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Characterizing information flows among spot, deliverable forward and non-deliverable forward exchange rate markets: A cross-country comparison ☆

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Using an innovative GMGARCH-MSKST model that allows for asymmetric generalized dynamic conditional correlation, this paper analyzes return and volatility interactions among spot, non-deliverable forward (NDF) and deliverable forward (DF) exchange rate markets for Korea and Taiwan. With the backdrop of these two very different regulatory and institutional regimes we examine how the inter-temporal dynamics of forward-directed currency market instruments are both influenced by, and influence, spot market exchange rates.

1. Introduction

Most research on exchange rate market behavior tends to focus on the spot market and its relationship to price revelation mechanisms operating in the deliverable forward (DF) markets. However, there is increasing interest in exploring non-deliverable forward (NDF) market behavior and its relationship to returns and volatility in the spot market. Despite targeting the same currency, DF and NDF markets exhibit important differences. DF pricing is theoretically governed by interest-rate parity conditions in which

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equivalent returns are measured over a set period according to their respective interest rates and current spot market exchange rates (Akram et al., 2008; Baba and Packer, 2009; Della Corte et al., 2011; Ahmad et al., 2012). Specifically, the forward premium is dictated by the interest rate differential between the two currencies. In most cases, covered interest rate parity holds true because it reflects a no-arbitrage condition between onshore and offshore currency markets in the absence of capital controls. However, when some forms of capital controls are adopted, non-residents may be restricted from full access to onshore currency markets. Dooley and Isard (1980) report that deviations from covered interest parity conditions can be partially explained by the existence of capital controls. Furthermore, Frankel and MacArthur (1988) discovered that the covered interest rate parity condition does not always hold for emerging market currencies. In order to more fully explore possible currency market misalignment with the covered interest rate parity condition (Fama, 1984), our research proposes a direct market information approach to gauge the influence of offshore non-deliverable forward (NDF) markets on onshore currency markets. Offshore trading has emerged as an important market phenomenon that is reflected in an increasing number of offshore foreign exchange transactions in NDF markets. This is particularly apparent in the cross-border flow of investment and trade in emerging economies, such as Asia, Latin America, and Eastern Europe, where restrictions on currency convertibility and controls on capital remain in effect (McCauley and Scatigna, 2011). This has led monetary policymakers to speculate that price movements in the NDF market could be a useful tool to monitor market expectations and uncover information about percolating pressure on an exchange rate regime that may not be fully manifested through traditional tools available in a country with capital controls.

In contrast to the central role of interest rate parity in pricing most forward currency exchange contracts, there are several conflating factors that undergird NDF pricing. These include such things as structural market frictions, trading restrictions, market segmentation, margin payment requirements, and market regulations. One of the key differences between the NDF market and more traditional DF forward contracts is that NDF markets are cash settled. This cash settlement feature reflects the fact that NDF markets are contracted offshore and is virtually beyond the direct jurisdiction purview of domestic monetary authorities that seek to impose currency convertibility restrictions. While offshore investors have only limited access to onshore interest markets, NDF prices are derived mainly from expectations regarding future spot exchange rates (Lipscomb, 2005). Recent estimates have indicated that a large proportion of NDF transactions are generated by speculative interest, and this purportedly reflects the growth of international hedge funds participating in NDF markets. NDF markets are structurally consistent with trades that both limit the exposure inherent in hedging strategies as well as attract traders speculating on currency movements.

Kong and Shao (2010) reported that DF markets are more heavily constrained by trade regulations and cost considerations, resulting in a slow reaction to information from NDF markets. Conversely, NDF markets are expected to react rapidly to information from DF markets due to the flexibility they provide in aligning price expectations with new market information. Discrepancies in pricing have been linked to the degree to which regulations impede the flow of information between markets. The less restrictive environment of NDF markets appears to incorporate a more robust set of information into currency exchange markets, suggesting the greater the extent to which NDF prices reflect spot prices. Maziad and Kang (2012) reported that developments in offshore markets could influence onshore spot markets in terms of level and volatility during periods of offshore market dislocation. They attributed this to the generally held belief by market participants that prices in the offshore market are likely to better reflect global market conditions. From a policy execution perspective, where NDF market returns exhibit a leading role in spot market pricing, Central Bank foreign exchange policy that is triggered by movements in spot market prices may be misaligned with the exchange rate generation mechanism that is actually occurring in the market. It is our hypothesis that policymakers and investors who employ global hedging strategies will benefit from a better understanding of the complex information flows in exchange rate markets that are governed by different regulatory regimes and institutional structures. Increasingly sophisticated empirical methods can be used to evaluate information transmission mechanisms that underlie price movement in exchange rate markets, however, such innovations are still quite rare in capturing a more comprehensive view of information flows between spot, DF and NDF currency markets.

With increased capital flows to emerging countries in the 1990s, the NDF market became more popular with Asian countries. In contrast to standard forward markets, Asian NDF markets are generally developed in offshore financial centers, primarily located in Singapore and Hong Kong, followed by London and Tokyo. An

Emerging Markets Traders Association (EMTA) survey shows that the turnover in Asian NDF markets has grown significantly over the past decade.[1] Historical data suggests that Asian NDF markets accounted for about 70% of NDF turnover globally in 2003 and 2004 (Ma et al., 2004). According to data reported by the Tokyo Foreign Exchange Market Committee (TFEMC), the average daily Asian NDF turnover increased rapidly from about US $2.3 billion in the middle of 2003 to US$54.4 billion in April 2007.[2] Even though there was a general downward movement in NDF turnover in the middle of 2008, trading activity in NDF markets picked up after the first half of 2009. While NDF trade is increasing generally within Asian NDF markets, the Korea and Taiwan NDF markets are regarded as the most active in the region. In particular, the Korean NDF market is considered to be the largest and most liquid in the world while the NDF market for the New Taiwan dollar (NTD) is the second most active in Asia (Debelle et al., 2006).

As a result of their significant influence in the region, this paper is focused on exchange rate markets in Korea and Taiwan. Both countries have experienced remarkable economic growth over the past few decades and provide an opportunity to distil important information about how differing institutional contexts affect NDF markets as they take on increasing importance in the facilitation of trade and currency speculation. In both countries, policymakers have implemented policies designed to facilitate cross-country investing and international capital flows and both countries have experienced rapid accumulation of foreign exchange reserves. However, we also note that both countries have taken different policy paths to financial openness. Following the Asian financial crisis in 1997, the Korean government has been more aggressive in adhering to IMF recommendations to eliminate restrictions on capital inflows and restraining government involvement in the market. Korean currency market liberalization was aimed at removing unnecessary regulations on transactions related to foreign exchange, either directly or indirectly, as well as improving the efficiency of market exchange in terms of liquidity and market participation (Chung et al., 2000). While foreign exchange dealings in the past had to be based on actual demand, speculative forward transactions were permitted. As an example, Korean regulations allow onshore entities greater access to offshore NDF markets and greater access to onshore counterparties for offshore contracts.[3] In contrast, the central bank of Taiwan has been more concerned about maintaining financial stability and preventing speculative activity, thus imposing more restrictions on the currency market and capital movement.[4] Furthermore, Taiwan regulations restrict non-residents from participating in the onshore deliverable forward markets and prevent onshore banks from booking some classes of NDF trades.

Given the differing regulatory policy frameworks that govern financial openness for these two countries, we investigate information flows between offshore and domestic currency markets. The empirical focus on Korean and Taiwan markets provides a consistent contextual foundation for evaluating and comparing exchange rate structures under different policy regimes. Understanding the nature of exchange rate market dynamics is critical for policymakers and market participants who are increasingly impacted by emerging markets in the Asian region. Our study will also assist policymakers who seek a better understanding of the relationship between offshore NDF markets and foreign exchange rate spot markets. This will prove particularly helpful as policymakers in emerging market economies consider capital controls and policies that support a convertible exchange rate regime.

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1 For details on the survey please refer to the EMTA website at http://www.emta.org/.

2 The Tokyo Foreign Exchange Market Committee's survey for the Japanese market is available at http://www.fxcomtky.com/index_e.html.

3 Although the accessibility of onshore counterparties to offshore NDF contracts has likely increased the "thickness" of Korean won currency markets, many structural differences remain between Korea's offshore and onshore markets. These structural differences are likely to result in potentially different price discovery mechanisms in the domestic and overseas exchange rate markets. For example: (1) Capital controls: At present, Korean markets are still constrained by a number of regulations aimed at limiting the transfer of speculative investment into the domestic market. The Korean government continues to employ macro-prudential measures that include supervisory action toward the position of derivatives, such as NDFs in Korean banks. (2) Intervention by the Bank of Korea: The Bank of Korea occasionally intervenes in the won market, resulting in a price difference between domestic and overseas exchange rates. (3) Substantial difference in characteristics of market participants and risk/cost structures: Most participants in NDF markets are international financial institutions with high credit ratings. Since NDF transactions generally only involve cash settlements these transactions result in lower credit risk and transactional costs than those in DF markets where local institutions must undertake capital settlement.

4 According to the Global Competitiveness Report 2010–2011 rankings, published by the World Economic Forum (Schwab, 2010), Taiwan ranks very poorly in terms of "restriction on capital flows," (pp. 317). Additional information on Taiwan's financial sector openness is available at Hwang et al. (2011).

2. Characteristics of exchange rate markets

Existing research on exchange rate markets generally finds evidence of a price transmission mechanism between spot and DF markets. However, little attention has been paid to the discovery of the dynamic properties of NDF markets (Park, 2001).[5] It is well established in the literature that empirical investigation of asset price behavior requires complete and comprehensive modeling of all relevant price transmission channels (Ehrmann et al., 2005). In this vein, the paper represents a substantial step forward in applying empirical modeling tools that explicitly account for information transmission linkages between spot, DF and NDF markets. Further, since one might expect that arbitrage of exchange rates across markets should result in volatility spillover between markets, our modeling framework allows us to address the cross-market volatility spillover transmission process as well.[6] This issue is of particular interest to financial analysts and practitioners because the direction of volatility spillovers conveys information about investment risk and signals the extent to which exchange rate markets show a greater tendency to transmit volatility when they are in an active phase. Insight on this subject also has implications for policymakers tasked with developing more informed decisions regarding the cross-border integration of exchange rate markets. In particular, as monetary policymakers attempt to stabilize foreign exchange markets, a practical implication drawn from our analysis is whether, and to what extent, a shock from an offshore NDF market triggers volatility in domestic exchange rate markets.[7] It is, therefore, critical for policymakers and practitioners working in emerging markets to understand the behavior of NDF markets and the potential impact of such markets on the onshore markets beyond what is conveyed by conventional economic fundamentals.

There are many reasons to suspect that exchange rate returns and volatility are impacted by economic fundamentals, in addition to spillover interactions within spot, NDF, and DF markets. In her overview of NDF market fundamentals, Lipscomb (2005) alludes to the importance of interest rate differentials and the suite of investor opportunities in related markets. Moreover, by evaluating the relative innovations from Japanese and US markets on Pacific-Basin countries as regional and world shocks respectively, Bekaert and Harvey (1997) and Ng (2000) found that there are significant regional spillovers from the Japan market to many of the Pacific-Basin country markets. [8] Gu and McNelis (2013) not only found that the Japanese yen directly and significantly influences the onshore won spot currency market, but also showed that the yen exerts strong direct pressure on Chinese financial markets through speculative movements in the offshore NDF markets. Their results show that RMB NDF market movements were driven by speculative psychology generated by the yen market, suggesting that the RMB NDF market plays a key role in transmitting information from the yen to RMB spot and financial markets. In addition, there may also be evidence of an error correction process imbedded in the exchange rate structure itself. We are interested in the question of whether return and volatility linkages between spot, NDF and DF markets remain after controlling for the impact of economic fundamentals in the exchange rate determination process as well as an error correction adjustment process.

Based on the nature of volatility observed in high frequency time series data, several studies (Wang et al., 2001; Zhong et al., 2004) document asymmetric responses in volatility resulting from both positive and negative shocks. It is generally recognized that volatility is typically more likely to rise during periods of falling prices, and fall during periods of rising prices. [9] While asymmetry is generally found in the volatility of financial asset data series, there has been little attention paid to the volatility dynamics of exchange rates.

[5] Park (2001) investigated the relationship between won–dollar spot and offshore NDF markets and reported evidence of information flows between the two markets. In addition, his research suggested that regulatory reform has changed the direction of the dynamic relationship.

[6] Modeling the volatility component of interrelated markets explicitly accounts for arrival information flows and their associated assessment. Ehrmann et al. (2005) suggest that the causal relationships of volatility within and across markets provide keen insight into the dynamic structure of price revelation mechanisms.

[7] One of the reasons why capital flows from the NDF market to the spot market may induce more volatility in the spot market is that NDF markets may make it easier for sizable speculative positions to be accumulated.

[8] Pacific-Basin currency markets have recently experienced dramatic liberalization events such as capital market reform that have served to facilitate increases in cross-country investing over time. If liberalization or deregulation enhances information flows between currency markets, one would expect stronger spillover over effects from the global markets (US), and regional markets (Japan), after the institutional changes have taken place.

[9] There are two popular theories documenting the potential source of asymmetric volatility of equities: the leverage effect (Black, 1976) and the volatility feedback effect (French et al., 1987).

Within this literature, Maya and Gómez (2008) found that the response of exchange rates to volatility shocks is characterized by symmetry in most Latin American emerging markets. Wang and Yang (2009) found that currency depreciation against the US dollar leads to significantly greater volatility than an appreciation for Australian dollar and British pound, whereas the opposite is true for Japanese yen. McKenzie (2002) argues that the presence of asymmetric responses to exchange rate volatility is attributable to intervention activity of the central bank. Fang et al. (2009) suggest that response asymmetries may arise from a host of different sources, including asymmetric risk perceptions by exporters, hedging behavior, US-dollar invoicing, and unsophisticated foreign exchange market intervention.

Another issue that has received increasing attention in the financial economics literature is the time-varying behavior of correlation and covariance in exchange rate markets. As these correlations are crucial inputs for international portfolio management and risk management decisions, understanding the behavior of time-varying structures between asset returns is crucial for investors. There are numerous studies that document time varying correlation. Specifically, Engle (2002) employs a multivariate dynamic conditional correlation GARCH (DCC-GARCH) model to estimate conditional correlation coefficients simultaneously with the conditional variance–covariance matrix. Cappiello et al. (2006) use an Asymmetric Generalized Dynamic Conditional Correlation GARCH model that facilitates the evaluation of asymmetric response of asset correlations to negative shocks. Recently, Li (2011) explored this type of asymmetry in exchange rate correlations and found that currencies co-move more closely during joint appreciations than joint depreciations. These different types of correlation asymmetries provide important financial implications for portfolio diversification and hedging strategy. By relaxing constraints on the symmetric covariance specification we expect to create a more finely focused picture of underlying exchange rate covariance dynamics.

Finally, there is a large volume of literature that challenges the adequacy of employing an assumed distribution for univariate GARCH models. In spite of this, it is still rare to find models that address proper parametric modeling in a multivariate framework. To better model the characteristics of exchange rates processes, we use a quasi-maximum likelihood GARCH estimation approach that employs the multivariate skewed-Student's t distribution (MSKST) proposed by Bauwens and Laurent (2005). This model structure addresses problems of non-normality and skewness in the return distribution. This modeling framework also provides a more general parametric specification that is capable of adapting to complexities in the exchange rate data generating process—in particular when data are observed with high frequency and exhibit non-Gaussian attributes.

3. Data description

Daily bilateral exchange rate spot prices for the Korean won, New Taiwan dollar (NTD) and Japanese yen are defined as units of local currency per US dollar. Spot currency prices reflect daily closing price. Both DF and NDF prices reflect the closing prices on a one-month forward contract. Exchange rate price and return data for markets in Korea and Taiwan are plotted in Fig. 1 for the period from the March 1, 2001 to Feb, 28, 2011.

One-month forward contracts are employed because they have the largest trading volume among the different maturities. Data for the Korean won and Taiwan NTD and their associated NDF and DF markets are taken from the Reuters database. We used Korea Composite Stock Price Index (KOSPI) and Taiwan Stock Exchange (TSE) composite index as stock market price data; and, Taiwan and Korea interbank offered rate relative to USD LIBOR rate as proxies for short-term nominal interest rate differentials. The Japanese yen spot price, stock index data and interest rate data are obtained from the DataStream database.

Since evidence suggests that financial time series data are typically characterized by non-stationary properties, an Augmented Dickey–Fuller (ADF) test is performed to examine the presence of a unit root in each raw data series. To preserve parsimony, selection of the optimal number of lags is based upon AIC (Akaike's Information Criterion) and SC (Schwartz's Criterion) criteria. The ADF statistics for spot, NDF, DF, interest rate, stock index, and Japanese yen level data are summarized in Table 1. All the ADF statistics, except the short term interest rate differentials variables, fail to reject evidence of unit-root phenomena. This result indicates that the data series—in the levels—exhibit non-stationary rather than mean-reverting properties. To achieve stationarity, all the data except interest rate differential variables are transformed from levels to rates of change by taking the first difference of the natural log. After transforming the data, ADF statistics (shown as ΔADF in the table) reject unit roots at the 1% level of significance. All subsequent analysis employs the transformed rate of return data as the unit of measurement in the market.

A wide range of descriptive statistics for the spot, NDF, and DF data series are reported in Table 2. Taken as a whole, the descriptive statistics suggest that daily exchange rate returns for the selected Korean and Taiwan foreign exchange markets deviate from conventional Gaussian assumptions and clearly reject an assumption of independence. These results are consistent with the empirical findings contained in Wang et al. (2001), and suggest a need for an appropriate ARMA-GARCH type specification that can correct the dependence structures existing in the first and second moment sequences. Daily price and return data plots presented in Fig. 1 suggest a sharp depreciation in exchange rates for both Korea and Taiwan beginning from July 2008 to March 2009. [10] At the bottom of Table 2, the results of Chow breakpoint tests suggest that the null of no structural change for the crisis period (July 2008 to March 2009) is rejected for spot, DF and NDF markets.[11] As such, we designate the structural change event period from July 2008 to March 2009—and refer to it as the crisis period. One of the features in our paper is to simultaneously model the crisis period breaks in the mean, variance and covariance structures.

4. Model specification

In specifying our econometric model, we take five primary issues into consideration: (i) multivariate interactions within spot, NDF, and DF markets that include an error correction process as an explanatory variable for both the conditional mean and the conditional variance; (ii) asymmetry in the volatility dynamics of exchange rates; (iii) several different aspects of asymmetry characteristics in exchange rate correlation processes; (iv) the subprime crisis period as a structural break in the mean, variance and correlation structures; and (v) potential non-normality in exchange rate regression errors. We tackle each of these modeling issues in turn as we introduce our model specification framework.

To facilitate a comprehensive examination of dynamic volatility interrelationships between spot, NDF, and DF markets, our paper adopts a multivariate asymmetric GJR GARCH modeling framework.[12] This framework not only captures own-market volatility asymmetry (i.e., volatility response asymmetries toward positive and negative shocks of the previous period) but also cross-market volatility asymmetry (i.e., asymmetric transmission of cross-market volatility to positive and negative shocks exhibited in other markets). The transmission mechanism of cross-market volatility is of critical importance for risk management strategies in markets that show a greater tendency to transmit volatility when they are in an active phase.

As a further step toward capturing a more comprehensive understanding of the information flows between offshore and domestic exchange rate markets we explicitly model short-run disequilibrium in the first and second moments.[13] Following Lee's (1994) GARCH-X framework—and the subsequent contributions by Choudhry (1997), Zhong et al. (2004) and Sakthivel et al. (2012)—we allow shocks to the system to propagate through the first and second moments. To investigate how exchange rate markets are impacted by short-run price disequilibrium our model employs a system of error correction models

[10] This specification of the crisis period corresponds to that used by Fratzscher (2009) and Aizenman and Hutchison (2012).

[11] Results from Chow breakpoint tests suggest that the null of no structural change for the crises periods could be rejected for all exchange rate markets. The Quandt–Andrews unknown breakpoints test (Andrews, 1993) also suggests rejection of the null of no structural changes for all exchange rate markets. Due to space considerations, the results for these tests are not reported here. They are available to readers upon request.

[12] The GJR-GARCH model has been demonstrated to be superior to other alternative competing asymmetric conditional variance specifications in some studies (Engle and Ng, 1993; Kim and Kon, 1994). Research by Antoniou et al. (1998) and Butterworth (2000) showed that the GJR-GARCH model outperformed EGARCH and APGARH models in capturing the dynamic behavior of exchange rate markets. Demonstrating superiority of a GJR-GARCH model in capturing volatility behavior for managed float currencies suggests that application of the model to Korean and Taiwan exchange rate markets is efficacious since the monetary authorities seem to be more inclined to operate according to managed float regimes during our study period.

[13] As Engle and Granger (1987) point out, if a given market's non-stationary sequences have relationships of cointegration, one should take corrections of short-term disequilibrium into consideration. Ignoring the presence of cointegration may result in inappropriate estimations due to mis-specification of pricing behavior between asset markets. We utilize a multivariate GARCH-X specification, first introduced by Lee (1994), to include an error correction term from a cointegrating type relationship for the underlying vector process as an explanatory variable in the conditional covariance matrix. Examining the behavior of the variances over time as a function of short-run deviations is reasonable when one expects increased volatility due to shocks to the system which propagate on both the first and the second moments.

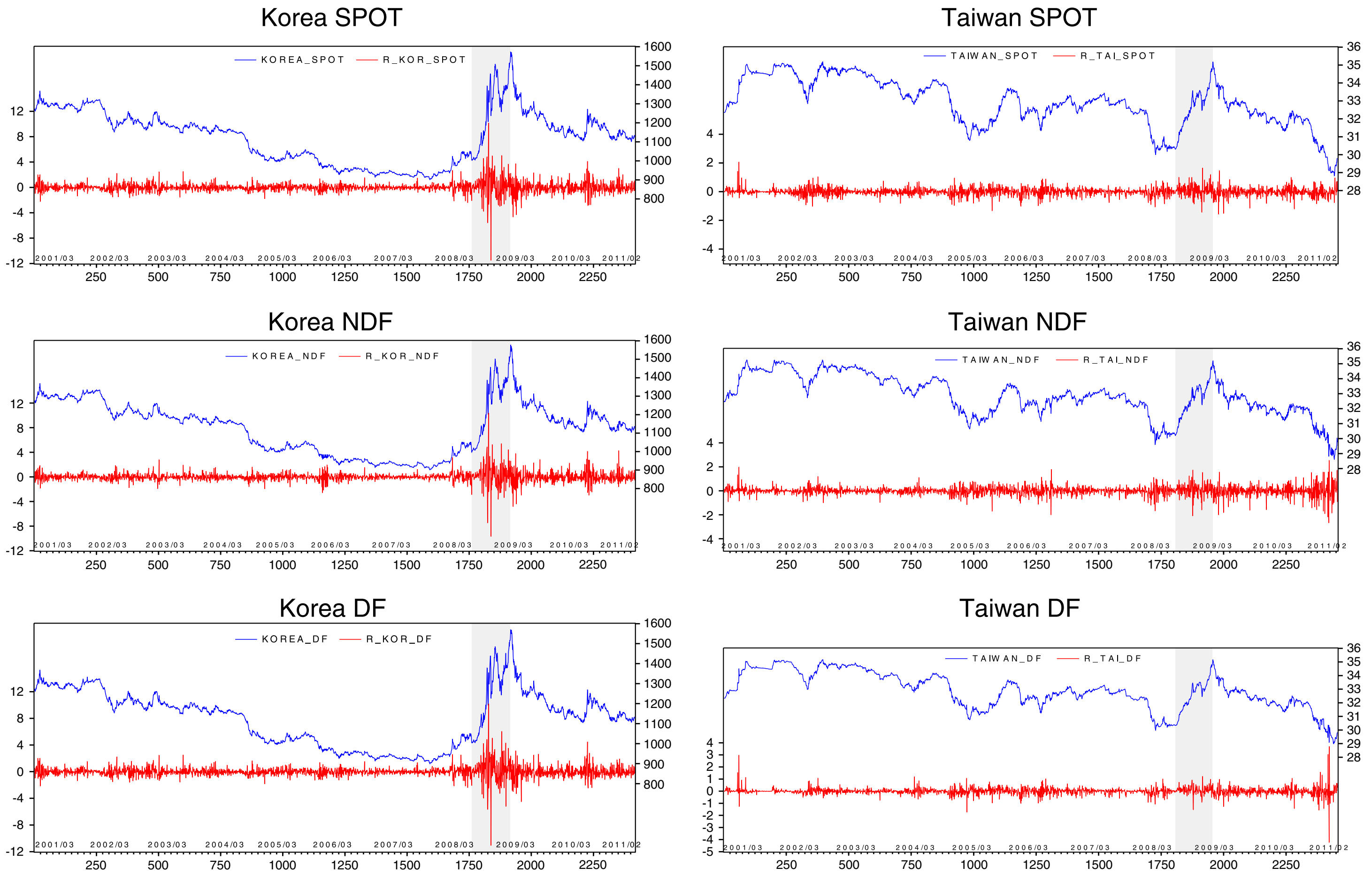

Fig. 1. Daily price and return for SPOT, NDF and DF markets: the plots show exchange rates on daily price and daily return for SPOT (top), NDF (middle), and DF (bottom) respectively. The shading area indicates the sub-prime crisis periods during July 2008 to March 2009.

Table 1
Augmented Dickey–Fuller unit root tests.

	Korea			Taiwan		
	SPOT	NDF	DF	SPOT	NDF	DF
ADF	−0.4670	−0.2553	−0.3513	−1.0057	−0.8413	−0.9508
	<9>	<9>	<10>	<6>	<5>	<5>
ΔADF	−31.035***	−30.772***	−30.959***	−48.784***	−40.881***	−19.434***
	<2>	<2>	<2>	<0>	<1>	<4>
	STOCK	RS	YEN	STOCK	RS	YEN
ADF	−1.2957	−4.6934***	−1.4544	−1.7955	−2.0602**	−1.1530
	<0>	<3>	<0>	<1>	<2>	<0>
ΔADF	−47.917***		−52.108***	−46.724***		−52.565***
	<0>		<0>	<0>		<0>

ADF is the Augmented Dickey Fuller (ADF) unit root test statistic in the level data, while ΔADF represents ADF statistics after undergoing natural log transformations and taking the first difference. For a discussion of the selection of their critical values, see MacKinnon (1996). Numbers in < > reflect the optimum lag taken according to AIC rules. The symbols ** and *** indicate significance at the 5% and 1% level, respectively.

(ECMs) for the conditional mean and an extended trivariate GARCH model with an error correction term for the conditional variance. This provides us with a useful tool to monitor market expectations and explore the potential pressure of disequilibrium forces on offshore and domestic currency market structures.

Our research also seeks to underscore the descriptive importance of both leverage and cross-asymmetry effects in exchange rate covariance dynamics. Within the generalized asymmetric dynamic conditional correlation (GADCC) framework we partition the combinations of asymmetric shocks into three quadrants:

Table 2
Summary statistics for daily SPOT, NDF, and DF market returns.

	Korea			Taiwan		
	SPOT	NDF	DF	SPOT	NDF	DF
Descriptive statistics						
Mean	−0.0047	−0.0049	−0.0047	−0.0033	−0.0035	−0.0032
Standard deviation	0.8210	0.8050	0.8116	0.2801	0.3796	0.2978
Skewness[a]	−0.1444#	0.3816#	0.0421	0.0987#	−0.3945#	0.3110#
Kurtosis[a]	35.142#	31.738#	35.544#	7.6896#	9.8773#	40.1627#
J-B test[b]	103969.2	83162.3	106579.1	2252.7	4899.9	141254.1
	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]
LB-Q(10)[c]	63.737	49.108	67.237	21.168	54.135	68.998
	[0.000]	[0.000]	[0.000]	[0.020]	[0.000]	[0.000]
LB-Q^2(10)	640.26	605.86	618.06	608.44	556.57	626.37
	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]
Chow breakpoint test						
F-statistic	11.2695***	10.3496***	11.0072***	3.8664***	10.8137***	12.7536***
	[0.0000]	[0.0000]	[0.0000]	[0.0007]	[0.0000]	[0.0000]

Values in brackets represent the P-values of estimates and the symbol *** indicates significance at the 1% level.

[a] The calculation method for quasi standard deviation of skewed and kurtosis coefficients are $\sqrt{6/T}$ and $\sqrt{24/T}$, where T is the sample number. Values in parentheses represent the standard deviation of estimated parameters; Under the normal distribution, the skewness is zero; the kurtosis coefficient is 3; #means skewness or kurtosis coefficient is over two times larger than the quasi standard deviation.

[b] The J–B test represents results for the Jarque–Bera normal distribution test. Values in brackets represent the P-values of estimates.

[c] LB-Q(10) and LB-Q^2(10) express the heteroscedasticity-consistent Ljung–Box Q(10) statistics used to test the joint significance of the serial correlations up to 10 lags for standardized residuals and their associated squared standardized residuals, respectively.

joint negative (appreciation) shocks of the same direction [−,−]; joint positive (depreciation) shocks of the same direction [+,+]; and one positive (depreciation) shock and the other negative (appreciation) shock [[−,+], [+,−]]. Ignoring the features of joint positive, joint negative and cross asymmetries in covariance dynamics would lead to underestimating the correlation and understating the benefits of diversification. We also incorporate a structural change variable into the GADCC specification that facilitates exploration of whether the crisis period played a significant role in explaining correlation changes between exchange rate markets.

Furthermore, given the prevalence of fat tail and skewness properties observed in the Korean and Taiwan exchange rate return data, we relax the usual multivariate normal distribution assumption. The proposed modeling framework, based on the MSKST density (Bauwens and Laurent, 2005), is sufficiently general to capture the statistical features of skewness and kurtosis reflected in the exchange rate return data used in our application. The tri-variate skewed-Student's t MGARCH model based on GADCC covariance dynamics (termed as GMGARCH-MSKST) used in our analysis is presented below in terms of conditional mean equations, conditional variance equations, conditional covariance equations, and the multivariate probability distribution.

4.1. Conditional means

The conditional mean equations are as follows:

$$\phi^i_{m^i}(B)R^i_t = c^i + \sum_{l=1}^{s^j}\psi^{j_i}_l R^j_{t-l} + \sum_{l=1}^{s^k}\psi^{k_i}_l R^k_{t-l} + \sum_{l=1}^{s^{RS}}\gamma^i_l RS_{t-l} + \sum_{l=1}^{s^{stock}}\eta^i_l Stock_{t-l} + \sum_{l=1}^{s^{jp}}\varpi^i_l R^{jp}_{t-l} + \iota^i_R DS_t + \tau^i_R Z^i_{t-1} + \theta^i_{n^i}(B)\varepsilon^i_t \quad (i,j,k = spot, ndf, df; i \neq j \neq k). \tag{1}$$

Eq. (1) represents the structural conditional mean equations in spot, NDF, and DF markets, respectively. R^{spot}_t, R^{ndf}_t, and R^{df}_t represent returns for spot, NDF, and DF markets, respectively; ε^{spot}_t, ε^{ndf}_t, and ε^{df}_t are the residuals of the corresponding market's mean equation; ϕ and θ are estimated parameters of the AR and MA process from the ARMA model, respectively. This paper relies upon the three principles proposed by Wang et al. (2001) to determine the optimal lags for rates of return in each market.[14] The parameters ψ^{i_j} determine the effects of cross-market return transmission. For example, in Eq. (1) ψ^{ndf_spot} and ψ^{df_spot} represent estimations of cross-market return transmission signals from NDF markets and DF markets to spot markets, respectively. In consideration of the voluminous literature on relevant structural linkages in exchange rate markets we limited our focus to modeling three core structural linkages, including: interest rate differentials (RS_t), stock price changes ($Stock_t$)[15], and spillovers from large regional currency markets (R^{jp}_t).

In our model, the Japanese yen is identified as the large regional currency that is creating spillover effects on the won and NTD markets. As context for this element of the model it is important to note that Japan accounts for the third largest share of imports for Korea and second largest share of imports for Taiwan. As such it is likely that currency movements in the Japanese yen are likely to influence Korean won and Taiwan NTD currency market movements. Also, we would expect that currency traders and policy makers will be sensitive to understanding the dynamic behavior of Japanese yen for pricing domestic securities, implementing global hedging strategies and how market participants are making asset allocation decisions.

[14] First, using the Box–Jenkins (Box and Jenkins, 1976) method, autocorrelation functions, and partial autocorrelation functions are specified. Second, we select from among this set of possible models using the Box–Jenkins rule, further selecting models in which the P-values of Ljung–Box (10) statistics are greater than 0.1, to ensure that residual terms carry white noise from the first moment sequence. Finally, having used these two rules to filter possible models, AIC and SC rules are adopted to select the most parsimonious models to be the basis of setting conditional mean equations.

[15] Kohler (2010) documented that interest rate differentials played an important role in explaining exchange rate movements during periods of financial crisis. Increasing activity in carry trade could be a potential source for this finding. From the standpoint of asset selection, the choices that people make with regard to domestic and foreign assets also has an influence on the determination of exchange rate returns.

Because of the close relationship between the Japanese yen and other Asian currencies we explore whether offshore NDF markets—that are characterized by more speculative trading—tend to be more sensitive to valuation changes in the yen than onshore DF and spot markets. In addition, the dummy variable DS_t is assigned a value of one for the period July 2008 to March 2009, referred as the crisis period; otherwise DS_t is zero.

Under conditions of long-term equilibrium among spot, NDF, and DF markets, the error correction terms (Z^i_{t-1}) are specified by $Z^i_{t-1} = EX^i_{t-1} - \vartheta_0 - \vartheta_1 EX^j_{t-1} - \vartheta_2 EX^k_{t-1}$, where $i, j, k =$ spot, NDF, and DF, respectively, and $i \neq j \neq k$; and EX^i_{t-1} is the logarithm of the level of the exchange rate.[16] These are incorporated into a market's first and second moments to describe the dynamic influence of short-term correction on market returns and volatility. The significance of the estimated parameter, τ^i_R, is used to explore the importance of disequilibrium in explaining the propagation of shocks through the first moments. Specifically, the τ^i_R parameters inform the speed of adjustment in the presence of deviation from long-run return equilibrium. The sign of the τ^i_R parameter tracks the short-term adjustment of long-run equilibrium in the exchange rate microstructure and is expected to be negative in the return equation due to arbitrage effects. While Zhong et al. (2004) mention that the sign of the error correction term in the return equation depends on the net outcome of the "arbitrage" and "momentum" effects, we attempt to empirically examine how traders respond to exchange rate disequilibrium.

4.2. Conditional variance

The conditional variance equations are specified as follows:

$$h^i_t = g^i + \sum_{l=1}^{u^i} \beta^i_l h^i_{t-l} + \sum_{s=1}^{v^i} \lambda^i_{1s}\left(\varepsilon^i_{t-s}\right)^2 + \lambda^i_2 D^i_{t-1}\left(\varepsilon^i_{t-1}\right)^2 + \tau^i_h\left(Z^i_{t-1}\right)^2 + \iota^i_h DS_t$$
$$+ \mu^{j_i}_1\left(\varepsilon^j_{t-1}\right)^2 + \mu^{j_i}_2 D^j_{t-1}\left(\varepsilon^j_{t-1}\right)^2 + \mu^{k_i}_1\left(\varepsilon^k_{t-1}\right)^2 + \mu^{k_i}_2 D^k_{t-1}\left(\varepsilon^k_{t-1}\right)^2 \quad (2)$$

where the dummy variable $D^i_t = 1$ if exchange rate market i at time t exhibits a negative shock; otherwise $D^i_t = 0$, ($i, j, k = spot, ndf, df$; $i \neq j \neq k$).

Eq. (2) is the conditional variance equation for spot, NDF, and DF markets, respectively. In order to examine volatility spillovers, we adopt the asymmetric GJR GARCH specification that allows conditional variance (h^i_t) to be influenced by lagged square residuals ($(\varepsilon^i_{t-s})^2$), lagged conditional variance (h^i_{t-1}), and asymmetric volatility effects ($D^i_{t-1}(\varepsilon^i_{t-1})^2$). Under conditions in which the exchange rate spot market experiences unexpected negative shocks ($\varepsilon^i_t < 0$), the dummy variable is $D^i_t = 1$; if the exchange rate showed unexpected positive changes ($\varepsilon^i_t < 0$), then the dummy variable $D^i_t = 0$. Under this specification, λ^i_2 captures information about asymmetric volatility effects. Based upon the significance of the volatility transmission parameter, $\mu^{i\text{-}j}_1$, and asymmetric volatility transmission parameter, $\mu^{i\text{-}j}_2$, one can describe both cross-market volatility and asymmetry of transmissions.[17] This modeling framework is designed to enhance understanding of the information transmission process among markets and helps inform policymakers and market participants about the source of instability in foreign exchange rate markets.

Lee (1994) shows that when short-term disequilibrium becomes greater, market volatility increases. The τ^i_h parameters measure the impact of a deviation from long-run equilibrium on the subsequent conditional volatility; the sign of the error correction term τ^i_h is expected to be positive in the variance

[16] Preliminary analysis using unit root tests suggests that spot, NDF, and DF markets exhibit a long run relationship. Due to space considerations, the results are not reported here but are available upon request.

[17] In other words, when $\mu^{i\text{-}j}_2$ is significantly different from zero, unexpected negative shocks in market i will lead to a higher conditional volatility spillover to market j compared with positive shocks of currency market i in the previous period. When $\mu^{i\text{-}j}_1$ is significant but $\mu^{i\text{-}j}_2$ is insignificant, it means that there exists a volatility spillover mechanism operating from market i to market j, but the positive or negative shock is not clearly distinguishable. If $\mu^{i\text{-}j}_1$ and $\mu^{i\text{-}j}_2$ are both insignificant, then market i has no influence on the volatility of market j.

equation due to the belief that temporary mispricing should lead to higher volatility in a subsequent period. As in Eq. (1) we include a dummy variable DS_t that captures information on volatility breaks during the crisis period.

4.3. Conditional covariance

Conditional covariance equations are specified as follows.

$$\hat{\boldsymbol{\varepsilon}}_t|\Omega_{t-1} \sim MSKST(0, \mathbf{H}_t) \tag{3}$$

$$\mathbf{H}_t = \mathbf{D_t R_t D_t} \tag{4}$$

$$\mathbf{D}_t = diag\left(h_t^i, h_t^j, h_t^k\right) \tag{5}$$

$$\hat{\mathbf{Z}}_t = \mathbf{D}_t^{-1/2}\hat{\boldsymbol{\varepsilon}}_\mathbf{t} \tag{6}$$

$$\mathbf{R}_t = diag\{\mathbf{Q}_t\}^{-1/2}\mathbf{Q}_t diag\{\mathbf{Q}_t\}^{-1/2} \tag{7}$$

$$\mathbf{Q}_t = \left\{q_t^{ij}\right\} \tag{8}$$

$$\begin{aligned} q_t^{ij} = {} & \left(1-\alpha_3^{ij}-\alpha_4^{ij}\right)\overline{\rho}^{ij}-\left(\alpha_1^{ij}\overline{\rho}^{ij-}+\alpha_2^{ij}\overline{\rho}^{ij+}\right)+\alpha_1^{ij}\left(z_{t-1}^i D_{t-1}^i\right)\left(z_{t-1}^j D_{t-1}^j\right) \\ & +\alpha_2^{ij} z_{t-1}^i\left(1-D_{t-1}^i\right) z_{t-1}^j\left(1-D_{t-1}^j\right)+\alpha_3^{ij} z_{t-1}^i z_{t-1}^j+\alpha_4^{ij} q_{t-1}^{ij}+\alpha_5^{ij} DS_t \end{aligned} \tag{9}$$

$$\rho_t^{ij} = q_t^{ij}/\sqrt{q_t^{ii}q_t^{jj}}, \quad i,j = spot, ndf, df \tag{10}$$

The conditional covariance $\mathbf{H_t}$ represents the conditional covariance of spot, NDF, and DF market returns respectively. $\mathbf{D}_t$ is the 3×3 diagonal matrix of time-varying standard deviations from univariate GARCH models with h_t^i on the ith diagonal. $\mathbf{R_t}$ denotes the conditional correlation matrix for standardized return residuals $\{z_t^i\}$, whose elements, ρ_t^{ij}, represent a time-varying conditional correlation between z_{t-1}^i and z_{t-1}^j. Before estimating the parameters of $\mathbf{R}_t$, a standardization of residuals is required. Engle (2002) suggests setting the standardized residuals vector $\hat{\mathbf{Z}}_t$ at mean zero and variance one. $\mathbf{Q}_t$ is the conditional covariance matrix of the standardized residuals. In the model we allow for different asymmetric dynamics in the responses of correlations to information shocks. With this model structure we can explore the coexistence of positive and negative leverage effects proposed by De Goeij and Marquering (2004) and Li (2011), and also cross-asymmetry effects in exchange rate covariance dynamics.

Our approach makes an effort to comprehensively describe the full dimension of asymmetric covariance dynamics. As such the dynamic conditional correlation coefficient matrix, depicted in Eq. (9), can be divided into two parts: The first part, $\left(1-\alpha_3^{ij}-\alpha_4^{ij}\right)\overline{\rho}^{ij}-\left(\alpha_1^{ij}\overline{\rho}^{ij-}+\alpha_2^{ij}\overline{\rho}^{ij+}\right)$ represents the unconditional expectation of the cross product $z_t^i z_t^j$, where $\overline{\rho}^{ij} = E\left(z_t^i z_t^j\right)$; $\overline{\rho}^{ij+} = E\left[\left(z_t^i\left(1-D_t^i\right)\right)\left(z_t^j\left(1-D_t^j\right)\right)\right]$; $\overline{\rho}^{ij-} = E\left[\left(z_t^i D_t^i\right)\left(z_t^j D_t^j\right)\right]$; and the second part of Eq. (9), $\alpha_1^{ij}(z_{t-1}^i D_{t-1}^i)(z_{t-1}^j D_{t-1}^j) + \alpha_2^{ij} z_{t-1}^i(1 - D_{t-1}^i)$ $z_{t-1}^j(1 - D_{t-1}^j) + \alpha_3^{ij} z_{t-1}^i z_{t-1}^j + \alpha_4^{ij} q_{t-1}^{ij} + \alpha_5^{ij} DS_t$, indicates the conditional time-varying covariance. The α_1^{ij} parameters determine two previous negative shocks $((D_{t-1}^i * D_{t-1}^j) z_{t-1}^i z_{t-1}^j)$; the α_2^{ij} param eters show two previous positive shocks $((1 - D_{t-1}^i) * (1 - D_{t-1}^j) z_{t-1}^i z_{t-1}^j)$; the α_3^{ij} parameters examine how the degree of conditional covariance is influenced by the shocks of opposite signs $(z_{t-1}^i z_{t-1}^j)$. Finally, the parameters α_4^{ij} are specified to control for the effect of conditional covariance persistence. As mentioned before, we incorporate a structural break dummy DS_t in the conditional correlations to accompany periods of increased covariance induced risk. The coefficient α_5^{ij} is used to capture the structural change properties in the conditional correlations. We expect to observe a tighter interrelationship during the financial crisis period, anticipating the coefficient α_5^{ij} to be positive. Additionally, α_1^{ij}, α_2^{ij}, α_3^{ij} and α_4^{ij} are constrained to be nonnegative parameters that satisfy $\alpha_1^{ij} + \alpha_2^{ij} + \alpha_3^{ij} + \alpha_4^{ij} < 1$.

The GADCC (Generalized Asymmetric DCC) GARCH model is estimated in two steps: the first step is to estimate asymmetric dynamics in variances and then estimate parameters of dynamic correlations.[18] The maximum likelihood function is denoted by $MSKST\left(0, \mathbf{H}_i, \hat{\xi}, nu\right)$ and is represented by Eq. (11) below:

$$\begin{aligned} L_c(\Theta,\phi) = & -\frac{1}{2}\sum_{t=1}^{T}[K\ln((nu-2)\pi) + \ln\Gamma(nu) - \ln\Gamma(nu+K)] \\ & + \sum_{t=1}^{T}\left[K\ln(2) + \ln\left(\prod_{i=1}^{k}\frac{\xi_i}{1+\xi_i^2}\right) + \ln(\prod_{i=1}^{k}s_i)\right] \\ & - \frac{1}{2}\sum_{t=1}^{T}\left[\ln|\mathbf{H}_t| + (nu+K)\ln(1+\frac{\hat{z}_t'\mathbf{H}_t^{-1}\hat{z}_t}{mu-2})\right], \end{aligned} \tag{11}$$

where $\hat{z}_t = (z_1^*, \ldots\ldots, z_k^*)'$; $z_i^* = (s_i z_i + m_i)\xi_i^{I_i}$; $z_i = \varepsilon_t^i/\sqrt{h_t^i}$; $\hat{\xi} = (\xi_1 \ldots \xi_k)$;

$$\begin{aligned} m_i &= \frac{\Gamma((nu-1)/2)\sqrt{nu-2}}{\sqrt{\pi}\Gamma(nu/2)}\left(\xi_i - \frac{1}{\xi_i}\right); \\ s_i^2 &= \left(\xi_i^2 + 1/\xi_i^2 - 1\right) - m_i^2; \\ I_t^i &= \begin{cases} -1 & if \quad z_t^i \geq -m_i/s_i \\ 1 & if \quad z_t^i \leq -m_i/s_i \end{cases}. \end{aligned}$$

$\hat{\boldsymbol{\varepsilon}}_t$ and $\mathbf{H}_t$ represent the error terms and the variance–covariance matrix, respectively; the symbol Θ is used to represent the unknown parameters of $\hat{\boldsymbol{\varepsilon}}_t$ and $\mathbf{H}_t$; T represents the number of observations; and K represents the number of variates.[19] The shape parameter, degree of freedom, is denoted by nu ($nu > 2$), which can be seen as an indication of how pervasive fat tail properties are in the data's distribution; the vector of skewness parameters are denoted by ξ_i ($\xi_i > 0$), where the sign of ln ξ_i indicates the direction of skewness (i.e., the density is skewed right (left) if ln $\xi_i > 0(<0)$). In addition, m_i and s_i^2 are functions of $\hat{\xi}$ and nu and do not add additional parameters to the estimation.

5. Empirical results

We start our empirical analysis with a focus on model specification by estimating the MGARCH model using three different multivariate distributional specifications: normal (MN), student t (MST), and skewed student t (MSKST).[20] Log likelihood ratio tests, AIC and SC ranking of alternative distributional specifications suggest the following ranking from most descriptive to least: MSKST, MST, and MN. This result validates the value of accounting for asymmetric and leptokurtic properties in specifying the exchange rate return density function. Table 3 summarizes model diagnostics for the GMGARCH-MSKST specification, including heteroscedasticity consistent Ljung–Box-Q (LB-Q) statistics for the standardized residuals, its associated square, the product of standardized residuals, variance specification tests (Engle and Ng, 1993), and covariance specification tests (Koutmos and Knif, 2002). The LB-Q statistics for the standardized residuals and their squares are not significant at conventional levels, indicating that serial correlation has been adequately addressed in the model specification.

Estimation results reported at the bottom of Table 3, indicate that the shape parameter—degree of freedom (nu)—of the MSKST structure for both Korea and Taiwan is significant. In addition, most of the

[18] As mentioned in Engle (2002), two-step estimation results in some efficiency loss, but enhances the feasibility of large covariance matrix computation.
[19] In this article, we employ a tri-variate GARCH; thus K = 3.
[20] Due to space considerations, the results discussed in this paragraph are not reported here. They are available upon request.

Table 3
Diagnostic tests for the GMGARCH-MSKST models.

Diagnostic tests						
	Korea			Taiwan		
Standardized residuals (z_t^i) [1]	SPOT	NDF	DF	SPOT	NDF	DF
LB-Q(10)[a]	8.2008	7.5399	8.7145	16.230	7.8865	13.498
	[0.609]	[0.674]	[0.559]	[0.093]	[0.640]	[0.197]
LB-Q^2(10)[b]	5.8713	15.250	14.178	12.381	4.4034	3.8887
	[0.826]	[0.123]	[0.165]	[0.260]	[0.927]	[0.952]
Variance specification test (Engle and Ng, 1993)						
Sign test $\chi^2(1)$	0.0566	0.6606	0.2065	3.5052	0.5523	1.6709
	[0.8118]	[0.4163]	[0.6495]	[0.0612]	[0.4573]	[0.1961]
Negative size bias $\chi^2(1)$	2.1680	1.7531	0.0058	1.5531	1.6680	1.9039
	[0.1409]	[0.1855]	[0.9389]	[0.2127]	[0.1965]	[0.1676]
Positive size bias $\chi^2(1)$	7.9575	6.6498	1.0541	2.9041	1.5994	3.5282
	[0.0548]	[0.0599]	[0.3046]	[0.0884]	[0.2061]	[0.0603]
Joint sign test $\chi^2(3)$	12.522	8.9823	1.0602	5.2230	4.2495	5.4907
	[0.0658]	[0.0495]	[0.7867]	[0.1562]	[0.2357]	[0.1392]
The product of standardized residuals[2]	$z_t^{spot} * z_t^{ndf}$	$z_t^{ndf} * z_t^{df}$	$z_t^{df} * z_t^{spot}$	$z_t^{spot} * z_t^{ndf}$	$z_t^{ndf} * z_t^{df}$	$z_t^{df} * z_t^{spot}$
LB-Q(10)[c]	11.123	11.475	10.067	7.7018	2.6579	3.0113
	[0.348]	[0.322]	[0.435]	[0.658]	[0.988]	[0.981]
Covariance specification test (Koutmos and Knif, 2002)						
Sign test $\chi^2(3)$	1.2448	1.2077	3.1033	3.3382	3.4611	3.4147
	[0.5366]	[0.5467]	[0.2119]	[0.1884]	[0.1772]	[0.1813]
Negative size bias $\chi^2(2)$	2.8963	1.1653	0.5370	0.3473	0.6157	0.2206
	[0.2350]	[0.5584]	[0.7645]	[0.8406]	[0.7351]	[0.8955]
Positive size bias $\chi^2(2)$	5.5811	4.7434	7.6050	2.5865	4.6483	5.1746
	[0.0614]	[0.0933]	[0.0623]	[0.2744]	[0.0979]	[0.0752]
Joint sign test $\chi^2(6)$	11.3355	8.2936	8.7407	6.3568	7.8686	7.2153
	[0.0785]	[0.2174]	[0.1887]	[0.3844]	[0.2479]	[0.3014]
Shape parameters[3]	SPOT	NDF	DF	SPOT	NDF	DF
ln ξ_1	0.0727***			−0.0541***		
	(0.0278)			(0.0197)		
ln ξ_2		0.0495**			−0.0401**	
		(0.0229)			(0.0180)	
ln ξ_3			0.0720***			0.0045
			(0.0277)			(0.0175)
nu		3.0255***			3.2824***	
		(0.0724)			(0.0708)	
The conditions for existence of the 2nd and 4th moments of GARCH estimations[4]						
2nd	0.9197	0.8998	0.8598	0.9144	0.9263	0.8891
4th	0.8319	0.8165	0.7518	0.8622	0.8722	0.7820

Notes: z_t^i denotes the standardized residual, which should have zero mean, unit variance, and be i.i.d if the model is correctly specified.
LB-Q(10)[a] and LB-Q^2(10)[b] express the heteroscedasticity-consistent Ljung–Box Q(10) statistics to test the joint significance of the serial correlations up to 10 lags for standardized residuals and their associated squared standardized residuals, respectively.
LB-Q(10)[c] indicates the heteroscedasticity-consistent Ljung–Box Q(10) statistics to test the joint significance of the cross product of standardized residuals ($z_t^i * z_t^j$) up to 10 lags for market i and market j.
The shape parameters ξ_1, ξ_2 and ξ_3 indicate the skewness parameters for SPOT, NDF, and DF market, respectively; *nu* denotes the degree of freedom. The entries in parentheses are the standard deviations of coefficient estimates.
2nd and 4th moments represents the existence conditions of these moments by plugging in estimated parameters of their estimated model into the expressions (A.1) and (A.2) individually for SPOT, NDF, and DF in Korea and Taiwan.
Values in brackets represent the P-values of estimates and the symbols ** and *** indicate significance at the 5% and 1% level, respectively.

skewed shape parameters (ξ_1,ξ_2,ξ_3) for spot, NDF, and DF series are significant at the 1% level, affirming the importance of accounting for asymmetric properties. As such, we adopt the GMGARCH-MSKST model as a baseline for the empirical discussion that follows.

Table 4
The estimated parameters of GMGARCH-MSKST model.

Panel A: Estimates of return transmissions					
SPOT vs. NDF		SPOT vs. DF		NDF vs. DF	
Korea					
ψ^{ndf_spot}	0.4463***	ψ^{df_spot}	0.0318**	ψ^{ndf_df}	0.3771***
	(0.0337)		(0.0151)		(0.0349)
ψ^{spot_ndf}	0.2422***	ψ^{spot_df}	0.1008**	ψ^{df_ndf}	−0.0179
	(0.0414)		(0.0533)		(0.0180)
Taiwan					
ψ^{ndf_spot}	0.2437	ψ^{df_spot}	0.1187	ψ^{ndf_df}	0.2557***
	(0.1727)		(0.1042)		(0.0020)
ψ^{spot_ndf}	0.1779***	ψ^{spot_df}	0.3696***	ψ^{df_ndf}	0.1884***
	(0.0048)		(0.1626)		(0.0082)
Panel B: Estimates of volatility transmissions					
SPOT vs. NDF		SPOT vs. DF		NDF vs. DF	
Korea					
$\mu_1^{ndf\_spot}$	0.0030	$\mu_1^{df\_spot}$	0.0769**	$\mu_1^{ndf\_df}$	0.0261
	(0.0005)		(0.0358)		(0.0493)
$\mu_2^{ndf\_spot}$	0.0251***	$\mu_2^{df\_spot}$	−0.0973	$\mu_2^{ndf\_df}$	0.0217***
	(0.0001)		(0.1255)		(0.0080)
$\mu_1^{spot\_ndf}$	0.0171***	$\mu_1^{spot\_df}$	0.0107**	$\mu_1^{df\_ndf}$	0.0455***
	(0.0002)		(0.0051)		(0.0050)
$\mu_2^{spot\_ndf}$	0.0623***	$\mu_2^{spot\_df}$	0.0546**	$\mu_2^{df\_ndf}$	−0.0886
	(0.0003)		(0.0233)		(0.0946)
Taiwan					
$\mu_1^{ndf\_spot}$	0.0023	$\mu_1^{df\_spot}$	0.0251	$\mu_1^{ndf\_df}$	0.0170***
	(0.0167)		(0.0290)		(0.0002)
$\mu_2^{ndf\_spot}$	0.0038	$\mu_2^{df\_spot}$	−0.0415	$\mu_2^{ndf\_df}$	−0.0100
	(0.0254)		(0.0500)		(0.0276)
$\mu_1^{spot\_ndf}$	−0.0586	$\mu_1^{spot\_df}$	−0.0064	$\mu_1^{df\_ndf}$	0.0688***
	(0.0130)		(0.0040)		(0.0005)
$\mu_2^{spot\_ndf}$	0.0405***	$\mu_2^{spot\_df}$	0.0157***	$\mu_2^{df\_ndf}$	−0.0684
	(0.0060)		(0.0070)		(0.0753)

	Korea			Taiwan		
	SPOT	NDF	DF	SPOT	NDF	DF
Panel C: Estimates of error correction terms in mean and variance equations						
τ_R^i	−2.2432***	1.4161**	−3.3220***	−0.3592***	0.3647	−0.5940***
	(0.5255)	(0.6189)	(0.5536)	(0.0123)	(0.2540)	(0.0102)
τ_h^i	0.0323	0.0379**	0.0329	0.0243	0.0579**	0.0399***
	(0.0221)	(0.0166)	(0.0250)	(0.0258)	(0.0289)	(0.0113)
Panel D: Estimates of macroeconomic variables						
γ^i	0.0379***	0.0106*	0.0379***	0.0316***	0.0367**	0.0320***
	(0.0114)	(0.0060)	(0.0106)	(0.0120)	(0.0182)	(0.0065)
η^i	−0.0052***	−0.0098*	−0.0037***	−0.0034*	−0.0036**	−0.0058***
	(0.0007)	(0.0051)	(0.0005)	(0.0019)	(0.0017)	(0.0007)
ϖ^i	0.0232***	0.0606***	0.0551***	0.0481***	0.0663***	0.0478***
	(0.0079)	(0.0120)	(0.0104)	(0.0127)	(0.0237)	(0.0166)
ι_R^i	0.2439***	0.2386***	0.1033**	0.1035***	0.1284***	0.0576***
	(0.0871)	(0.0897)	(0.0567)	(0.0076)	(0.0340)	(0.0219)
Panel E: Estimates of ARCH models						
g^i	0.7134***	0.5112***	0.7918***	0.1657***	0.0892**	8.0236***
	(0.2191)	(0.1934)	(0.2449)	(0.0008)	(0.0392)	(0.0269)
β^i	0.8497***	0.8579***	0.8277***	0.8746***	0.9022***	0.8362***
	(0.0019)	(0.0196)	(0.0115)	(0.0040)	(0.0027)	(0.0018)

Table 4 (*continued*)

	Korea			Taiwan		
	SPOT	NDF	DF	SPOT	NDF	DF
Panel E: Estimates of ARCH models						
λ_1^i	0.0450***	0.0353***	0.0237***	0.0132***	0.0136***	0.0489***
	(0.0078)	(0.0085)	(0.0067)	(0.0041)	(0.0026)	(0.0036)
λ_2^i	0.0539	0.0138*	0.0180	0.0504	0.0202	0.0080
	(0.0458)	(0.0076)	(0.0112)	(0.0309)	(0.0384)	(0.0061)
ι_h^i	0.0655***	0.0226**	0.0949***	0.0978***	0.0924***	0.0208***
	(0.0085)	(0.0113)	(0.0103)	(0.0395)	(0.0337)	(0.0082)

Panel F: Estimates of conditional covariance equations

	Korea	Taiwan
α_1^{ij}	0.1256**	0.0218***
	(0.0025)	(0.0050)
α_2^{ij}	0.0906***	0.0240***
	(0.0024)	(0.0050)
α_3^{ij}	−0.0138***	−0.0058***
	(0.0029)	(0.0016)
α_4^{ij}	0.3906**	0.4761**
	(0.1959)	(0.2315)
α_5^{ij}	0.1776***	0.1465***
	(0.0215)	(0.0162)

Notes: The entries in parentheses are the standard deviations of coefficient estimates and the symbols *, ** and *** indicate significance at the 10%, 5% and 1% level, respectively.

Following Ling and McAleer (2002, 2003), we evaluate whether necessary and sufficient conditions for the existence of the unconditional second and fourth moments are satisfied. We derive the existence conditions of moments for our proposed GMGARCH-MSKST model.[21] Specifically, expressions (A.1) and (A.2) in our appendix are conditions for the existence of the second and fourth unconditional moment, respectively. Calculated estimates provided in the last two columns of Table 3 verify the existence conditions of these moments for spot, DF and NDF data for both Korean and Taiwan markets.

To facilitate a discussion of important contributions to the literature, Table 4 is presented as a concatenation of six panels (A to F) for the GMGARCH-MSKST estimation results. Each panel of estimates is discussed below and corresponds to the major research contributions of this work—as outlined in previous sections of this paper.[22]

5.1. Panel A: price transmission mechanism

Significant price feedback effects appear between non-deliverable forward and spot markets in Korea; whereas there is only unidirectional causality from spot to NDF markets in Taiwan. While there may be many explanations for this result, one possibility is that regulatory regimes that place restrictions on derivatives trading, such as those that restrict trades in the Taiwan NDF market, lead to less robust information flows between derivatives markets and spot markets. Results for the feedback and unidirectional causal effect from spot to DF markets are consistent with the analysis above between spot and NDF markets.

5.2. Panel B: cross-market volatility

Results for Korea indicate that a shock from an offshore NDF market triggers significant volatility in spot exchange rate markets. This insight is potentially valuable to policymakers who are tasked with managing volatility in onshore currency markets and further suggests that monetary authorities should carefully consider how cross-market volatility may impact policy directives that seek to reduce cross-market spillover

21 Necessary and sufficient conditions are presented in the Appendix A.
22 Tables that summarize all estimated parameters of the GMGARCH-MSKST model are available from the authors upon request.

effects to onshore markets. Moreover, estimates suggest that the volatility of NDF and DF exchange rate market returns for the Korean won are relatively more sensitive to news originating from spot appreciation shocks than to news from spot depreciation shocks. In addition, we find that volatility transmission from spot to NDF and DF markets in Korea are more persistent than that of other cross-market volatility transmission estimates, implying greater information spillovers generated from spot markets in Korea.

Regarding the cross-market volatility transmissions in Taiwan, the parameters from NDF to spot markets are insignificant, suggesting that the shocks emanating from NDF markets do not have a significant impact on the volatility of spot market returns. On the other hand, volatility transmission from the spot market does induce increased volatility in the NDF and DF markets. Moreover, spot appreciation shocks generally generate greater volatility in NDF and DF markets than depreciation shocks do in NDF and DF markets. This phenomenon corresponds to evidence from Korean markets. Our analysis generally reveals not only the causality from spot to NDF and DF in terms of price transmissions but also a unidirectional relationship in volatility transmissions from spot to NDF and DF markets. This result highlights the importance of spot markets in Taiwan and the spot market's substantial role of transmitting information to other markets.

5.3. Panel C: error correction process in first and second moments

Most estimated coefficients of error correction terms in spot, NDF, and DF return equations are significant at the 5% level. The significance of error correction terms suggests that deviation from long-run equilibrium is corrected gradually through a series of partial adjustments interacting within the three markets.

Our results support the validity of a long-run prediction hypothesis, suggesting that the error correction process can help predict subsequent movements in market returns. Following the arguments presented in Zhong et al. (2004), it is worthwhile to note that the sign of error correction terms are negative in the spot and DF return equations in both Korean and Taiwan markets, implying that the arbitrage effect appears to dominate in the error correction process of these markets. It is also noteworthy that adjustments back to exchange rate equilibriums are primarily concentrated in spot and DF markets, suggesting that currency traders in Korea and Taiwan can exploit arbitrage opportunities that result from temporary mispricing in these two markets.[23] However, the sign of the error correction term in the Korean NDF market is positive, suggesting that the momentum effect dominates in the NDF market (Zhong et al., 2004). Statistical insignificance of the Taiwan NDF error correction term is consistent with the view that Taiwan's central bank has one of the most restrictive NDF regulation regimes in Asia.

There are two additional findings that are notable. We found the absolute magnitude of the estimated error correction term for the DF market to be consistently larger than that of spot and NDF markets.[24] This suggests that DF markets tend to be the primary conduits for speeding adjustments back to long-run equilibrium. Second, the statistical significance of the error correction term in the NDF conditional variance equation suggests that NDF markets tend to respond with higher volatility as the deviation from long-run equilibrium increases. This result indicates that short-run disequilibrium in the spot market influences not only exchange rate market returns but also volatility in the NDF market.

[23] While arbitrage opportunities exist in the market—whether or not the opportunities can be exploited is another question. As the covered interest parity (CIP) formula suggests, a currency trader can in theory realize arbitrage profit if the forward rate deviates from CIP. Despite increasing efficiency in foreign exchange rate markets in the past decade, a number of studies document persistent violations of CIP during periods of uncertainty and turmoil before they are arbitraged away. Specifically, Baba and Packer (2009) provide a detailed explanation for the cause of such deviations and find that funding liquidity and counterparty risks are important determinants for dislocation in the foreign exchange swap markets during the financial crisis. In addition, Coffey et al. (2009) contend that CIP violations are partly due to liquidity constraints and partly due to heightened counterparty credit risk. In addition, deviations from covered interest parity might also arise from capital controls that restrict onshore currency markets from adjusting to external market dynamics.

[24] We examined various hypotheses using the likelihood ratio test to determine whether the error correction process influenced spot, NDF, and DF markets in an equivalent manner. Restrictions were placed on the relevant parameters and the following two sets of null hypotheses were tested: (1) H_0: No significant difference exists between the estimated error correction terms for DF and spot markets: i.e. $\tau_R^{df} = \tau_R^{spot}$; (2) H_0: No significant difference exists between the estimated error correction terms for DF and NDF markets: i.e. $\tau_R^{df} = \tau_R^{ndf}$. Likelihood ratio test statistics for the error correction term of the DF market relative to that of the spot and NDF markets were as follows: 4.98 and 43.59 (Korea); and 5.31 and 61.06 (Taiwan), respectively. These results were considered statistically significant at the 5% level, and suggested that the influence of the error correction term on DF markets is significantly greater than its influence on spot and NDF markets.

5.4. Panel D: structural effects

The interest rate parity theory suggests the importance of interest rates on currency market returns. This is affirmed in our finding that estimated coefficients for interest rate variables (γ^i) for all markets are statistically significant. In addition, we find estimates on stock returns (η^i) to be significant and negative for all markets in both Korea and Taiwan. This highly significant result suggests that exchange rate markets exhibit stock-oriented properties, implying that investors' prospects relative to the local equity market exchanges are highly correlated with returns in currency markets.

Estimates of the effect of the Japanese yen (ϖ^i) on spot, NDF, and DF markets are significant at the 1% level. Considering the intense export competition and increasing integration among Pacific-Basin financial markets, this result suggests that Japanese yen movements play a potentially crucial role in informing currency exchange rates for Korea and Taiwan. Since the magnitude of the estimated coefficients for yen influence on NDF markets in Taiwan and Korea are consistently larger than that of yen to spot and DF markets,[25] participants in the NDF markets should take changes in the yen exchange rate as a more fundamental reference point than that of spot and DF markets. Our results indicate that the yen has greater influence for NDF markets than that of spot and DF markets, suggesting NDF movements are potentially exposed to speculative psychology generated by the yen. In contrast to spot markets that have been saddled with a number of regulations, the less restrictive NDF market is found to be relatively more responsive to regional shocks arising from movements in the Japanese yen. These results suggest that the NDF market is likely to play an important role in transmitting yen-related information to the spot and onshore markets. These results offer insight into the transmission dynamics that arise from yen to Asian NDF markets—as the yen becomes more volatile, market players might want to structure investments in Taiwan and Korean currencies to reflect expectations of future appreciation/depreciation.

Finally, regarding the effect of the crisis period on exchange rate markets, estimated coefficients (ι_R^i) are all significantly positive at the 5% level. This suggests evidence of an increase in depreciation shocks for Korea and Taiwan during the specific crisis period. One more finding stands out. The magnitude of Korea's estimated coefficients are all larger than those for Taiwan, suggesting that Korean currency markets have been more sensitive to this episode of turbulence in global financial markets.

5.5. Panel E: GARCH effects

The statistically significant ARCH parameter (λ_1^i) in the conditional variance equation indicates that the foreign exchange markets for both Korea and Taiwan exhibit the phenomenon of volatility clustering. The coefficient λ_2^i, designed to capture information on asymmetries in each market's volatility toward its own innovations, is not significant. This result is consistent with findings reported by Tsui and Ho (2004) that do not find support for asymmetric volatility when these currencies are measured against the dollar. The sum of ARCH and GARCH parameters represent the persistence of the conditional volatility and indicate high persistence in the time varying volatility process for all currency markets in our study. Finally, most of the structural break estimates, ι_h^i, are statistically significant, suggesting that the conditional variance experienced sharp increases during the crisis period. Identifying structural change in conditional volatility is useful for applications in finance that rely on estimates of conditional volatility, such as option pricing, portfolio optimization, value at risk and hedging.

5.6. Panel F: covariance dynamics

Most estimates indicate broad-based rejection of the assumption of constant correlation coefficients. This result is consistent with the findings of De Goeij and Marquering (2004), Ferreira and Lopez (2005),

[25] We examined various hypotheses using the likelihood ratio test to determine whether the Japanese yen influenced spot, NDF, and DF markets in an equivalent manner. We test two sets of null hypotheses as follows: (1) H_0: No significant difference exists between the estimated effect of Japanese yen on NDF and spot markets: i.e. $\varpi^{ndf} = \varpi^{spot}$; and (2) H_0: No significant difference exists between the estimated effect of Japanese yen on NDF and DF markets: i.e. $\varpi^{ndf} = \varpi^{df}$. The likelihood ratio statistics for Japanese yen on the NDF market relative to that of the effect on the spot and DF markets are as follows: 6.02 and 8.30 (Korea); and 5.67 and 4.42 (Taiwan), respectively. These results were considered statistically significant at the 5% level, and suggested that the influence of the Japanese yen on NDF markets is significantly greater than its influence on spot and DF markets.

and Koutmos and Knif (2002). Investigation of the asymmetry properties observed in covariance structures suggests significant leverage effects involving conditional covariance structures. Estimates of most of the parameters for the product of two negative shocks (α_1^{ij}) in covariance equations are significant and positive, suggesting that contemporaneous movement is higher when related exchange rate markets experience joint negative (appreciation) shocks of the same sign in the previous period. Similarly, most estimates are significant and positive for the product of two previous positive shocks (α_2^{ij}) in the covariance equation. These results suggest that the conditional covariance tends to increase as exchange rate markets experience joint positive (depreciation) shocks of the same direction in the previous period. Furthermore, the cross asymmetry effect (α_3^{ij}) is significant and negative. The negative signs on estimated exchange rate covariance parameters suggest that, in general, conditional covariance tends to be lower when a negative (positive) shock in a market is combined with a positive (negative) shock in the related market. Interestingly, exchange rate covariance exhibits momentum spillover—where previous shocks of the same sign tend to increase the expected conditional covariance. In contrast, two shocks of the opposite sign in the previous period tend to decrease the contemporaneous conditional covariance. These results provide evidence that not only supports the coexistence of positive and negative leverage effects in exchange rate co-movements, but also successfully captures the cross asymmetries in the exchange rate covariance structures. In addition, estimates of the coefficient for the lagged conditional covariance (α_4^{ij}) are all significant. This result suggests that not only variance but also covariance exhibits clustering behavior; indicating that a higher (lower) degree of conditional covariance is usually followed by a higher (lower) subsequent conditional covariance in currency markets.

Moreover, the coefficient that captures evidence of structural change in conditional correlation (α_5^{ij}) is significant and positive. The crisis period generates significant increases in the bivariate conditional correlations within spot, DF and NDF markets. More specifically, the magnitude of Korea's estimated structural change coefficient is higher than that for Taiwan, indicating that Korea's conditional correlations within exchange rate markets are higher relative to those of Taiwan during the crisis period. This insight into the nature of excess correlation during the crisis period is potentially useful for investors who seek to structure a diversifiable portfolio of currency investments that hedge against structural-change risk in markets.

Finally, the average of dynamic conditional correlations generated from the GMGARCH-MSKST model estimation results are reported in Table 5. The average correlation coefficient between spot and DF markets is the highest among all three coefficients while that of NDF and DF markets is the lowest. Data plots in Fig. 2 reflect the fitted conditional correlations between the spot and NDF, the spot and DF, and the NDF and DF markets respectively. The plots illustrate how the estimated correlation coefficients between exchange rates fluctuate over time and are generally consistent with research that has documented the properties of currency markets and other asset markets, such as those of stocks and bonds (Cappiello et al., 2006; Li, 2011). Further examination of each pair of the plots shows that exchange markets closely correspond with each other and possess strong positive correlation.

6. Conclusion

As a methodological innovation, we proposed using a GMGARCH-MSKST model that is flexible in its capacity to accommodate the data generating characteristics found in currency market returns. Using lessons drawn from Korean and Taiwan domestic and offshore currency markets it appears that the more liberal currency market policies of Korea are associated with evidence of more market integration. While

Table 5
The average of dynamic conditional correlations calculated by the GMGARCH-MSKST model.

	Korea			Taiwan		
	SPOT vs. NDF	SPOT vs. DF	NDF vs. DF	SPOT vs. NDF	SPOT vs. DF	NDF vs. DF
Mean	0.9894	0.9964	0.9873	0.9918	0.9933	0.9915
Std. dev.	0.0208	0.0117	0.0238	0.0107	0.0125	0.0141
MAX	0.9998	0.9999	0.9997	0.9996	0.9997	0.9997
MIN	0.7280	0.6764	0.6674	0.8305	0.7804	0.7976

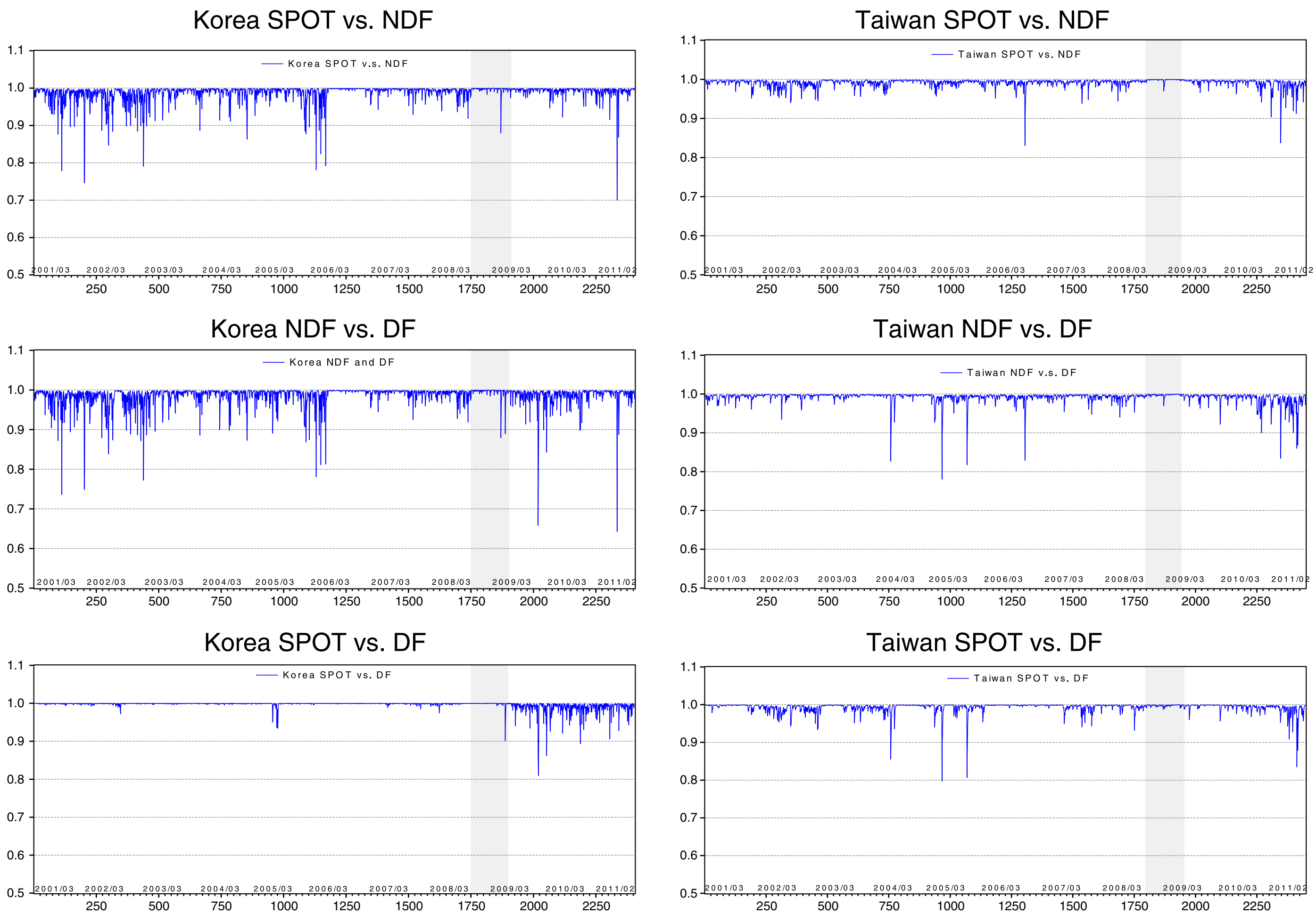

Fig. 2. The conditional correlations of exchanges rates currency pairs: the plots show respective conditional correlations of exchange rate currency pairs for SPOT and NDF (top), NDF and DF (middle), and SPOT and DF (bottom) by employing the GMGARCH-MSKST model analyzed in the study.

not conclusive, our research contributes to accumulating evidence about information transmission between offshore and domestic exchange rate markets under different regulatory and institutional regimes. It also creates a framework for future cross-country analysis of offshore and domestic market integration under different regulatory frameworks. Empirical contributions are summarized as follows.

- First, there is evidence that spot, NDF, and DF markets are related to each other with price feedback effects in Korea. Our observations suggest that offshore Korean NDF market factors have become a potentially important information tool to gauge the spot exchange rate regime going forward. In Taiwan, we find that the spot market exerts influence over NDF and DF markets. In contrast to Korea, the offshore Taiwan NDF market has less influence on the spot market and follows blindly or lags behind returns in the spot market.
- Second, our estimations show greater intensity of cross-market volatility spillovers within markets in Korea, reflecting considerable information transmission through spot, NDF, and DF markets. However, this may also expose the Korean spot market to higher volatility as appreciation/depreciation shocks propagate from offshore markets. In Taiwan we find that shocks in spot markets induce increased volatility in NDF and DF markets. Our results highlight the importance of the spot market in transmitting information to other markets for Taiwan's currency. The irrelevance of information flows from NDF to spot markets might be beneficial to Taiwanese policymakers who conduct a monetary policy that targets domestic currency market stability. Nevertheless, this might eliminate the usefulness of NDF markets for hedging purposes in periods of market stress.
- Third, our results indicate that the error correction process is a useful price discovery vehicle to predict subsequent price movements of spot, NDF, and DF markets. Specifically, DF markets play a significant role in triggering a quicker adjustment back to equilibrium. These results underscore the importance of external factors in determining the dynamics of Asian exchange rate markets, within local as well as regional influences. In particular, our results suggest evidence of yen/$ movements impacting Korean and Taiwan currency markets. Traders in the NDF markets should take changes in the yen exchange rate as a fundamental reference point for spot and DF market returns.
- Fourth, this research proposes a flexible asymmetric conditional covariance specification, which contributes to the literature not only in characterizing the leverage effects but also cross asymmetry effects in exchange rate covariance dynamics. Our results show that the conditional covariance has a tendency to increase as markets have experienced shocks of the same direction in the previous period; whereas the conditional covariance decreases as markets experience shocks of opposite signs in the previous period. Importantly, we find not only that variance is clustering but also that covariance is clustering. This suggests that a higher (lower) degree of conditional covariance is usually followed by a higher (lower) subsequent conditional covariance in the Korean and Taiwan currency markets.
- Fifth, while each exchange rate pair shows positive contemporaneous correlation, the average dynamic conditional correlation of spot and DF market pairings are the highest among all three in-country pairings. The NDF and DF market pairings are the lowest.
- Finally, we not only find supportive evidence of significant depreciation shocks and sharp increases in volatility during the U.S. subprime financial crisis, but also observe evidence that correlation coefficients across markets are likely to increase during high external volatility in crises periods.

Appendix A

Consider the following general class of GARCH(1,1) model: [26]

$$\varepsilon_t = z_t \sigma_t;$$
$$\sigma_t^k = g(z_{t-1}) + c(z_{t-1})\sigma_{t-1}^k$$

where $\{z_t\}$ is a sequence of independent identically distributed random variables with zero mean. k equals 1 or 2. $g_t = g(z_t)$ and $c_t = c(z_t)$ are well-defined functions of z_t. The necessary and sufficient condition for the *mk th* unconditional moment to exist is $E[c(z_t)^m] < 1$.

Referencing the MSKST–GJR-GARCH(1,1) model of Bauwens and Laurent (2005), let $k = 2$, $g(z_{t-1}) \equiv \alpha_0$ and $c(z_{t-1}) = \beta + (\alpha + wI(z_{t-1}))z_{t-1}^2$ where $I(z_{t-1}) = 1$ if $z_{t-1} < 0$ and $I(z_{t-1}) = 0$, otherwise. The density

[26] For derivation details please refer to Ling and McAleer (2002, 2003).

of $z_t \sim SKST(0,1,\xi, nu)$ (i.e. distributed as a standardized skewed-Student's t with parameters nu (the number of degrees of freedom) and $\xi > 0$ (a parameter related to skewness)) is given by

$$f(z_t|\xi, nu) = \begin{cases} \frac{2}{\xi + 1/\xi} g[\xi z_t|nu] & if z_t < 0 \\ \frac{2}{\xi + 1/\xi} g\left[{}^{z_t}/_{\xi}|nu\right] & if z_t \geq 0 \end{cases}$$

where $g(.\,|nu)$ is a symmetric (zero mean and unit variance) Student's t density with nu degrees of freedom. Thus, $f(z_t|\xi, nu)$ is a unimodal density with the same mode as $g(z_t|nu)$ and skewness parameter $\xi > 0$ such that the ratio of probability masses above and below the mode is:

$$\frac{\Pr(z_t \geq 0|\xi)}{\Pr(z_t < 0|\xi)} = \xi^2 \rightarrow \Pr(z_t < 0|\xi) = \frac{1}{1+\xi^2}$$
$$E[I(z_{t-1})] = E[I(z_{t-1})^2] = 1/\left(1+\xi^2\right).$$

By extension the second moment condition is given by:

$$\begin{aligned} E[c(z_{t-1})] &= \beta + \alpha E[z_{t-1}^2] + wE[I(z_{t-1})]E[z_{t-1}^2] \\ &= \beta + \alpha + w\frac{1}{1+\xi^2} < 1. \end{aligned} \quad (A.1)$$

And the fourth moment condition is given by:

$$\begin{aligned} E[c(z_{t-1})^2] &= E[\left(\beta + [\alpha + wI(z_{t-1})]z_{t-1}^2\right)^2] \\ &= E[\beta^2 + \left(\alpha^2 + 2\alpha wI(z_{t-1}) + w^2 I(z_{t-1})^2\right)z_{t-1}^4 + 2\alpha\beta z_{t-1}^2 + 2\beta wI(z_{t-1})z_{t-1}^2] \\ &= \beta^2 + \left\{\alpha^2 + 2\alpha wE[I(z_{t-1})] + w^2 E[I(z_{t-1})^2]\right\}E[z_{t-1}^4] + 2\alpha\beta + 2\beta wE[I(z_{t-1})]. \end{aligned}$$

According to Fernandez and Steel (1998), if the r^{th} ($r \in \mathfrak{R}$) order moment of $g(.)$ exists, the associated skewed distribution also has a finite r^{th} moment. In particular,

$$E(z_t^r|\xi) = M_r \frac{\xi^{r+1} + \frac{(-1)^r}{\xi^{r+1}}}{\xi + \frac{1}{\xi}}$$

where $M_r = \int 2u^r g(u)du$, and M_r is the r^{th} order moment of $g(.)$ truncated to positive real values. By extension, the following equation reflects the existence condition for the fourth moment of the proposed GMGARCH-MSKST model.

$$\begin{aligned} E[c(z_{t-1})^2] &= \beta^2 + \left\{\alpha^2 + 2\alpha wE[I(z_{t-1})] + w^2 E[I(z_{t-1})^2]\right\}E[z_{t-1}^4] + 2\alpha\beta + 2\beta wE[I(z_{t-1})] \\ &= \beta^2 + 2\alpha\beta + \alpha^2 s + \frac{1}{1+\xi^2}\left(2\beta w + 2\alpha ws + w^2 s\right) < 1 \end{aligned} \quad (A.2)$$

where $s = E\left[z_{t-1}^4\right] = \frac{\xi^5 + \frac{1}{\xi^5}}{\xi + \frac{1}{\xi}} 3(nu-2)/(nu-4)$.

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