parameters except for the underlying asset price in the B-S model, the implied stock index that equates the observed option prices to their calculated values can be determined with the implied standard deviation.

If these derivative prices actually depend on the models, the implied values will sensitively reflect the expected information in the markets, that is, the *SC* and *ISP* will reflect the investors assessments of the equilibrium prices. Manaster and Rendleman (1982) show that the implied value reflects the equilibrium stock prices and contains information about stock returns on the following day, using data for the United States. Asano (1993) also examines the implied stock prices calculated by using the equilibrium put-call parity in the Japanese option market. Using these implied stock prices, serial correlations are shown to occur over short time periods. Former empirical studies do not compare the three markets: the cash stock, futures, and option markets, nor the long-run movements of the three indices: the actual stock index, the implied stock index from the futures market, and the implied stock index from the option market, especially using time series analysis.

Here, two main problems are examined. First, can the actual stock index reflect the equilibrium stock value in the same way as *SC* or *ISP*? Second, are there any arbitrage opportunities among the three markets? In order to investigate these two questions, the comovement among the three variables is examined statistically using tests for unit-roots and cointegration. As the three variables might reflect the same information about the equilibrium value of the underlying stock index, it is likely that cointegrating relationships will be observed. Further, if these variables are cointegrated at different levels, there will be some arbitrage opportunities and a basis that will be stationary.

The analysis uses time series data on the *Nikkei225* from February 1994 when the *Nikkei300* option market opened until December 1999[1]. There appear to be cointegrating relationships among the observed index, implied index from futures, and the implied index from option during the sample period, which shows a linkage among the three markets: the cash, futures, and option markets. It is also found that between the cash and the each derivative market positive or negative basis exists, which is an implication of the arbitrage conditions.

The paper is organized as follows. Section 2 contains details of the models that are used to calculate the implied futures and option indices. The time series behavior of the observed stock prices, the calculated index for the futures, and the implied index for the option is discussed in section 3. Section 4 contains a short conclusion.

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1) The earlier version of this paper (Kii, 1994) used the data from November 1989 to November 1992.

2 THEORETICAL FRAMEWORK

2.1 Estimation of Implied Stock Indices

2.1.1 Futures Market

The best-known model for pricing stock index futures is the cost of carry model. Its derivation relies on a simple no-arbitrage argument in which a trader replicates a futures position with spot positions in the stock market (see Modest and Sundaresan, 1983)[2].

This model implies

$$F_t = e^{r_t T_t} S_t \tag{1}$$

where F_t is the futures price at time t for an instrument, S_t is the current value of the underlying stock at time t, r_t is the risk-free interest rate, and T_t is the time to maturity of the futures.

The implied stock index for futures valuations (SC) is obtained by inverting (1) to give

$$SC_t = e^{-r_t T_t} F_t. \tag{2}$$

2.1.2 Option Market

Black and Scholes (1973) derived an option-pricing model (the B-S model) that was extremely useful in catalyzing much of the research on option-like financial instruments. This formula is

$$c = SN(d_1) - Ke^{-r\tau} N(d_2), \tag{3}$$

$$p = Ke^{-r\tau} N(-d_2) - SN(-d_1) \tag{4}$$

where c is the current call option price, p is the current put option price , S is the current price of the underlying stock, r is the risk-free interest rate, σ is the standard deviation of the stock's rate of return (volatility), K is the exercise

2) This model is originally used as a forward pricing model. To apply it to stock index futures, it is necessary to assume that forward and futures prices are equal.

price, τ is the time to maturity of the option, and N (●) is the standard normal cumulative density function. The variables d_1 and d_2 are defined as

$$d_1=\frac{\ln(S/K)+r\tau}{\sigma\sqrt{\tau}}+\frac{\sigma\sqrt{\tau}}{2},$$

$$d_2=d_1-\sigma\sqrt{\tau}$$

In this model, the variance of the stock's rate of return (volatility) is the only unknown and the option value is highly sensitive to estimates of this volatility. According to this model, the implied standard deviation (*ISD*) and the implied stock index (*ISP*) are estimated simultaneously using data from several options on the same stock, while at the same time avoiding difficulties associated with errors of measurement in the standard deviation.

The implied stock index for futures (*SC*) is calculated directly from (2). However, the implied stock index for options (*ISP*) is obtained by applying numerical techniques to (3) and (4). In this study, the estimates of *ISD* and *ISP* are computed by applying nonlinear least square to option prices in the following models,

$$c_t=\Psi_{ct}(\hat{S},\ \hat{\sigma};K,\ r,\ \tau)+\varepsilon_{ct} \tag{5}$$

$$p_t=\Psi_{pt}(\hat{S},\ \hat{\sigma};K,\ r,\ \tau)+\varepsilon_{pt} \tag{6}$$

where c_t (p_t) represents the observed price of the call (put) option at time t, Ψ_{ct} (●) (Ψ_{pt} (●)) is the model's price, ε_{ct} (ε_{pt}) is a random disturbance, and $\hat{S}$ and $\hat{\sigma}$ represent *ISP* and *ISD*, respectively. Estimates of S_t and σ_t are obtained by solving the following minimization problem (see Whaley, 1981):

$$\min_{\hat{S}_t,\hat{\sigma}_t} G_t=\sum_{i=1}^{N_{ct}}(c_{it}-\Psi_{ict}(\bullet))^2+\sum_{j=1}^{N_{pt}}(p_{jt}-\Psi_{jpt}(\bullet))^2 \tag{7}$$

where N_{ct} and N_{pt} represent the number of call and put options at time t, respectively. The solution to (7) minimizes the sum of the squared deviations between the observed and calculated option prices. The estimates of *ISP* and

ISD are obtained using two methods. First, the initial value of both S_t and σ_t can be estimated using the multidimensional golden section method (*GSS*)(see Iwata, 1989 and Gallant, 1987), then, *ISP* and *ISD* are calculated at the same time using a Gauss search method[3].

In the next section the time-series behavior of the calculated values of the implied stock data (SC_t, ISP_t) and the observed stock prices (S_t) are examined.

2.2 Cointegration among the Stock Indices

The cash index implied by the futures or option contract is the value of the underlying stock for which a continuously revised futures or option portfolio would be a perfect substitute for the stock. However, the actual stock price cannot be adjusted simultaneously to make the index individually to arrange the portfolio. Hence, it can be said that there are three concurrent indices for an underlying stock value, that is, the observed cash index (S), the index implied by the futures price (SC), and the index implied by the option price (ISP).

If the actual stock price reflects the implicit equilibrium value, S, SC, and ISP move simultaneously and there is a linkage among the three markets, the cash, futures, and option markets. In this case, S and SC (or S and ISP) will be cointegrated such that $S-SC$ (or ($S-ISP$)) is stationary. Moreover, if the level of the implied stock index (SC or ISP) differs from that of the observed index (S), some arbitrage opportunities will exist between the cash stock and the derivative instruments. Even then, it is possible that S and SC or S and ISP move together because market traders will attempt to profit from such price differences, that is, a basis. This implies that S and SC or S and ISP

3) The error tolerance values of the two methods are 1.0×10^{-8}.

are cointegrated such that the basis is stationary. In contrast, if the general level of S is consistent with the level of both SC and ISP when S, SC, and ISP are cointegrated, there is no arbitrage condition among the three markets.

Suppose that $\boldsymbol{X}'_t = (S_t,\ SC_t,\ ISP_t)$ and each variable in $\boldsymbol{X}_t$ contains a unit-root. The previous arguments suggest that the observed stock prices (S_t) might be cointegrated with the calculated prices (SC_t, ISP_t) with cointegration vectors $(1, 0, -1)'$ and $(1, -1, 0)'$. For example, the spread between S_t and SC_t or S_t and ISP_t might be a stationary linear combination of $\boldsymbol{X}_t$.

3 EMPIRICAL TESTS

3.1 Data

All of the date, including measures of the all monthly maturities and all exercise prices of *Nikkei225* index options at time t (c_t, p_t), the all monthly maturity *Nikkei225* index futures at time t (F_t), and the yield of *CD gensaki* (*bit*) at time t (r_t) are taken from the *Nikkei Quick Data*. The data are all daily closing values, and r_t, T_t, and τ_t are per annual values. The data ovar 1439 trading days runs from February 14, 1994 to December 30, 1999.

At February 14 in 1994, the new option market, *Nikkei300* option market, was open. In order to examine an effect of the new derivative market, the implied volatilities (*ISD*) are estimated before and after the introduction of the new market and are compared with each other[4]. The period before is from January 7, 1993 to February 10, 1994 and the period after is from February 14, 1994 to December 30, 1994. **Table 1** and **Figure 1** show the summary statistics and the movements of *ISD*. They show that *ISD* is smaller after the introduction of the *Nikkei300* option market, which indicates a possibility

4) A detailed analysis of the effect of the new derivative market to the spot price volatility is left for a future study.

Table 1: Summary Statistics for *ISD*

	Before	*After*
MEAN	0.096931	0.064543
S.D.	0.030465	0.024648
MIN	0.035384	0.00001
MAX	0.19942	0.14639

Note: MEAN, S.D., MIN, and MAX denote the sample mean, the standard deviation, the minimum value, and the maximum value of the data, respectively. *Before* indicates the period from 7/1/93 to 10/2/94 and *After* indicates the period from 14/2/94 to 30/12/94.

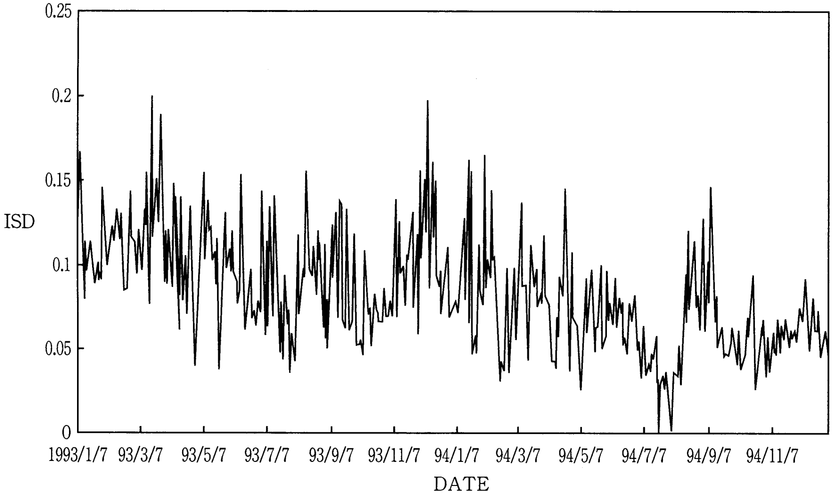

Figure 1

Time Series Data; The Implied Volatilities for Options
(7/1/1993-30/12/1994)
—ISD

that the new derivative market has a role of increased risk sharing if the other state of the Japanese economy does not change before the establishment of the market and behind. Next, the actual stock index that will be influenced by the derivatives is investigated with the implied stock prices from the derivatives markets after the introduction of the new market.

3.2 Linkage between *S*, *SC*, and *ISP*

3.2.1 A Unit-root

Some time series invstigations both of the actual spot price (*S*) and the implied stock prices (*SC*, *ISP*) are shown below. The movements of *S*, *SC*, and *ISP* are shown from **Figure 2** to **Figure 4**. The means and standard deviations of the three variables, *S*, *SC*, and *ISP*, are contained in **Table 2**. *ISP* has the smallest mean and maximum values and is more correlated with *S* rather than *SC*. This result shows that the investors in the cash stock option market regard for the cash stock index as the underlying asset value not for the futures price, which is in contrast to some previous studies (for example, Harvey and Whaley, 1991).

Next, the time series properties of the data are explored using tests for a unit-root. A number of statistics have been proposed as tests for the exist

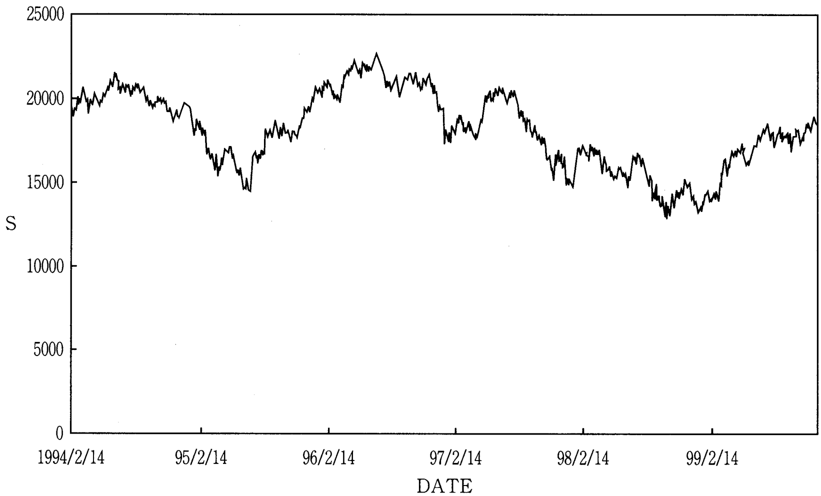

Figure 2

Time Series Data ; *Nikkei 225* Stock Prices
(14/2/1994-10/12/1999)
—S

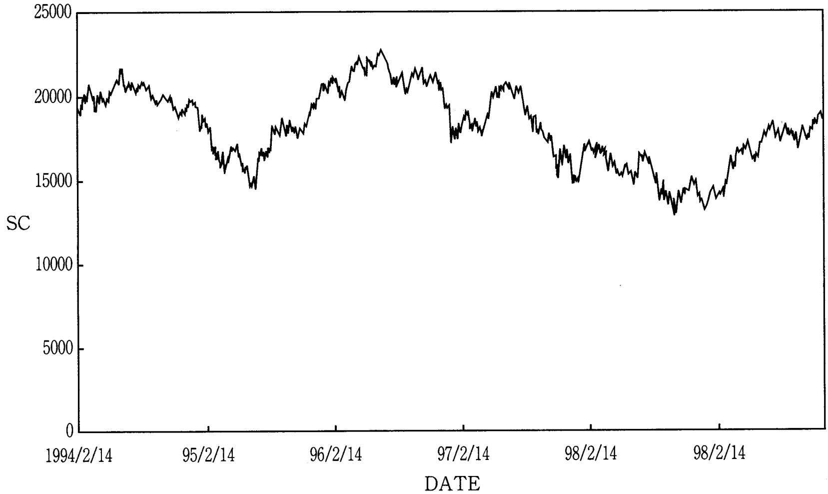

Figure 3

Time Series Data ; The Implied Stock Prices for Futures
(14/2/1994-10/12/1999)
—SC

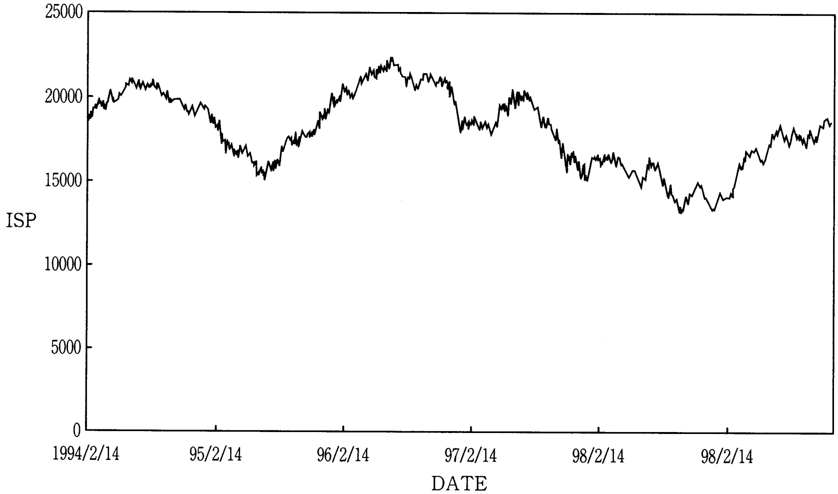

Figure 4

Time Series Data ; The Implied Stock Prices for Options
(14/2/1994-10/12/1999)
—ISP

Table 2: Summary Statistics

	S	SC	ISP
MEAN	18154.42	18162.39	18091.19
S.D.	2364.78	2380.52	2294.77
MIN	12879.97	12868.22	13130.82
MAX	22666.80	22704.08	22441.99

	S	SC	ISP
S	1.00000		
SC	0.99976	1.00000	
ISP	0.98535	0.98495	1.00000

Note: See the *Note* for Table 1.

-ence of a unit-root (see, for example, Dickey and Fuller, 1979, 1981, and Phillips, 1987). In this study, the well-known Dickey-Fuller (*DF*) and Augmented Dickey-Fuller (*ADF*) statistics are used.

An asymptotically valid method of testing for a unit-root in the process generating a variable Y is to employ the *ADF* regression with drift:

$$\Delta Y_t=\alpha_0+\alpha_1 t+\alpha_2 Y_{t-1}+\sum_{i=1}^{m}\beta_i \Delta Y_{t-i}+\eta_t$$

where the η_t is assumed to be identically and independently distributed random variables. In this experiment, four lags of the differenced variables are included ($m=4$) to account for serial correlation in the error term[5].

Table 3 reports Augmented Dickey-Fuller unit-root tests for levels of each variable, and the results of the unit-root tests for their first differences are also presented in Table 3. All statistics are consistent with the null hypothesis of a unit-root for the levels of all variables at the 5 percent significance level. In contrast, the null hypothesis that there is a unit-root is clearly rejected at the 5 percent significance level for their differences. An appropriate conclusion is that each stock variable, *S*, *SC*, and *ISP*, follows an integrated process of order one over the samle period[6].

5) The choice of m largely depends on the number of available observations and the serial correlation pattern present in η_t.

6) The hypothesis that the error term in each of their regressions is serially uncorrelated is accepted when tested using the Q statistic due to Box and Pierce (1970). Using the *ARCH* statistic proposed by Engle (1982), the hypothesis of homoscedastic errors was accepted.

Table 3: Unit-Root Tests

	S	ΔS	SC	ΔSC	ISP	ΔISP
DF	−2.0217	−40.5335*	−2.0816	−42.2631*	−1.8830	−45.6528*
ADF	−1.8120	−19.9829*	−1.8528	−29.6154*	−1.4542	−24.7850*

Note: Each ADF statistic is based on a vector autoregression with a lag length of 3 for S, 1 for SC, or 2 for ISP, respectively. The 5% critical values for the DF and ADF without trend are −2.8640 and the superscript * indicates that the statistic can reject the null at the 5% level (source: Dickey and Fuller, 1981).

3.2.2 Cointegration

Given these results, the next hypothesis of interest is to test for cointegration. Three main approaches are currently available to test for cointegration. One is Johansen's Full Information Maximum Likelihood approach developed by Johansen (1988) and Johansen and Juselius (1992), which estimates the cointegration vectors and permits testing of restrictions on the cointegrating vectors. This procedure seems to be most prevalent because it allows for potentially stationary variables in contrast with the other two methods that need non-stationarity. Another approach is the Augmented Dickey-Fuller (*ADF*) method recommended by Engle and Granger (1987), which is a test for unit-roots in the residuals from a cointegrating regression. Phillips (1991) has introduced a mixed Gaussian estimator of the cointegrated regression, which allows for some serial correlation. In this case, ordinary least squares estimates of the coefficients of the unconstrained linear equation which contains lagged and forwarded differences of the independent variables can be used to test simple null hypotheses using t statistics, which are asymptotically distributed as $N(0, 1)$. These three techniques are applied to test the hypotheses of interest.

First, the results of Johansen tests to determine the rank of the cointegration vector for the variables are presented in **Table 4**[7]. Over the sample,

Table 4: Johansen Cointegration Tests

Null	λ-max	95% Critical Value	*trance*	90% Critical Value
r=0	270.1662	21.0740	351.1065	31.5250
r≤1	77.6863	14.9000	80.9403	17.9530
r≤2	3.2541	8.1760	3.2541	8.1760
CHI−SQ(2)=37.7849*				

Note: All statistics are based on a vector autoregression with a lag length of 3, and critical values are taken from Osterwald-Lenum (1992). The statistic for the test of the null hypothesis that the two cointegration vectors between *S*, *SC*, and *ISP* are (1, −1, 0)′ and (1, 0, −1)′ is denoted by CHI−SQ(2). The superscript ∗ indicates that the statistic can reject the null at the 5% level.

Johansen's λ-max and *trace* statistics for the levels case accept the null hypothesis that the rank of the cointegration space is not more than two, but strongly reject the null hypothesis that the rank is not more than one at the 5 percent significance level. The results of the estimated cointegration vectors normalized by *S* are

$$S_t = 0.97876SC_t + 0.015230ISP_t$$

$$S_t = -3.1705SC_t + 4.2747ISP_t$$

which shows that *S* is more correlated with *SC* than *ISP*.

The restriction on the cointegrating vector

$$\begin{pmatrix} 1 & -1 & 0 \\ 1 & 0 & -1 \end{pmatrix}'$$

is strongly rejected at the 5 percent significance level with the *LR* test. This shows that the joint hypothesis that $S-SC$ and $S-ISP$ are stationary is strongly rejected in the sample period. These two results indicate that the three variables are cointegrated but there is a possibility that the difference

7) In this paper, only the results for a vector autoregression with three lags chosen with the *AIC* criteria reported but repeating the analysis with other lag lengths leads to the same conclusions.

Table 5: DF and ADF Cointegration Tests

	DF	ADF (3)
S-SC	−30.6014*	−14.2428*
S-ISP	−11.9402*	−6.5042*
SC-ISP	−12.1183*	−6.4333*

Note: See the *Note* for Table 3.

between any two variables contains a unit-root.

Second, *DF* and *ADF* tests are used to test whether individually the variables $S-SC$ and $S-ISP$ are stationary. **Table 5** contains the results of the *DF* and *ADF* tests where a lag length of 3 is used for the *ADF* test to account for serial correlation in the error term (denoted by *ADF* (3)). Both *DF* and *ADF* (3) statistics show that all pairs of variables do not have a unit-root and are stationary in the sample period, in contrast to the results of the Johansen type tests.

Third, the Phillips (1991) or Stock and Watson (1993) test is used to test for cointegration between *S*, *SC*, and *ISP* and is also used to examine whether each pair, $S-SC$, $S-ISP$ or $SC-ISP$, has comovements at exactly the same level. At first, the following equation is used to test for the three variables:

$$X_t=\alpha_0+\alpha_1 Y_t+\alpha_2 Z_t+\sum_{i=-3}^{i=+3}\beta_i \Delta Y_{t+i}+\sum_{j=-3}^{j=+3}\gamma_j \Delta Z_{t+j}+\eta_t \qquad (8)$$

where (X, Y, Z) is (S, SC, ISP) and η_t is a disturbance. The result of this estimation is

$$S_t=102.2523^*+0.97892^* SC_t+0.015105^* ISP_t,$$

where the superscript * indicates that the t statistic can reject the null hypothesis that α_0 (or α_1, α_2) is 0 at the 5% level. This result shows that *S* is more influenced by *SC* rather than *ISP*, which is similar to the result of the Johansen type test.

Table 6: Phillips Tests

	α_0	α_1	χ^2
S, SC	113.7421	0.99334	
	(10.5198)*	(1679.7)*	(110.6653)*
S, ISP	−244.1666	1.0167	
	(3.1368)*	(237.8394)*	(9.8391)*
SC, ISP	−353.2134	1.0232	
	(4.4495)*	(234.6897)*	(19.7973)*

Note: The absolute values of t statistics for testing the null hypotheses that $\alpha_0=0$, $\alpha_1=0$, and $\alpha_2=0$ are reported in parentheses. The χ^2 is used in the *Wald* tests: $\boldsymbol{H_0}$: $\alpha_0=0$, $\alpha_1=1$. The superscript * indicates that the statistic can reject the null at the 5% level.

Next, the following equation is estimated to examine $S-SC$, $S-$ISP, or $SC-ISP$ has comovements at exactly the same level.

$$X_t=\alpha_0+\alpha_1 Y_t+\sum_{i=-3}^{i=+3}\beta_i \Delta Y_{t+i}+\eta_t \tag{9}$$

where (X, Y) is (S, SC), (S, ISP), or (SC, ISP). The *Wald* test is also used to examine the hypothesis:

$$\boldsymbol{H_0}: a_0=0,\ a_1=1,$$

Table 6 contains estimates of α_0 and α_1, and the absolute values of the t statistics for all variables. The χ^2 statistics used in the *Wald* tests are also shown in the table. All statistics for all pairs of variables reject the hypotheses at the 5 percent level. In addition, the values of both $\hat{\alpha}_0$ and $(1-\hat{\alpha}_1)$ are positive for (S, SC), contrary to (S, ISP) or (SC, ISP) which has negative values for these estimates. These results suggest that there is a non-stationary basis between the variables, (S, SC), (S, ISP), or (SC, ISP), in contrast to the results of the *DF* and *ADF* tests.

4 CONCLUSION

This paper provides new evidence regarding the dynamics of stock prices, stock index futures prices, and stock index option prices in Japanese markets. In particular, the observed stock prices and the implied stock prices from the derivatives instruments are examined in a period after the new option market: *Nikkei300* option market was introduced in the Japanese financial markets.

The empirical results of this study provide three important conclusions. First, the observed value, the implied value for futures prices, and the implied value for option prices are cointegrated in the period. These results indicate that cash stock, futures, and option markets can be integrated and the actual spot price can reflect the market information of the derivatives markets even after the establishment of *Nikkei300* option market. It is also found that the spot stock market is more correlated with the futures market rather than the option market, which is consistent with a fact that the futures is regarded as the more active trade than the option.

Second, a positive basis exists between the spot stock index and the implied stock index in the futures market, suggesting that there is an arbitrage chance between the two markets. Between the cash stock index and the calculated index from options a negative basis is observed, suggesting that the actual spot index might be underestimated and that there is a possibility for arbitrage remaining. These findings are caused by some characteristic of the derivative trades. Indeed, the derivatives are regarded as the more superior trades than the cash stock because of low trading costs, fewer restrictions, and their high leverages.

Third, a similar cointegration relationship is found between the implied index for the futures and the implied index for the options, which offers an interesting finding to some studies suggesting that the investor in the cash stock option market do not consider the cash stock index but for the futures price as the underlying asset value (see Harvey and Whaley, 1991).

For the futures pricing model, for example, a generalized Cox, Ingersoll and Ross (CIR) model is proposed by Cox et al. (1981) and it is supposed to be more appropriate than the cost of carry model used in this study because the CIR model is able to capture a dynamic of a futures price applying a stochastic process directly. However, this model is not used widely by market traders, and an implied stock index from the model is likely not to reflect their expectations. Moreover, the CIR model is too complex to obtain the implied index, and these problems concerning the futures pricing model are left for a future study.

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