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The Impact of Introduction of QFIIs Trading on the Lead and Volatility Behavior: Evidence for Taiwan Index Futures Market

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This paper investigates whether the introduction of trading by qualified foreign institutional investors (QFIIs) has impacted the lead and volatility behavior of the futures market when the macroeconomic effects and some major economic events are controlled. First, we detect that some market inefficiency exists in Taiwan index futures market. Second, the evidence shows a strengthening in the lead of index futures over index spot markets following the introduction of trading by QFIIs. Third, we find evidence of an increase in the level of futures market volatility, implying that the quantity of information flowing into the futures market increases following the onset of trading by QFIIs. Finally, the asymmetries do not reduce after the opening up of the futures market to QFIIs. This finding is inconsistent with the view that the introduction of informed foreign investors may improve the

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reliability and quality of information and mitigate the effect of noise traders on market volatility.

Keywords: Lead-lag; volatility; stock index futures; VECM; GJR-GARCH; switching GJR-GARCH.

1. Introduction

The first futures contract (TX) based on Taiwan Stock Exchange Value-Weighted Stock Index is allowed to trade in Taiwan Futures Exchange (TAIFEX) on July 21, 1998. Nevertheless, Taiwan futures market is not active in earlier periods. The reason for the thin trading in TX may be that local investors were less familiar with the properties and trading mechanisms of futures market. In order to increase the width and depth of the Taiwan futures market, the regulators allowed qualified foreign institutional investors (QFIIs)[1] to participate in the Taiwan futures market on November 1, 1999. They hope that the increased participation of foreign institutional investors should demonstrate the effects of rational and professional investment strategies, increase the rate of information, improve the reliability and quality of information and, hence, mitigate the extraordinarily irrational and speculative mood caused by individual investors (i.e. noise traders).

Prior research concerning the impact of the introduction of foreign institutional investors on Taiwan's stock and foreign exchange markets have been conducted (e.g., Yu and Lai, 1999; Holmes and Wong, 2001; Wang and Shen, 1999), but relatively little empirical work has been done to study the issue about the Taiwan futures market. Therefore, this study will bridge a gap in the literature. One of the distinguishing characteristics of this paper is that the effects of macroeconomic factors and some major economic events other than the introduction of QFIIs to futures market are controlled so that the impact of the introduction of QFIIs on the local futures market is correctly investigated without contamination.

The purpose of this paper is to investigate the impact of the introduction of trading by QFIIs on the lead and volatility behavior of the local Taiwan futures market when the macroeconomic effects and some major economic events are controlled. First, we test the market efficiency hypothesis (MEH) in Taiwan by examining whether the cointegrated relationship (i.e., long-run equilibrium

[1]QFIIs denote foreign banks, insurance companies, fund management institutions, securities firms, and other investment institutions meeting the qualifications set by Taiwan Securities and Futures Commission (SFC). QFIIs always have sound financial resources and are regarded as informed traders.

relationship) among futures, spot markets and several macroeconomic factors exists before and after the opening up of futures market to QFIIs.

Second, this study employs the vector error correction model (VECM) with control variables to examine whether the lead-lag relationship between the futures and spot markets differs before and after the opening up of futures market to QFIIs. Additionally, in order to captures the effect of structural change due to the introduction of QFIIs, a dummy variable is incorporated into the VECM framework for the whole period of analysis to investigate the impact of the introduction of QFIIs on the lead-lag relationship between futures and spot markets.

Generally, most of the empirical studies document that futures markets are more likely to incorporate information more efficiently than spot markets, and serve a more important role in price discovery.[2] On the contrary, Huang and Shyu (1998) and Hsu and Ho (2000) report that the spot market leads the futures market in Taiwan. This study is primarily motivated by the conflicting findings. Fleming, Ostdiek, and Whaley (1996) introduce trading cost hypothesis, which argues that informed traders are attracted to derivatives markets in consequence of the transaction cost and leverage benefits provided by these markets. Therefore, on average, informed traders are more likely to trade in stock index futures markets, and price movements in stock index futures markets may lead those in stock markets. This study recognizes that this is possible to occur after the introduction of QFIIs to the Taiwan futures market. Consequently, we argue that the lead of the futures market will increase following the opening up of futures market to QFIIs.

Third, this study adopts the standard GJR-GARCH and switching GJR-GARCH models to examine the level of futures volatility and the asymmetric response of futures volatility to news before and after the opening up of futures market to QFIIs. There are two competing arguments about the possible impact of introduction of foreign investments on local market volatility, namely the stabilization-destabilization role of foreign investments. Based on Merton's (1987) market segmentation theories, there are two main factors that can explain why foreign participation may affect local share prices. First, Merton (1987) shows that foreign participation broadens the investors base in the local market, thus increases risk sharing, lowers the risk premium of stocks, and decreases volatility. The increased risk sharing, sometimes called the "base broadening" effect, provides an important

[2]This includes work by Stoll and Whaley (1990), Cheung and Ng (1990), Chan, Chan, and Karolyi (1991), Chan (1992), Iihara, Kato, and Tokunaga (1996), and Fleming, Ostdisk, and Whaley (1996) among others.

theoretical underpinning of the benefit of market liberalization (Clark and Berko, 1997). Second, foreign participation increases demand and liquidity of domestic shares, and the liquidity risk may be lowered by the flow of new investors. In contrast, Keynesians doubt whether liberalization will have a positive impact on savings and investment and suggest that short-run volatility will be increased and the degree of influence from international markets will be enlarged because of the quicker pace of transactions and inherent uncertain characteristic of trading by foreign investor.

Previous empirical research has documented considerable variation in the statistical significance of this stabilization-destabilization issue both through time and across markets. The supporters (Merton, 1987; Bekaert and Harvey, 1997; Kwan and Reyes, 1997; Choe, Kho and Stulz, 1999; Henry, 2000; Kim and Singal, 2000; Holmes and Wong, 2001; Kassimatis, 2002) suggest that the trades by well-informed foreign institutional investors improve market efficiency, increase the rate of information flow, mitigate the impact of noise trading, and tend to stabilize the financial market. On the contrary, Aitken (1996), Singh (1997), Wang and Shen (1999), and Wang (2005) document that foreign investors can have a destabilizing effect on the local financial markets.

Nevertheless, even if volatility increases, this may not be damaging to the markets. Bollerslev, Chou, and Kroner (1992), Ross (1989), Lamoureux and Lastrapes (1990), Antoniou and Holmes (1995), and Chiang and Wang (2002) argue that it is the volatility of an asset's price, not only the asset's simple price change, that is correlated to the rate of information flow, so increased volatility could make the market more efficient. Ross (1989) provides a theoretical model which shows that the variance of price change would be equal to the rate of information flow under a no-arbitrage condition, implying that volatility of the asset price will increase as the rate of information flow increases. Hence, increasing volatility is not necessarily a "bad thing" caused by speculative activity and may be the result of an increase in the rate of flow of information. Although empirical results show no consistency in this stabilization-destabilization issue, the introduction of foreign investment into local financial markets in the host country still continues all over the world.

Additionally, Black (1976), Christie (1982), French, Sewert, and Stambaugh (1987), Schwert (1989), Nelson (1991), and Engle and Ng (1993) find that in both stock and futures markets, there exists an asymmetric response to news, meaning that a negative shock to returns will increase volatility more than a positive shock of equal magnitude. Because noise traders usually do not trade on information but on noise and may overreact to new information, especially bad news, Black (1976) argues that the asymmetric effect is much more obvious if there are many noise traders in financial markets. The empirical research regarding the impacts of introduction of foreign investors on the asymmetry of local futures market is very limited. Therefore, this study will also examine whether the introduction of well-informed QFIIs into the Taiwan futures market affects the asymmetry of local futures market.

The remainder of this paper is organized as follows. Section 2 provides a description and some preliminary statistics of the sample data. Section 3 then describes the empirical methodology. Subsequently, we present the empirical results and discussion in Sec. 4. Finally, Sec. 5 makes concluding remarks.

2. DATA and Preliminary Statistics

2.1. *Macroeconomic factors and some major economic events*

Prior research has examined the relationship between macroeconomic factors and the stock prices.[3] The empirical results concerning the relationship between macroeconomic factors and the Taiwan stock index[4] show that the influences of the leading indicator, money supply (M1B) and exchange rate (TWD$/USD$) on the Taiwan stock index are always obvious and significant. Therefore, in order to clean off the effects of macroeconomic factors on the lead and volatility behavior of the futures market, this study employs three major macroeconomic factors as control variables to catch the economic systematic effects, so that the impacts of introduction of trading by foreign investors on the futures return and volatility is correctly investigated without contamination.

In addition, both the stock and futures markets use daily price-limit systems in Taiwan. The price limit on TSE and TAIFEX is 7% of the previous day's close. That is, the daily return of individual stocks and futures index cannot exceed 7% in absolute value (i.e., $+7\% \sim -7\%$). During the sample period of this study, there occurred some major events[5] that cause

[3] See Schwert (1989), Wasserflallen (1989), Fung and Lie (1990), Mookerjee and Yu (1997), and Huang (1998).
[4] For example, Huang (1998), Chiang and Wang (2002) and Liu and Lai (2002).
[5] For examples, the gigantic earthquake the rotation of political parties, the halt of the 4th nuclear power-plant and so on.

Table 1. The change of price limits.

Periods	New Price Limits
1. 1999/09/27 ∼ 1999/10/07	+7% ∼ −3.5%
2. 2000/03/20 ∼ 2000/03/24	+7% ∼ −3.5%
3. 2000/10/04 ∼ 2000/10/11	+7% ∼ −3.5%
4. 2000/10/20 ∼ 2000/11/07	+7% ∼ −3.5%
5. 2000/11/21 ∼ 2000/12/31	+7% ∼ −3.5%

Note: The data source is from the Taiwan Stock Exchange.

economic turbulence on the Taiwan stock and futures markets. Therefore, the regulators narrow down the daily price limits in five periods to reduce the impact of these major events on the price behavior of Taiwan's stock and futures markets. Table 1 lists the five periods and new price limits during the sample period. To minimize the effect of these major events on the estimation results, this study excludes the sample data listed in Table 1.

2.2. *Data description*

The sample period under investigation is from July 21, 1998 to April 30, 2001, which is split into two sub-periods (pre-QFIIs and post-QFIIs periods). The cut-off point is November 1, 1999, when the regulators allowed the participation of QFIIs in Taiwan futures market. Consequently, the pre-QFIIs period is the period without QFIIs, which is from July 21, 1998 to October 31, 1999. The post-QFIIs period covers from November 1, 1999 to April 30, 2001, which is the period when QFIIs was permitted in the local futures market.

The reason for our interest in empirical work on the TX index futures market is that it is the most active futures contract in Taiwan futures markets. The other futures contracts, such as electronic and financial sector index futures markets, are not included in our sample due to thin and discrete trading volume by foreign investors. Daily data of the Taiwan Stock Index Futures contract (TX), its corresponding underlying spot index and exchange rate is used in this study. Consistent with prior research, the nearby futures contract is used to construct futures returns. The data for the spot prices, exchange rate and futures prices are retrieved from the

Taiwan Economic Journal (TEJ)[6] and the TAIFEX database, respectively. Monthly leading indicator and M1B (money supply) data is constructed by the Council for Economic Planning and Development (CEPD) and Central Bank in Taiwan, respectively. Following Chiang and Wang (2002), we transform the leading indicator and M1B to daily basis. The spot returns and futures returns are obtained by taking the natural logarithmic difference of the price levels, respectively. That is, $R_{Ft} = F_t - F_{t-1}$ and $R_{st} = S_t - S_{t-1}$ where F_t is the natural logarithm of the futures price, and S_t is the natural logarithm of underlying spot price. Ex, L and M1b denote the natural logarithmic transformation of the TWD/USD exchange rate, leading indicator and M1B, respectively.

Table 2 reports the summary statistics for the $\mathrm{R_F}$, Rs, Ex, L and M1b variables for the pre-QFIIs, post-QFIIs and whole periods, respectively. The standard deviations of the futures returns are higher in the post-QFIIs period than in the pre-QFIIs and whole periods, implying that futures market volatility increases in the post-QFIIs period. The Jarque-Bera statistics are statistically significant at the 1% level for the three periods, indicating that none of the five series is normally distributed. All the figures of 6 and 12 lags Ljung-Box statistics of the futures and spot return series are not significant for the pre-QFIIs period. The figures of 12 lags Ljung-Box statistics of futures returns series are significant for both the post-QFIIs and whole periods, indicating that the autocorrelations in futures returns are present. All the figures of Ljung-Box statistics of the Ex, L and m1b series are statistically significant at the 1% level for the three periods. Additionally, the significant Ljung-Box statistics of the squared returns for the three periods indicate that nonlinear dependence exists in the futures and spot return series and the variances of the futures and spot returns change over time. Hence, we will utilize ARCH-type process (Engle, 1982; Bollerslev, 1986; Bollerslev, Chou, and Kroner, 1992) to model the time varying nature of futures price volatility.

This paper uses the ADF test by Dickey and Fuller (1981) to detect whether unit roots exist in the five time series for the three periods. Table 3 reports results for unit root tests. The evidence shows that the unit root hypothesis in the natural logarithm of each level series cannot be rejected at the 1% significance levels, but all the five time series are transformed

[6]TEJ is a private data-source company. It provides the most comprehensive and reliable economic and financial data base.

Table 2. Summary statistics.

	Pre-QFIIs (1998/07/21 ~ 1999/10/31)					Post-QFIIs (1999/11/01 ~ 2001/04/30)					The Whole Period (1998/07/21 ~ 2001/04/30)				
	R_F	R_S	Ex	L	M1b	R_F	R_S	Ex	L	M1b	R_F	R_S	Ex	L	M1b
Mean	−0.000049	−0.000035	32.83	99.7	38623	−0.001278	−0.000986	32.56	104.5	48135	−0.000597	−0.000487	32.19	102.4	42572
Std. Dev.	0.017	0.016	0.90	2.68	1971.1	0.036	0.031	0.68	4.21	1893.2	0.028	0.025	1.19	3.81	3361.5
Skewness	0.009	0.038	1.04	0.32	−0.031	0.251	1.152	0.873	−0.78	−0.487	0.489	1.239	0.732	0.15	−0.027
Kurtosis	5.30	4.40	2.94	1.55	1.558	13.83	16.92	3.115	3.63	2.79	12.97	14.51	3.12	2.35	1.932
J-B	73.8*	27.71*	61.21*	34.86*	29.04*	110.2*	59.92*	55.68*	41.5*	17.94*	198.2*	74.39*	55.97*	37.29*	37.28*
LB Q(6)	0.885	2.819	1879*	1954*	1983*	28.52*	7.64	1683*	1672*	1686*	23.27*	6.41	3759*	3791*	3862*
LB Q (12)	8.323	5.972	3478*	3758*	3859*	29.16*	12.9	2898*	3167*	3145*	26.97**	8.63	7108*	7467*	7403*
LB Q^2 (6)	14.6**	19.39*				22.72*	18.59**				48.17*	26.19*			
LB Q^2 (12)	18.3	22.42**				27.96*	21.17**				57.69*	36.37*			

Note: The Jarque-Bera statistic tests whether a series is normally distributed under the null hypothesis of normality.

Ljung-Box Q (k) statistic tests the joint significance of the autocorrelations of the daily return series up to the kth order. Ljung-Box Q^2 (k) statistic tests the joint significance of the autocorrelations of the squared daily return series up to the kth order.

* and ** indicate statistically significant at 1% and 5% level, respectively.

Table 3. Unit root test results of the five time series.

	Pre-QFIIs	Post-QFIIs	The Whole Period
Levels	ADF Test	ADF Test	ADF Test
F_t	−2.84	−1.92	−1.27
S_t	−2.65	−1.79	−1.38
Ex_t	−1.68	−1.72	−1.29
L_t	−1.53	−1.35	0.96
$M1b_t$	−2.31	−1.89	−1.73
First-difference			
ΔF_t	−17.98*	−10.82*	−11.52*
ΔS_t	−16.77*	−19.51*	−25.74*
ΔEx_t	−8.58*	−13.81*	−14.17*
ΔL_t	−4.51*	−17.98*	−6.36*
$\Delta M1b_t$	−4.48*	−7.19*	−7.95*

Note: The critical values for ADF test at the 5% and 1% levels are −3.42 and −3.97, respectively (see Mackinnon, 1996). H_0 : unit root, H_A : no unit root.
* and ** indicate that the statistic is significant at the 1% and 5% levels, respectively.

to achieve stationarity by taking the first difference of the natural logarithm of each level series. These results suggest that the five variables are nonstationary I (1) series, which fulfills the necessary condition for the cointegration test.

3. Methodology

3.1. *Cointegration test and vector error correction model*

First, we test the market efficiency hypothesis (MEH) in Taiwan index futures market by examining whether the cointegrated relationship (i.e., long-run equilibrium relationship) among futures, spot prices and several macroeconomic factors exists before and after the opening up of futures market to QFIIs. Given that the five variables are integrated of order one, the cointegration test proposed by Johansen and Juselius (1990) is performed. If there are cointegrated relationships among futures, spot prices and several macroeconomic factors, then we suggest that some market inefficiency exists in Taiwan index futures market.

Second, for cointegrated series, Granger causality tests need to be performed in the corresponding VECM framework according to the Granger Representation Theorem proposed by Engle and Granger (1987). This study employs the VECM to examine whether the lead-lag relationship between the futures and spot markets differs for the pre- and post-QFIIs periods. To control effects of macroeconomic factors on the relationship between the futures and spot markets, we incorporate the macroeconomic factors into the VECM. Therefore, this paper adopts the following VECM[7] framework with five variables to study the lead-lag relationship between the futures and spot markets for the pre-QFIIs, post-QFIIs, and whole periods, respectively.

$$\Delta F_t = \mu_1 + \beta_1 Z_{t-1} + \sum_{i=1}^{m1} a_{1i} \Delta F_{t-i} + \sum_{i=1}^{n1} b_{1i} \Delta S_{t-i} + \sum_{i=1}^{o1} c_{1i} \Delta Ex_{t-i} + \sum_{i=1}^{p1} d_{1i} \Delta L_{t-i} + \sum_{i=1}^{q1} e_{1i} \Delta M1b_{t-i} + \varepsilon_{1t} \quad (1)$$

$$\Delta S_t = \mu_2 + \beta_2 Z_{t-1} + \sum_{i=1}^{m2} a_{2i} \Delta F_{t-i} + \sum_{i=1}^{n2} b_{2i} \Delta S_{t-i} + \sum_{i=1}^{o2} c_{2i} \Delta Ex_{t-i} + \sum_{i=1}^{p2} d_{2i} \Delta L_{t-i} + \sum_{i=1}^{q2} e_{2i} \Delta M1b_{t-i} + \varepsilon_{2t} \quad (2)$$

where Z_{t-1} (e.g., $Z_{t-1} = F_{t-1} - \alpha - \omega_1 S_{t-1} - \omega_2 Ex_{t-1} - \omega_3 L_{t-1} - \omega_4 M1b_{t-1}$) is error correction term and ΔF_t, ΔS_t, ΔEx_t, ΔL_t and $\Delta M1b_t$ denote the natural logarithmic difference of the futures price, spot price, exchange rate, leading indicator and M1b, respectively. The Z_{t-1} error correction term means the previous period's equilibrium error, which expresses the deviations from a "long-run" cointegration equilibrium in the last period. We employ the Akaike Information Criterion (AIC) to determine the appropriate lag length in the VECM framework. The coefficients β_1 and β_2 represent the adjustment speeds of error correction mechanism in ΔF_t and ΔS_t, respectively, toward long-run equilibrium cointegration relationship. If the coefficient on Z_{t-1} in the process of ΔF_t (ΔS_t) is small, F_t (S_t) has little tendency to adjust a disequilibrium situation. That is, most of the adjustment will be done by S_t (F_t), futures (spot) play a more important role in price discovery. Moreover, we implement the Granger-causality test between futures and spot markets by computing a standard F-test with the null hypothesis that all the lagged coefficients b_{1i} or a_{2i} are jointly equal to zero. If some b_{1i} (a_{2i}) are statistically different from zero but some a_{2i} (b_{1i}) not, the spot (futures) market Granger-causes futures (spot) market. In other

[7]Because the main focus of this study is to examine the lead-lag relationship of price discovery between the index spot and index futures, the specifications of macroeconomic factors in the VECM are not showed here for saving space.

words, the spot (futures) market shows a lead effect if the spot (futures) market Granger-causes futures (spot) market. This indicates that there is a unidirectional Granger causality between futures and spot markets. Furthermore, if both a_{2i} and b_{1i} are significantly different from zero, there exists a bidirectional causality between futures and spot markets.

Additionally, in order to test the impact of structural change due to the introduction of QFIIs on the short term dynamics and long-run error correction term between stock index and stock index returns, a dummy variable (d_t) is introduced into Eqs. (1) and (2) for the whole period.[8] The modified model may be specified as follows:

$$\begin{aligned}\Delta F_t = {}& \mu_1 + \beta_1 Z_{t-1} + \sum_{i=1}^{m1} a_{1i}\Delta F_{t-i} + \sum_{i=1}^{n1} b_{1i}\Delta S_{t-i} + \sum_{i=1}^{o1} c_{1i}\Delta Ex_{t-i} \\ & + \sum_{i=1}^{p1} d_{1i}\Delta L_{t-i} + \sum_{i=1}^{q1} e_{1i}\Delta M1b_{t-i} + \beta_{1d} Z_{t-1} d_t \\ & + \sum_{i=1}^{n1} b_{1di}\Delta S_{t-i} d_{t-i} + \varepsilon_{ft} \end{aligned} \tag{3}$$

$$\begin{aligned}\Delta S_t = {}& \mu_2 + \beta_2 Z_{t-1} + \sum_{i=1}^{m2} a_{2i}\Delta F_{t-i} + \sum_{i=1}^{n2} b_{2i}\Delta S_{t-i} + \sum_{i=1}^{o2} c_{2i}\Delta Ex_{t-i} \\ & + \sum_{i=1}^{p2} d_{2i}\Delta L_{t-i} + \sum_{i=1}^{q2} e_{2i}\Delta M1b_{t-i} + \beta_{2d} Z_{t-1} d_t \\ & + \sum_{i=1}^{m2} a_{2di}\Delta F_{t-i} d_{t-i} + \varepsilon_{st}, \end{aligned} \tag{4}$$

where d_t is a dummy variable that takes on a value of 1 if observation t lies within the post-QFIIs period, otherwise 0. If the dummy is statistically significant, then the introduction of QFIIs has an impact on the lead-lag relationship between futures and spot markets.

3.2. *Switching GJR-GARCH (1,1) and standard GJR-GARCH (1,1) models*

In analyzing the impact of the opening up of the futures market to QFIIs on the level and nature of futures price volatility, there are two issues that need

[8]See Chang and Nieh (2001). In order to investigate the impact of the US stock crash of October 1997 on some markets, a dummy variable is incorporated into the VECM model.

to be addressed. First, does the existence of QFIIs trading in itself have any effect on volatility? Second, if the existence of QFIIs trading affects volatility, how does it?

To address the first issue, a switching GJR-GARCH (1,1) model is employed to examine the impact of the existence of QFIIs trading on the level and nature of futures price volatility for the whole sample period. Lee and Ohk (1992) present the modified GARCH model, which imposes an autoregressive structure on conditional variance and captures the change in the level and slope of time-varying volatility using dummy variables. This modified model is called the switching GARCH model. But the switching GARCH model is connected with the shortcoming that it assumes a symmetric response to news and fails to account for observed asymmetry in the market. Glosten, Jagannathan, and Runkle (1993) propose a GJR-GARCH model, which can capture the asymmetric impact of shocks on volatility.

Hence, in the spirit of Lee and Ohk (1992) and Glosten, Jagannathan, and Runkle (1993), we propose the conditional variance equation of GJR-GARCH (1, 1) with a dummy variable d_t. As stated in the previous section it is necessary to remove the influences of macroeconomic factors on futures market volatility by incorporating control variables into the mean equation. Consequently, the modified model, switching GJR-GARCH (1,1), can be specified as follows:

$$\begin{aligned} R_{F,t} &= \theta_0 + \sum_{i=1}^{p} \theta_{1,i} R_{F,t-i} + \sum_{j=0}^{m} \theta_{2,j} R_{S,t-j} + \sum_{j=0}^{m} \theta_{3,j} \Delta Ex_{t-j} \\ &\quad + \sum_{j=0}^{m} \theta_{4,j} \Delta L_{t-j} + \sum_{j=0}^{m} \theta_{5,j} \Delta M1b_{t-j} + \varepsilon_t \end{aligned} \tag{5}$$

$$\begin{aligned} h_t &= \omega_0 + \omega_1 \varepsilon_{t-1}^2 + \omega_2 h_{t-1} + \omega_3 \varepsilon_{t-1}^2 D_{t-1} + \gamma_0 d_t + \gamma_1 \varepsilon_{t-1}^2 d_t \\ &\quad + \gamma_2 h_{t-1} d_t + \gamma_3 \varepsilon_{t-1}^2 D_{t-1} d_t, \end{aligned} \tag{6}$$

where d_t is a dummy variable, namely switching point t^* (November 1, 1999), taking on the value zero in the pre-QFIIs period and one in the post-QFIIs period. D_{t-1} is a dummy variable which is 1 in response to bad news ($\varepsilon_{t-1} < 0$) and zero in response to good news ($\varepsilon_{t-1} > 0$). Equation (5) is the conditional mean equation. $R_{F,t}$ is regressed on its AR (p) process to eliminate the serially correlated residuals. The spot returns, exchange rate, leading indicator and M1b are included in Eq. (5) to filter out the effect of macroeconomic factors on futures market volatility. Ω_{t-1} is the information set available up to time t. Equation (6) is the conditional variance equation and h_t represents the conditional variance term at time t. If the dummy is statistically significant then the existence of QFIIs trading has had an impact on futures market volatility. A positive γ_0 coefficient indicates increased unconditional volatility in the post-QFIIs period, whereas a negative γ_0 coefficient indicates decreased unconditional volatility in the post-QFIIs period. Furthermore, we may test the impact of QFIIs trading on conditional futures price volatility through the γ_1 and γ_2 coefficients. Finally, the impact of QFIIs trading on the asymmetry can be assessed through γ_3. The Berndt–Hall–Hall–Hausman (BHHH, 1974) optimization algorithm is used to calculate maximum likelihood estimates of each of the coefficients in the conditional mean and variance equations.

To address the second issue, the following standard GJR-GARCH (1, 1) model is estimated for the pre-QFIIs, post-QFIIs, and whole periods, thereby, allowing a comparison of the nature of volatility before and after the onset of QFIIs trading.

$$R_{F,t} = \theta_0 + \sum_{i=1}^{p} \theta_{1,i} R_{F,t-i} + \sum_{j=0}^{m} \theta_{2,j} R_{S,t-j} + \sum_{j=0}^{m} \theta_{3,j} \Delta Ex_{t-j} + \sum_{j=0}^{m} \theta_{4,j} \Delta L_{t-j} + \sum_{j=0}^{m} \theta_{5,j} \Delta M1b_{t-j} + \varepsilon_t \tag{7}$$

$$h_t = \omega_0 + \omega_1 \varepsilon_{t-1}^2 + \omega_2 h_{t-1} + \omega_3 \varepsilon_{t-1}^2 D_{t-1} \tag{8}$$

First, we can examine the level of unconditional volatility through the ω_0 coefficient. If the value of ω_0 increases in the post-QFIIs period, the level of futures market volatility increases after the onset of QFIIs trading. Furthermore, coefficient ω_3 measures the extent to which there is an asymmetric response of volatility to news, and change of its value is important in our analysis. If ω_3 is positive, an asymmetric effect exists in the data as a negative return will increase the volatility more than a positive return of the same magnitude does. Therefore, a decrease in the value of ω_3 in the post-QFIIs period will imply that asymmetric effect has decreased following the introduction of QFIIs, whereas a rise in ω_3 will suggest the opposite.

4. Empirical Results

4.1. *Cointegration test and vector error correction model*

Panel A of Table 4 lists the results of the multivariate cointegration test for the three periods. Both the trace statistics and max-eigen statistics show that there exists one cointegrating vector among these five variables at the 1% significance level for the three periods, indicating that there exists a long-run equilibrium relationship between these five series. The evidence of the cointegrated relationship among futures, spot prices and several macroeconomic factors violates the market efficiency hypothesis (MEH).

Panel B of Table 4 reports the estimation results of the VECM framework.[9] In the pre-QFIIs period, the F-statistics testing the null hypothesis that all the lagged coefficients on the spot variables are jointly equal to zero in the futures equation is significant at the 5% level, whereas the F-statistics of all the lagged coefficients on the futures variables in the spot equation are insignificant. It appears that spot market Granger-causes futures market in the pre-QFIIs period. Furthermore, the error correction term coefficient β_1 has the right sign and is statistically significant, but β_2 is not significant. This means that, futures market will make all the adjustment to reestablish the equilibrium situation during the next period in cases of short-run deviations from the long-run cointegrated equilibrium in the pre-QFIIs period. These findings show that spot market leads futures market in the pre-QFIIs period.

The evidence, however, is different in the post-QFII period. The F-statistics of all the lagged coefficients on the spot variables in the futures equation are significant at 5% level, and the F-statistics of all the lagged coefficients on the futures variables in the spot equation are also significant. The error correction term coefficients β_1 and β_2 are statistically significant at the 1% significance level, suggesting that both the spot and futures markets will make the adjustment together to reestablish the equilibrium situation when the long-run disequilibrium occurs. This evidence indicates that there exists bi-directional Granger-causality between spot and futures markets in the post-QFII period. Notably, however, the error correction term in the spot equation is greater in magnitude than that of the futures equation, implying that the spot market makes more adjustment to reestablish the equilibrium than does futures market. This provides some evidence that the causality from futures to spot is stronger than the reverse direction in the post-QFII period.

In the whole period, the estimation results of error correction terms of the futures and spot markets are similar to those for the post-QFIIs

[9]Because the main focus of this study is to examine the lead-lag relationship of price discovery between the index spot and index futures, the estimation results of macroeconomic factors for the VECM are not reported here for saving space.

Table 4. Estimation results of cointegration test and vector error correction model.

$$\Delta F_t = \mu_1 + \beta_1 Z_{t-1} + \sum_{i=1}^{m1} a_{1i}\Delta F_{t-i} + \sum_{i=1}^{n1} b_{1i}\Delta S_{t-i} + \sum_{i=1}^{o1} c_{1i}\Delta Ex_{t-i} + \sum_{i=1}^{p1} d_{1i}\Delta L_{t-i} + \sum_{i=1}^{q1} e_{1i}\Delta M1b_{t-i} + \varepsilon_{1t}$$

$$\Delta S_t = \mu_2 + \beta_2 Z_{t-1} + \sum_{i=1}^{m2} a_{2i}\Delta F_{t-i} + \sum_{i=1}^{n2} b_{2i}\Delta S_{t-i} + \sum_{i=1}^{o2} c_{2i}\Delta Ex_{t-i} + \sum_{i=1}^{p2} d_{2i}\Delta L_{t-i} + \sum_{i=1}^{q2} e_{2i}\Delta M1b_{t-i} + \varepsilon_{2t}$$

	Pre-QFIIs		Post-QFIIs		The Whole Period	
	λ_{Trace}	$\lambda_{\text{max-eigen}}$	λ_{Trace}	$\lambda_{\text{max-eigen}}$	λ_{Trace}	$\lambda_{\text{max-eigen}}$
Panel A: Cointegration test						
H_0: $r = 0$	103.41*	54.16*	97.67*	54.28*	173.16*	108.52*
H_0: $r \leq 1$	42.25	24.11	43.78	25.32	52.34**	29.95**
H_0: $r \leq 2$	25.13	13.34	21.49	13.62	24.87	17.86
H_0: $r \leq 3$	11.78	10.97	6.83	6.72	6.72	7.23
Error Correction Term	$z_t = F_t - 0.67 - 1.02\,S_t + 0.15\,Ex_t + 0.4\,L_t - 0.14\,M1b_t$		$z_t = F_t + 2.41 - 0.98\,S_t - 0.26\,Ex_t - 0.21\,L_t - 0.05M1\,b_t$		$z_t = F_t - 1.37 - 0.98\,S_t + 0.20\,Ex_t - 0.02\,L_t + 0.49\,M1b_t$	

Table 4. (*Continued*)

Variable	ΔF_t Coeff.	ΔF_t t	ΔS_t Coeff.	ΔS_t t	ΔF_t Coeff.	ΔF_t t	ΔS_t Coeff.	ΔS_t t	ΔF_t Coeff.	ΔF_t t	ΔS_t Coeff.	ΔS_t t
Panel B: Vector error correction model												
Constant	−0.001	−0.63	−0.001	−0.68	−0.0008	0.61	−0.0028	−0.30	−0.00081	−0.98	−0.0038	−0.55
ΔZ_{t-1}	−0.51	−3.92*	−0.09	−0.67	−0.41	−3.45*	0.51	5.52*	−0.38	−4.25*	0.41	5.43*
ΔF_{t-1}	−0.28	−2.09**	−0.01	−0.09	0.23	1.91	0.12	1.29	0.14	1.88	0.19	2.81*
ΔF_{t-2}	−0.13	−1.21	−0.04	−0.39	0.42	3.61*	0.31	4.19*	0.20	3.02*	0.23	4.13*
ΔF_{t-3}	0.06	0.58	0.05	0.51	0.35	3.72*	0.17	2.73*	0.21	3.34*	0.11	2.05**
ΔS_{t-1}	0.36	2.58*	0.08	0.61	−0.41	−3.43*	−0.25	−3.24*	−0.13	−1.90	−0.15	−2.38**
ΔS_{t-2}	0.13	2.21**	0.02	0.12	−0.31	−2.89**	−0.17	−2.36**	−0.15	−2.29**	−0.17	−2.98 *
ΔS_{t-3}	−0.03	−1.27	−0.04	−0.37	−0.89	−1.43	−0.03	−0.71	−0.13	−2.26**	−0.07	−1.66
ΔEx_{t-1}	−0.03	−0.06	−0.08	−0.21	−0.21	−0.28	0.18	0.39	0.29	0.07	−0.14	−0.43
ΔEx_{t-2}	0.11	0.29	−0.25	−0.63	−0.72	−0.81	−0.18	−0.32	−0.09	−0.24	−0.43	−1.31
ΔEx_{t-3}	0.06	0.16	−0.04	−0.11	0.58	0.75	0.47	0.82	0.26	0.67	−0.15	−0.47
ΔL_{t-1}	2.23	1.83	3.41	2.78*	−0.37	−0.29	1.72	2.16**	0.48	0.62	1.64	2.51**
ΔL_{t-2}	−0.41	−0.32	−1.29	−1.02	1.68	1.48	0.68	0.80	1.83	2.35**	0.04	0.07
ΔL_{t-3}	0.67	0.51	1.91	1.42	−0.66	−0.62	−0.73	−0.99	−0.53	−0.72	−0.61	−0.99
$\Delta M1b_{t-1}$	−1.03	−0.76	−2.40	−1.62	−0.05	−0.04	−0.18	−0.39	0.25	0.52	−0.26	−0.62
$\Delta M1b_{t-2}$	2.09	1.56	3.82	2.83*	−0.07	−0.11	0.16	0.38	0.12	0.25	0.18	0.41
$\Delta M1b_{t-3}$	−1.21	−0.83	−2.09	−1.42	1.82	2.73*	0.51	0.98	1.41	2.82*	0.13	0.38
Granger causality test:												
$H_0 : b_{1i} = 0$	F-Stats = 3.31**				F-Stats = 3.12**				F-Stats = 2.32			
$H_0 : a_{2i} = 0$			F-Stats = 0.08				F-Stats = 8.75*				F-Stats = 6.21*	

Note:
r denotes the cointegrating vectors number in Panel A.
The appropriate lag-length specification of each equation is determined using Akaike's Information Criterion (AIC). The lag-length is 2 in the pre-QFIIs period and 3 in the post-QFIIs period.
*and ** indicate that the statistic is significant at the 1% and 5% levels, respectively.

period. The F-statistics of all the lagged coefficients on the spot variables in the futures equation are insignificant at the 5% level, while the F-statistics of all the lagged coefficients on the futures variables in the spot equation are significant at the 1% significance level. This means that unidirectional causality running from the futures market to the spot market, indicating that futures market Granger-causes spot market in the whole period.

Table 5 reports the estimation results of the modified VECM with dummy variables for the whole period. F-statistics testing the null hypothesis that all the lagged coefficients on the dummy variables are jointly equal to zero is rejected at the 5% significance for both the futures and spot equations. This indicates that the introduction of QFIIs has an impact on the lead-lag relationship between futures and spot markets and there exists bi-directional Granger-causality between spot and futures markets in the post-QFIIs period. The coefficient on error correction term dummy of the futures equation is not statistically significant, while that of the spot equation is significant, indicating that spot market makes all the adjustment to reestablish the equilibrium situation after the onset of QFIIs trading. The results reported in Table 5 are similar to those in Table 4.

In general, the lead of futures markets has long been documented in the theoretical and empirical literature.[10] Overall, Tables 4 and 5 provide evidence that Granger-causality between Taiwan futures and spot markets differs before and after the onset of QFIIs trading, documenting a strengthening in the lead of index futures over index spot markets following the opening up of the Taiwan futures market to QFIIs.

4.2. *Switching GJR-GARCH (1,1) and standard GJR-GARCH (1,1) models*

Table 6 lists the estimation results of the switching GJR-GARCH (1,1) model for the whole period. The coefficients on dummy variables (γ_0 to γ_2) are statistically significant, indicating that the existence of QFIIs trading has had an impact on futures market volatility. The γ_0 coefficient is significantly positive, implying that the mean level of unconditional volatility increases following the opening up of the local futures markets to QFIIs. This evidence supports a more favorable view of QFIIs trading in which the introduction of QFIIs trading enhances information flows. The γ_1 and γ_2

[10] See, for examples, Stoll and Whaley (1990), Cheung and Ng (1990), Chan, Chan and Karolyi (1991), Chan (1992), Iihara, Kato and Tokunaga (1996), Fleming, Ostdiek, and Whaley (1996) and so on.

Table 5. Estimation results of vector error correction model with dummy variables.

$$\Delta F_t = \mu_1 + \beta_1 Z_{t-1} + \sum_{i=1}^{m1} a_{1i}\Delta F_{t-i} + \sum_{i=1}^{n1} b_{1i}\Delta S_{t-i} + \sum_{i=1}^{o1} c_{1i}\Delta Ex_{t-i} + \sum_{i=1}^{p1} d_{1i}\Delta L_{t-i} + \sum_{i=1}^{q1} e_{1i}\Delta M1b_{t-i} + \beta_{1d} Z_{t-1} d_t + \sum_{i=1}^{n1} b_{1di}\Delta S_{t-i} d_t + \varepsilon_{1t}$$

$$\Delta S_t = \mu_2 + \beta_2 Z_{t-1} + \sum_{i=1}^{m2} a_{2i}\Delta F_{t-i} + \sum_{i=1}^{n2} b_{2i}\Delta S_{t-i} + \sum_{i=1}^{o2} c_{2i}\Delta Ex_{t-i} + \sum_{i=1}^{p2} d_{2i}\Delta L_{t-i} + \sum_{i=1}^{q2} e_{2i}\Delta M1b_{t-i} + \beta_{2d} Z_{t-1} d_t + \sum_{i=1}^{m2} a_{2di}\Delta F_{t-i} d_t + \varepsilon_{2t}$$

	ΔF_t		ΔS_t	
Variable	Coeff.	t	Coeff.	t
Constant	−0.0004	−0.39	−0.00035	−0.37
Z_{t-1}	−0.52	−3.28*	−0.17	−1.31
ΔF_{t-1}	−0.39	−2.28**	−0.18	−0.11
ΔF_{t-2}	−0.24	−1.35	−0.004	−0.03
ΔF_{t-3}	0.11	0.19	0.03	0.32
ΔS_{t-1}	0.36	2.01**	−0.03	−0.19
ΔS_{t-2}	0.28	1.731	0.06	0.51
ΔS_{t-3}	0.12	0.72	0.0002	0.005
ΔEx_{t-1}	0.15	0.38	0.12	0.39
ΔEx_{t-2}	−0.08	−0.22	−0.24	−0.78
ΔEx_{t-3}	0.39	0.98	0.14	0.41
ΔL_{t-1}	0.55	0.72	1.79	2.82*
ΔL_{t-2}	1.85	2.52**	0.48	0.80
ΔL_{t-3}	−0.71	−0.94	−0.55	−0.78
$\Delta M1b_{t-1}$	0.25	0.49	−0.38	−0.83
$\Delta M1b_{t-2}$	0.17	0.36	0.24	0.61
$\Delta M1b_{t-3}$	1.53	3.13*	0.16	0.47
$Z_{t-1}d_t$	0.19	1.23	0.73	4.85*
$S_{t-1}d_{t-1}$ or $\Delta F_{t-1}d_{t-1}$	−0.58	−2.68**	0.43	2.63*
$S_{t-2}\ d_{t-2}$ or $\Delta F_{t-2}d_{t-2}$	−0.51	−2.24**	0.45	2.58*
$S_{t-3}\ d_{t-3}$ or $\Delta F_{t-3}d_{t-3}$	−0.19	−1.41	0.09	0.71
Granger causality test:				
$H_0 : b_{1di} = 0$	F-Stats = 3.67**			
$H_0 : a_{2di} = 0$			F-Stats = 6.82*	

Note: * and ** indicate that the statistic is significant at the 1% and 5% levels, respectively.

coefficients are significant, revealing that the conditional variance for futures price underwent some form of change after the onset of QFIIs trading. The γ_3 coefficient on the asymmetric dummy is not significant but positive, showing that there is no evidence of significant asymmetric volatility response in the post-QFIIs period.

Table 6. Estimation results of the switching GJR-GARCH (1,1) model.

$$R_{F,t} = \theta_0 + \sum_{i=1}^{p} \theta_{1,i} R_{F,t-i} + \sum_{j=0}^{m} \theta_{2,j} R_{S,t-j} + \sum_{j=0}^{m} \theta_{3,j} Ex_{t-j} + \sum_{j=0}^{m} \theta_{4,j} L_{t-j} + \sum_{j=0}^{m} M1b_{t-j} + \varepsilon_t$$

$$h_t = \omega_0 + \omega_1 \varepsilon_{t-1}^2 + \omega_2 h_{t-1} + \omega_3 \varepsilon_{t-1}^2 D_{t-1} + \gamma_0 d_t + \gamma_1 \varepsilon_{t-1}^2 d_t + \gamma_2 h_{t-1} d_t + \gamma_3 \varepsilon_{t-1}^2 D_{t-1} d_t$$

	Coefficient	p-value
θ_0	0.0018	0.583
θ_{11}	−0.5281	0.000*
θ_{12}	−0.2997	0.000*
θ_{13}	−0.1497	0.0001*
θ_{20}	0.8572	0.000*
θ_{21}	0.5082	0.005*
θ_{22}	0.1425	0.078**
θ_{23}	0.0918	0.1221
θ_{30}	−0.5243	0.0075*
θ_{31}	0.1934	0.3484
θ_{32}	0.2857	0.1461
θ_{33}	−0.2841	0.1491
θ_{40}	−0.4893	0.4672
θ_{41}	−0.9465	0.1223
θ_{42}	1.0186	0.051
θ_{43}	−0.2491	0.6294
θ_{50}	−0.7229	0.1875
θ_{51}	0.2287	0.5731
θ_{52}	−0.7852	0.1176
θ_{53}	0.6872	0.1985
ω_0	0.000011	0.001*
ω_1	0.5572	0.002*
ω_2	0.1067	0.4623
ω_3	0.0374	0.8377
γ_0	0.000047	0.031**
γ_1	−0.4729	0.0128**
γ_2	0.7491	0.000*
γ_3	0.1976	0.462
Model Diagnostics Test on the Standardized Residuals		
LB Q(6) = 8.34	LB Q^2(6) = 6.29	ARCH(6) = 5.96

Notes: LB Q(6), LB Q^2(6) are the Ljung-Box statistics applied on the standardized and squared standardized residuals, respectively.
ARCH(6) is the statistics used to test whether standardized residuals exhibit the ARCH effect up to the order 6.
* and ** indicate that the statistic is significant at the 1% and 5% levels, respectively.

Table 7 shows the parameter estimates of the standard GJR-GARCH(1,1) models for the three periods. The coefficient ω_0 is statistically significant and increases from 0.0000012 to 0.000021, revealing that the level of unconditional volatility increases after the onset of QFIIs trading. This is consistent with the argument that the introduction of QFII trading may increase the speed of information transmission to the local futures market. The coefficient ω_3 for the asymmetric effect is statistically significant at the 1% level for both pre- and post-QFIIs periods, but not for the whole period. However, the pattern of asymmetries is different for the pre- and post-QFIIs periods. In the pre-QFIIs period, a negative ω_3 coefficient indicates that "good news" results in more volatility than "bad news" of the same magnitude, which is inconsistent with the negative relationship between current returns and future volatility observed by Black (1976). On the contrary, a positive ω_3 coefficient shows that "bad news" results in more volatility than "good news" of the same magnitude in the post-QFIIs period. The sum of coefficients ω_1 and ω_3 for the impact of bad news increases from -0.2018 to 0.3188, indicating that there exists an increased asymmetric response of volatility to news following the introduction of QFIIs trading. The increased asymmetric effects may result from the fact that the major proportion of investors on the TAIFEX is individual traders who are poorly trained and more inclined to overreact to bad news even if the QFIIs are introduced in Taiwan futures market. The standard diagnostic tests of the standardized and squared standardized residuals from the switching and standard GJR-GARCH (1, 1) models confirm the absence of any further ARCH effects, suggesting an appropriate model specifications.

Overall, these results reported in Tables 6 and 7 suggest that the speed and quantity of information flowing into the local futures market increase following the onset of trading by QFIIs, but the asymmetric response behavior does not reduce. This evidence is inconsistent with the view that the introduction of informed foreign investors may improve the reliability and quality of information and mitigate the effect of noise traders on market volatility.

5. Conclusion

The purpose of this paper is to examine the influence of the opening up of the Taiwan futures market to well-informed QFIIs on the lead and volatility behavior of the local futures market when macroeconomic effects are controlled.

Table 7. Estimation results of the standard GJR-GARCH (1,1) model.

$$R_{F,t} = \theta_0 + \sum_{i=1}^{p} \theta_{1,i} R_{F,t-i} + \sum_{j=0}^{m} \theta_{2,j} R_{S,t-j} + \sum_{j=0}^{m} \theta_{3,j} Ex_{t-j} + \sum_{j=0}^{m} \theta_{4,j} L_{t-j} + \sum_{j=0}^{m} M1b_{t-j} + \varepsilon_t$$

$$h_t = \omega_0 + \omega_1 \varepsilon_{t-1}^2 + \omega_2 h_{t-1} + \omega_3 \varepsilon_{t-1}^2 D_{t-1}$$

	Pre-QFIIs		Post-QFIIs		The Whole Period	
	Coefficient	p-value	Coefficient	p-value	Coefficient	p-value
θ_0	0.0015	0.0093**	−0.0014	0.2084	0.0004	0.4726
θ_{11}	−0.5533	0.0000*	−0.3368	0.0000*	−0.5357	0.0000*
θ_{12}	−0.2169	0.0001*	−0.0851	0.3732	−0.3004	0.0000*
θ_{13}	−0.0874	0.1174	−0.0812	0.3562	−0.0895	0.0463**
θ_{20}	0.9209	0.0000*	0.3567	0.0000*	0.8687	0.0000*
θ_{21}	0.6160	0.0000*	0.1572	0.0868	0.1148	0.0197**
θ_{22}	0.3259	0.0000*	0.0437	0.7128	0.0832	0.9265
θ_{23}	0.1159	0.0691	−0.0401	0.5875	0.0762	0.834
θ_{30}	−0.1126	0.5869	−2.0126	0.0047*	−0.3589	0.0721
θ_{31}	0.0226	0.9044	0.5429	0.4329	0.0965	0.6439
θ_{32}	0.3840	0.0290**	0.3187	0.6763	0.1764	0.4003
θ_{33}	−0.2504	0.1564	−0.1295	0.8874	−0.1162	0.5528
θ_{40}	0.4242	0.6446	−1.3762	0.6753	0.0899	0.9012
θ_{41}	−1.2366	0.1690	−0.3247	0.6783	−0.6572	0.3664
θ_{42}	0.7580	0.1784	2.5642	0.3872	0.2276	0.6553
θ_{43}	−1.2310	0.0072*	−0.0817	0.9985	−0.3197	0.4473
θ_{50}	−0.8642	0.3819	−0.6652	0.3976	−0.6168	0.1921
θ_{51}	2.4651	0.0094*	0.8074	0.0781	0.0375	0.9423
θ_{52}	−1.9186	0.0651	−0.0967	0.9231	−0.5576	0.3213
θ_{53}	0.3794	0.6437	0.9789	0.1765	0.8113	0.0289**
ω_0	0.0000012	0.1031	0.000021	0.0278**	0.0000021	0.0421**
ω_1	0.1094	0.0000*	0.0613	0.0361**	0.1834	0.0003*
ω_2	0.9662	0.0000*	0.9428	0.0000*	0.8479	0.0000*
ω_3	−0.3112	0.0000*	0.2573	0.0001*	0.0058	0.9137
Model diagnostics test on the standardized residuals						
	Stats	p-value	Stats	p-value	Stats	p-value
LB Q (6)	9.3726	0.154	1.8734	0.8532	7.7137	0.045**
LB Q^2(6)	5.9817	0.425	4.8725	0.7653	8.4328	0.192
ARCH(6)	6.0325	0.419	3.1623	0.821	9.8382	0.106

Note: LB Q(6) LB Q^2(6) are the Ljung-Box statistics applied on the standardized and squared standardized residuals, respectively.

ARCH(6) is the statistics used to test whether standardized residuals exhibit the ARCH effect up to the order. 6.

* and ** indicates that the statistic is significant at the 1% and 5% levels, respectively.

First, the result of cointegration test shows that the stock index spot is cointegrated with index futures and a set of macroeconomic factors for the pre-QFIIs, post-QFIIs and whole periods, suggesting some market inefficiency exists in Taiwan.

Second, the lead behavior is examined by means of the vector error correction model. As a result, the evidence shows unidirectional causality running from the spot market to futures market for the pre-QFIIs period. Nevertheless, after the opening up of the futures market to QFIIs, the empirical result shows that there exists bi-directional Granger-causality between index futures and index spot, but the causality from futures to spot is stronger. Therefore, it can be concluded that the opening up of the local futures market to QFIIs has indeed improved the lead of the Taiwan futures market.

Third, the finding shows that the inception of informed QFIIs trading alters the mechanism of the local futures index volatility. The evidence shows that the level of futures price volatility increases in the post-QFIIs period, suggesting that the opening up of the local futures market to QFIIs may enhance information flows running from the futures to spot markets. However, the asymmetric effects increase for the post-QFIIs period under investigation, implying that the introduction of informed QFIIs does not reduces the asymmetric response behavior. The increased asymmetric effects may result from the fact that the major proportion of investors on the TAIFEX is individual traders who are poorly trained and more inclined to overreact to bad news even if the QFIIs are introduced in Taiwan futures market during our sample period.

Overall, the above evidence documents that the introduction of well-informed QFIIs improves the efficiency of information transmissions of the local futures market, suggesting that deregulation is appropriate.

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