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Price cointegration between sovereign CDS and currency option markets in the financial crises of 2007–2013

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ABSTRACT

The sovereign credit default swap (CDS) spreads and exchange rates of the developed economies including the US, Japan, Switzerland and the eurozone with the first three countries' currencies conventionally considered as safe-haven varied in a wide range during the financial crises since late 2007. This raises the question of any interconnectivity between the anticipated sovereign credit risks of these economies and the market expectations of their exchange rates. Using a bivariate vector error-correction model with random coefficients, this paper finds evidence of cointegration and time varying conditional correlation between the prices in the sovereign CDS and currency option markets. This suggests that the relative sovereign credit risk of these developed economies impacts the market expectations of their exchange rates in the long run, but in the short run the impact changes drastically in times of crisis, resulting in drastic and persistent price deviations from their long-run equilibrium amid central banks' monetary measures and market turbulence.

1. Introduction

A sovereign credit default swap (CDS) is an over-the-counter credit protection contract in which a protection seller pays compensation to a protection buyer to make a payment in the case of a pre-defined credit event. For credit protection buyers who pay a fixed premium called the CDS spread, the CDS market offers the opportunity to reduce credit risk. For protection sellers, it offers the opportunity to take credit exposure to an entity and earn income without having to fund the position. Sovereign CDS spreads have been used as a direct measure of the creditworthiness of the underlying sovereign, for example in Pan and Singleton (2008).[1] A change in the credit risk of a sovereign borrower reflected in its sovereign CDS spread can thus be considered as an indicator of the country's economic–political stability, which is linked to country-specific macro-economic variables, such as output growth, foreign exchange reserves, budget deficit, real effective exchange rate deviation, and foreign direct investment. A substantial increase in sovereign risk due to economic–political instability would lead investors to sell securities denominated in the country's currency and to repatriate funds, hence putting downward pressure on the currency and heightening its volatility. The relationship between sovereign risk and exchange rate stability has long been the subject of interest in international finance including those of Eichengreen, Rose, and Wyplosz (1996), Frankel and Rose (1996), Kaminsky, Lizondo, and Reinhart (1998), and Kumar, Moorthy, and Perraudin (2003), who use macro-economic indicators to estimate the probability of currency crashes.

In view of the linkage between a country's credit risk and the strength of its currency, interactions between emerging market sovereign CDS spreads and corresponding exchange rate expectations anticipated in currency option markets are studied by Carr

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[1] The sovereign CDS market expanded rapidly in 2009 and 2010. The gross notion of protection was around US$2 trillion as of 2010. See the IMF's Global Financial Stability Report (Meeting New Challenges to Stability and Building a Safer System, April 2010).

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and Wu (2007). Currency option markets have the desirable property of being forward-looking in nature and are thus useful sources of information for gauging market sentiments about future exchange rates. Options, whose payoff depends on a limited range of the expected exchange rate, offer broad information about market expectations. Potential extreme movements of the exchange rate can be inferred from out-of-the-money option prices. Carr and Wu (2007) investigate the relationship between currency option-implied volatilities and sovereign creditworthiness for Mexico and Brazil from 2002 to 2005. They find that the level and skew of the option-implied volatility display significant co-movement with the sovereign CDS spreads of the two countries. This suggests that the currency option market has consistently set prices considering the probability of a currency crash triggered by a corresponding sovereign default of the two countries.

The subprime crisis in the US during 2007–2008 was closely followed by the European sovereign debt crisis which began in 2009 and may still be unfolding. During this period, there have been several bouts of financial turbulence, causing sharp changes in risk assessment globally, with large swings in the developed economies' foreign exchange market. The US dollar (USD) had once depreciated about 30% against the Japanese yen during the subprime crisis. As concerns spread during the European sovereign debt crisis, the euro fell sharply against the USD by about 19% as of April 2010 since November 2009. Hui and Chung (2011) show that the creditworthiness of euro-area countries distinct from other macro-financial factors can affect market expectations on the stability of the euro. They find evidence of information flow from the sovereign CDS market to the USD-euro currency option market during September 2009–April 2010. The impact has been considerable and even disturbing enough to cause some policymakers (e.g., the Swiss National Bank (SNB)) to resort to drastic policy action (e.g., changing the country's exchange rate regime in the case of Switzerland).

Given the fact that Japan's and the US's government debt to GDP ratios were the highest (233%) and the third (100%) among the G20 economies at the end of 2011, their level and trend of government debt have raised concerns over their sovereign risk. The IMF also pointed out that Japan and the US may not be immune to the European-debt-crisis-style risks.[2] In addition to the concerns of their sovereign risks, the depreciation of the USD during the crisis period was quite inconsistent with the conventional wisdom that the USD together with the Japanese yen and Swiss franc are safe-haven currencies during financial crises according to the findings in Ranaldo and Söderlind (2010) and Kohler (2010).[3] In view of these observations, the exchange rate movements, in particular in favor of safe-haven currencies, show that the prices in the developed economies' sovereign CDS market and corresponding currency option market may have appreciable impact on each other.

This paper investigates how the sovereign risk affects market expectation on the exchange rates embedded in their currency options in the developed economies including the US, Japan, Switzerland and the eurozone. The market expectations of the exchange rates are reflected by the price of risk reversals quoted in the currency option market. The risk reversal is a directional option strategy that takes the view of the skewness of the exchange rate distribution by simultaneously buying an out-of-the-money put and selling an out-of-the-money call. It measures the implied volatility difference between an out-of-the-money put and call at the same (absolute) delta. The delta is a measure of a change in the option price with respect to a small change in the underlying exchange rate.[4] Fig. 1a and d shows the generally positive risk reversals of USD–yen and euro–yen options at the 25% delta (25-delta risk reversals) during the period from 2007 to 2013.[5] A positive risk reversal implies that the risk-neutral exchange rate distribution is negatively skewed. This reflects that the USD-put (euro-put) implied volatility is higher than the USD-call (euro-call) implied volatilities.[6] The asymmetry in the implied volatility occurs because market participants think that a depreciation of the USD and euro against the yen is more likely than an appreciation of the same size. One reason for such asymmetry in the expectation of the exchange rates is that only 5% of the Japanese government bonds are held by foreign investors, which is much smaller than 48% of the US Treasury securities held by foreign investors.[7] As foreign investors are more likely to sell a country's government bonds with a rising default risk compared with home investors, when both the US and Japan's sovereign credit risks are rising, the US Treasury securities would be under a larger selling pressure by foreign investors relative to the Japanese government bonds. This makes the yen safer than the USD.

Fig. 2 reports the net notional amounts outstanding (i.e., net protection bought) and average daily amounts of transactions of the CDS contracts of the US, Japan, Switzerland and the four highly indebted European counties including Ireland, Italy, Portugal and Spain as at 9 August 2013. The Japanese sovereign CDS contracts had a total net notional value of US$9.5 billion which is relatively small compared to those of Italy and Spain, but higher than those of the US, Portugal, and Ireland.[8] While the amounts of net notional outstanding of the US and Japanese sovereign CDS are limited, changes in their spreads are commonly used by policy makers to monitor for signals of concerns about their sovereign risks anticipated by market participants.[9]

In this paper, we use cointegration to model the joint behavior of the sovereign CDS spreads of the US, Japan, Switzerland and the eurozone and the risk reversals in the currency option market during the financial crisis period of 2007–2013. We investigate how the

[2] See IMF's "Global Financial Stability Report" and "World Economic Outlook", April 2012.

[3] There is no well-accepted definition of a safe haven asset. It could mean an asset with low risk or high liquidity and have a common characteristic that one would expect the relative price of such an asset to increase during crises.

[4] The delta of the option is roughly equal to the risk-neutral probability of the underlying ending in-the-money. For example, a 25% delta put option has approximately 25% probability of in-the-money at maturity. The Black–Scholes delta provides a normalized measure of option moneyness where the delta of a European option increases monotonically from 0–100, with the moneyness moving from out-of-the-money to in-the-money.

[5] The option data are downloaded from JPMorgan Chase.

[6] A dollar put (call) option against the yen here is a European option of selling (buying) dollars at the contractual option strike price in an exchange of yens at the option maturity.

[7] The figure for the Japanese government bonds is from the Bank of Japan as at March 2011 and the figure for the US Treasury securities is from the Bureau of Public Debt as at June 2011.

[8] Details of the structure of the sovereign CDS market are in Pan and Singleton (2008).

[9] For example, see Box 1.2 "How Concerned are Markets About US Sovereign Risks" in the IMF's September 2011 Global Financial Stability Report. However, the US sovereign CDS spread is not considered very reliable for extracting mathematical default probability.

sovereign credit risk and exchange rate risk interact with each other by testing the following three hypotheses: (i) in the long run the relative sovereign credit risk has considerable impact on the market expectations of exchange rates in the developed economies; (ii) in the short run the impact changes drastically and persistently in times of crisis; and (iii) the sovereign CDS spreads move

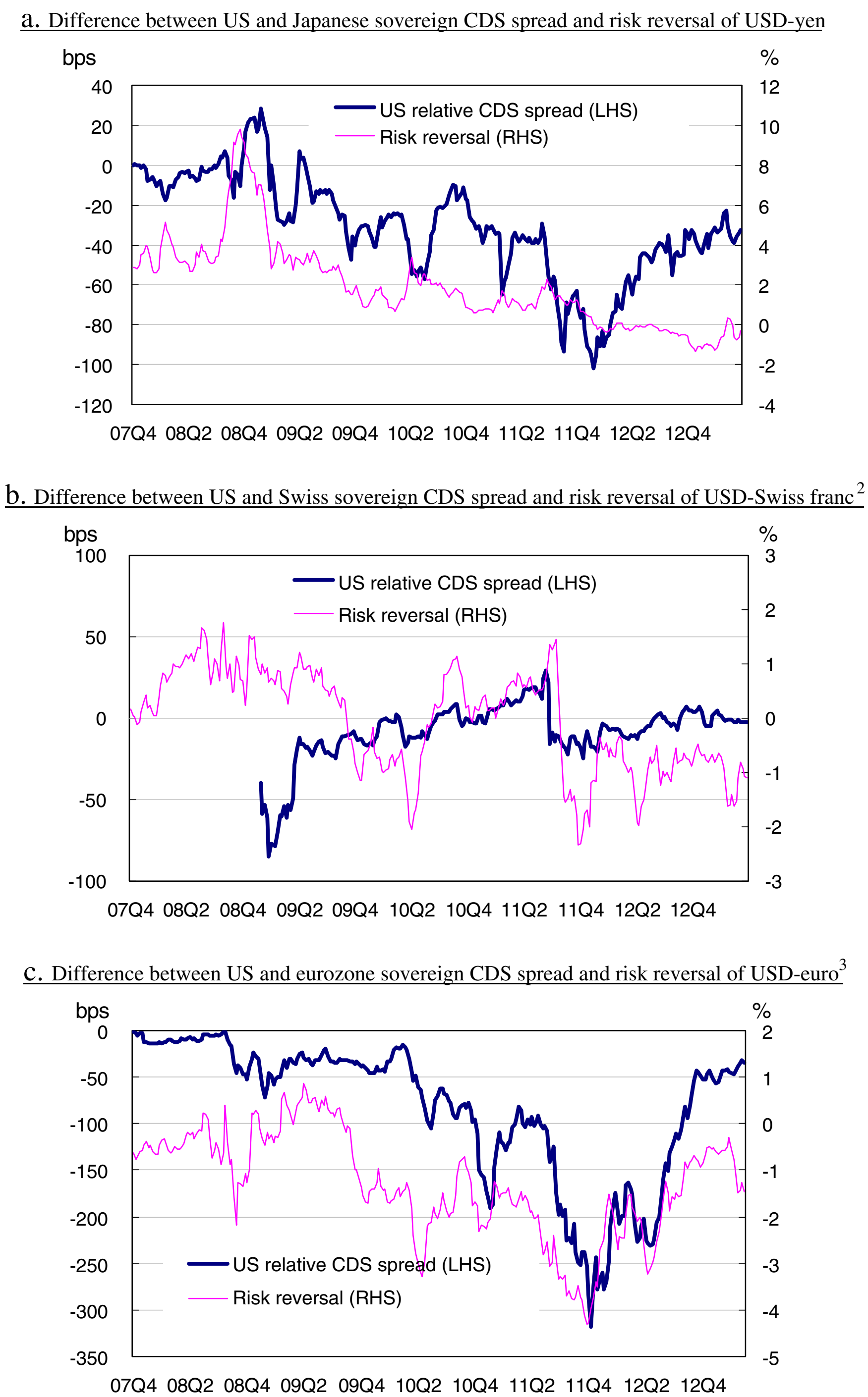
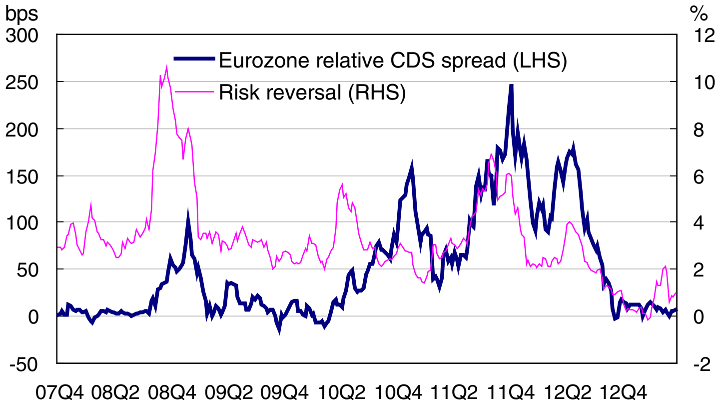


Fig. 1. Relative sovereign CDS spreads and risk reversals[1].
Notes:

1. The risk reversal is the implied volatility of an out-of-the-money USD (euro) put minus that of an out-of-the-money USD (euro) call at the 25% delta.
2. The Switzerland's sovereign CDS spreads is only available since January 2009 in the data source (JP Morgan Chase).
3. The USD-euro risk reversal and the US-eurozone sovereign CDS spreads can offer perspectives of both USD- and euro-based investors at the same time.
   Sources: JP Morgan Chase and Bloomberg.

d. Difference between eurozone and Japanese sovereign CDS spread and risk reversal of euro-yen



e. Difference between eurozone and Swiss sovereign CDS spread and risk reversal of euro-Swiss franc[2]

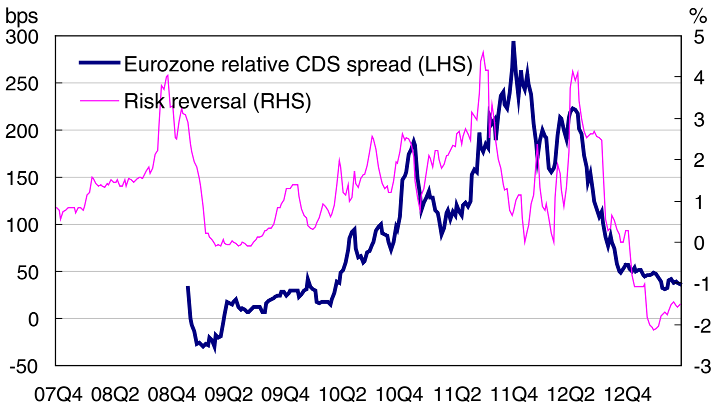


Fig. 1 (*continued*).

ahead of the currency option prices in the price discovery process. The test results could have implications for the predictability of the economies' exchange rate risk through the use of the information in the sovereign CDS market and the effects of the monetary policies adopted by their central banks on the currency and sovereign CDS markets.

This paper demonstrates that interconnectivity not only appears between the corporate CDS market and the corresponding stock (or stock option) market, but also exists between the sovereign CDS market and the currency option market in the developed economies. In the corporate sector, Acharya and Johnson (2007) find that the corporate CDS market leads the stock market to anticipate adverse credit information of the reference firm and this finding is linked to informed-trading in credit derivatives. This is reflected by incremental information flow from the corporate CDS market to the stock market. Cao, Yu, and Zhong (2010) document that implied volatility of deep out-of-the-money put options of stocks is closely related to corporate CDS spreads, because the options provide investors with similar protections against downside risk. They conclude that stock options play an important role in the price discovery process for firms' credit risk.

This paper is organized as follows. The following section discusses the cointegration and error-correction model. Section 3 presents the data. Section 4 examines the estimation results for testing the hypotheses. Section 5 contains the conclusion.

2. Methodology

The principal feature of cointegration is that a linear combination of non-stationary variables is stationary. This implies that cointegrated variables do move independently of each other but they are linked by the stationary linear combination. This stationary relationship is regarded as a long-run equilibrium among the cointegrated variables. Under this equilibrium, a short-term deviation of a cointegrated variable from the others is expected to be temporary, and this cointegrated variable will gradually revert to the long-run relationship.

This section presents the model to test any cointegration relation between the prices in the currency option and sovereign CDS markets. Such relation can be illustrated by considering a rise in the sovereign CDS spread of the US relative to the CDS spread of another country when the CDS spreads have an equilibrium relationship with foreign exchange (against USD) option prices in terms of risk reversals. This triggers a gap between the prices of the two markets. If this gap is large enough, the gap will ultimately be closed by

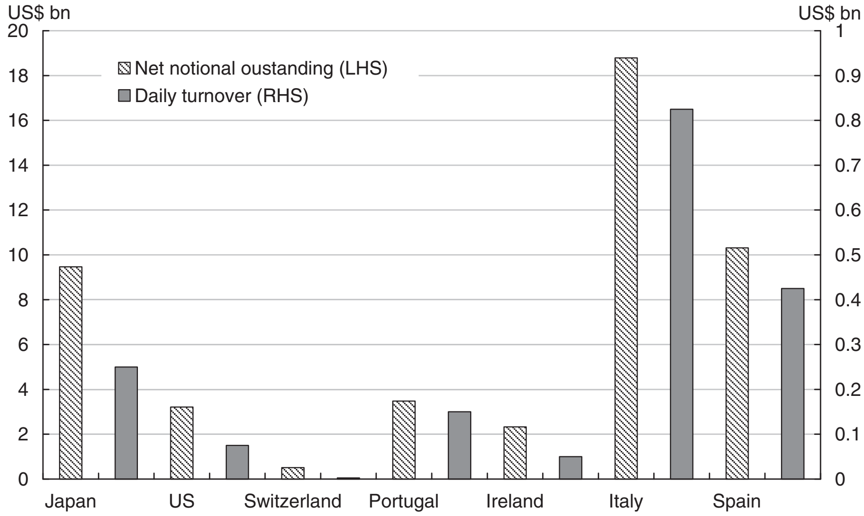


Fig. 2. Net notional amounts outstanding and average daily turnover of sovereign CDS contracts of the US, Japan, Switzerland and the four highly indebted European countries.
Notes:

1. The data are from Depository Trust and Clearing Corporation. See www.dtcc.com.
2. Net notional amounts outstanding are the aggregate net protection bought (or equivalently sold) across counterparts. Except for Switzerland, the net notional outstanding shown is as at 9 August 2013 and the average amounts of daily turnover are the average during the period from 20 December 2012 to 19 March 2013.

(1) a rise in an appreciation expectation of the foreign currency against USD (an increase in the risk reversal) and/or a fall in the US sovereign CDS spread; or (2) a further increase in the US sovereign CDS spread with a commensurately larger increase in the appreciation expectation; or (3) a fall in the US sovereign CDS spread with a smaller reduction in the appreciation expectation.

The above illustration is a dynamical error-correction model. In the model, the short-term dynamics of the variables in the system are influenced by deviations from equilibrium. Assuming that the US sovereign CDS spread and risk reversal are "integrated of order 1" denoted by $I(1)$ (i.e., non-stationary in levels, but stationary in changes) and cointegrated, a simple error-correction model for the corresponding variables can be expressed as:

$$\Delta y_t = \alpha_y(y_{t-1} - \beta x_{t-1}) + \varepsilon_{yt} \tag{1}$$

$$\Delta x_t = \alpha_x(y_{t-1} - \beta x_{t-1}) + \varepsilon_{xt} \tag{2}$$

where y_t and x_t are the risk reversal of the exchange rate of the USD against another country's currency, and their relative sovereign CDS spread respectively at time t, and both α_y and α_x are greater than zero. As specified, the two variables will change in response to stochastic shocks (represented by ε_{yt} and ε_{xt}) and to the previous period's deviation from the long-run equilibrium. If this deviation is positive (with $y_{t-1} - \beta x_{t-1} > 0$), the appreciation expectation of the foreign currency (i.e., risk reversal) rises and the relative US sovereign CDS spread falls. The long-run equilibrium is attained when $y_t = \beta x_t$. It is noted that α_y has to be negative and α_x has to be positive in order to ensure that when a deviation occurs, y_t will adjust downward and x_t will adjust upward subsequently to restore the long-run equilibrium.[10]

The parameters α_y and α_x are the speeds of adjustment. In absolute terms, the larger α_y is, the greater the response of y_t to the previous period's deviation from the long-run equilibrium. At the opposite extreme, a very small value of α_y in absolute terms implies that the risk reversal is unresponsive to the last period's equilibrium error. If both α_y and α_x are equal to zero, the long-run equilibrium relationship does not appear and the model is not error-correction or cointegration. Thus, for a meaningful cointegration and error-correction model, at least one of the speeds of adjustment terms in Eqs. (1) and (2) must be non-zero.

The discussion above is unaltered if we formulate a more general model by introducing the lagged changes of each variable into the equations:

$$\Delta y_t = a_{10} + \alpha_y(y_{t-1} - \beta x_{t-1}) + \sum_k b_{1k}\Delta y_{t-k} + \sum_k c_{1k}\Delta x_{t-k} + \varepsilon_{yt} \tag{3}$$

[10] The relationship between the error-correction model and the cointegrated variables can also be illustrated by the following argument. By assumptions, Δy_t is stationary (i.e. "integrated of order zero" denoted by $I(0)$), so that the left-hand side of Eq. (1) is stationary. For Eq. (1) being sensible, the right-hand side must be $I(0)$ as well. Given that ε_{yt} is stationary, it follows that the linear combination $y_{t-1} - \beta x_{t-1}$ must be stationary; hence, the two variables must be cointegrated. This identical argument can also be applied to Eq. (2).

$$\Delta x_t = a_{20} + \alpha_x(y_{t-1} - \beta x_{t-1}) + \sum\nolimits_k b_{2k}\Delta y_{t-k} + \sum\nolimits_k c_{2k}\Delta x_{t-k} + \varepsilon_{xt}. \tag{4}$$

The above specification is very similar to a vector autoregressive model (VAR). This two-variable error-correction model can be regarded as a bivariate VAR in the first differences augmented by the error-correction terms $\alpha_y\,(y_{t-1} - \beta x_{t-1})$ and $\alpha_x\,(y_{t-1} - \beta x_{t-1})$. In a vector form, the model can be re-written as:

$$\begin{pmatrix} \Delta y_t \\ \Delta x_t \end{pmatrix} = \begin{pmatrix} a_{10} \\ a_{20} \end{pmatrix} + \begin{pmatrix} \alpha_y \\ \alpha_x \end{pmatrix} \cdot \eta_{t-1} + \sum_{k=1}^{K} \begin{pmatrix} b_{1k} & c_{1k} \\ b_{2k} & c_{2k} \end{pmatrix} \begin{pmatrix} \Delta y_{t-k} \\ \Delta x_{t-k} \end{pmatrix} + \begin{pmatrix} \varepsilon_{yt} \\ \varepsilon_{xt} \end{pmatrix} \tag{5}$$

where $\eta_t = y_t - \beta x_t$. In this model, the cointegrating vector is said to be $(1, \beta)$.

We can basically apply the conventional model specified in Eq. (5) to test the hypotheses. However, the conventional model assumes that the long-run equilibrium (i.e. $y_t = \beta x_t$) does not change over time, which may be too strong for modeling financial time series. Hansen (1992) tested this assumption by checking some published cointegrating regressions. By calculating simple split tests on the variance of the regression error, strong evidence for non-constancy of the error variances is found in these cointegrating regressions. In view of this, we employ a generalized error-correction model proposed by Fong and Li (2004). Specifically, the models for country *i* at time *t* with reference to the US's and eurozone's conditions are represented respectively by the following specifications:

$$\begin{pmatrix} \Delta \mathrm{RR}(i,\mathrm{US})_t \\ \Delta \mathrm{CDS}(i,\mathrm{US})_t \end{pmatrix} = \begin{pmatrix} a_{\Delta \mathrm{RR}} \\ a_{\Delta \mathrm{CDS}} \end{pmatrix} + \begin{pmatrix} \alpha_1 \\ \alpha_2 \end{pmatrix} \cdot \eta^*_{\mathrm{US},t-1} + \sum_{k=1}^{K} \Phi_k \begin{pmatrix} \Delta \mathrm{RR}(i,\mathrm{US})_{t-k} \\ \Delta \mathrm{CDS}(i,\mathrm{US})_{t-k} \end{pmatrix} + \Theta X_t + \begin{pmatrix} \varepsilon_{\Delta \mathrm{RR},t} \\ \varepsilon_{\Delta \mathrm{CDS},t} \end{pmatrix} \tag{6}$$

and

$$\begin{pmatrix} \Delta \mathrm{RR}(i,\mathrm{EU})_t \\ \Delta \mathrm{CDS}(i,\mathrm{EU})_t \end{pmatrix} = \begin{pmatrix} a_{\Delta \mathrm{RR}} \\ a_{\Delta \mathrm{CDS}} \end{pmatrix} + \begin{pmatrix} \alpha_1 \\ \alpha_2 \end{pmatrix} \cdot \eta^*_{\mathrm{EU},t-1} + \sum_{k=1}^{K} \Phi_k \begin{pmatrix} \Delta \mathrm{RR}(i,\mathrm{EU})_{t-k} \\ \Delta \mathrm{CDS}(i,\mathrm{EU})_{t-k} \end{pmatrix} + \Theta X_t + \begin{pmatrix} \varepsilon_{\Delta \mathrm{RR},t} \\ \varepsilon_{\Delta \mathrm{CDS},t} \end{pmatrix} \tag{7}$$

with the long-run equilibrium errors between the two markets of $\eta^*_{\mathrm{US},t} = \mathrm{RR}(i,\mathrm{US})_t - \beta_{\mathrm{US},t}\mathrm{CDS}(i,\mathrm{US})_t$ and $\eta^*_{\mathrm{EU},t} = \mathrm{RR}(i,\mathrm{EU})_t - \beta_{\mathrm{EU},t}\mathrm{CDS}(i,\mathrm{EU})_t$ respectively, where RR(*i*, US) is the 25-delta risk reversal (i.e., the implied volatility of a put minus that of a call at the 25% delta) of the USD against the country *i*'s currency, and RR(*i*, EU) is the risk reversal for the euro; CDS(*i*, US) is the difference between the sovereign CDS spreads of the US and country *i*, and CDS(*i*, EU) is the difference between the sovereign CDS spreads of the eurozone and country *i*; α_1 and α_2 are speeds of cointegration adjustment for RR and CDS; β_t is a random coefficient in the long-run cointegration relationship following a normal distribution with mean β and variance σ^2_β; X_t is a vector of exogenous macroeconomic and financial variables; $\varepsilon_{\Delta \mathrm{RR}}$ and $\varepsilon_{\Delta \mathrm{CDS}}$ are two *r*-correlated error terms with variances $\sigma^2_{\Delta \mathrm{RR}}$ and $\sigma^2_{\Delta \mathrm{CDS}}$ respectively; a_Y is a constant term for the variable *Y*; and Δ is the first difference operator.

Eqs. (6) and (7) imply that the long-run cointegration coefficient is allowed to vary at β with variance σ^2_β, or $\beta_t = \beta + \xi_t$ where $\xi_t \sim N(0, \sigma^2_\beta)$. Thus, the long-run equilibrium can be rewritten as $\eta^*_t = \mathrm{RR}_t - \beta \cdot \mathrm{CDS}_t - \xi_t\mathrm{CDS}_t = \eta_t - \xi_t\mathrm{CDS}_t$ where $\eta_t = \mathrm{RR}_t - \beta \cdot \mathrm{CDS}_t$. Hence, Eqs. (6) and (7) can be viewed as a specification of conventional error correction model plus with heteroskedastic error terms, that is,

$$\begin{pmatrix} \Delta \mathrm{RR}(i,\mathrm{US})_t \\ \Delta \mathrm{CDS}(i,\mathrm{US})_t \end{pmatrix} = \begin{pmatrix} a_{\Delta \mathrm{RR}} \\ a_{\Delta \mathrm{CDS}} \end{pmatrix} + \begin{pmatrix} \alpha_1 \\ \alpha_2 \end{pmatrix} \cdot \eta_{\mathrm{US},t-1} + \sum_{k=1}^{K} \Phi_k \begin{pmatrix} \Delta \mathrm{RR}(i,\mathrm{US})_{t-k} \\ \Delta \mathrm{CDS}(i,\mathrm{US})_{t-k} \end{pmatrix} + \Theta X_t + \begin{pmatrix} \varepsilon^*_{\Delta \mathrm{RR},t} \\ \varepsilon^*_{\Delta \mathrm{CDS},t} \end{pmatrix} \tag{8}$$

and

$$\begin{pmatrix} \Delta \mathrm{RR}(i,\mathrm{EU})_t \\ \Delta \mathrm{CDS}(i,\mathrm{EU})_t \end{pmatrix} = \begin{pmatrix} a_{\Delta \mathrm{RR}} \\ a_{\Delta \mathrm{CDS}} \end{pmatrix} + \begin{pmatrix} \alpha_1 \\ \alpha_2 \end{pmatrix} \cdot \eta_{\mathrm{EU},t-1} + \sum_{k=1}^{K} \Phi_k \begin{pmatrix} \Delta \mathrm{RR}(i,\mathrm{EU})_{t-k} \\ \Delta \mathrm{CDS}(i,\mathrm{EU})_{t-k} \end{pmatrix} + \Theta X_t + \begin{pmatrix} \varepsilon^*_{\Delta \mathrm{RR},t} \\ \varepsilon^*_{\Delta \mathrm{CDS},t} \end{pmatrix} \tag{9}$$

where $\begin{pmatrix} \varepsilon^*_{\Delta \mathrm{RR},t} \\ \varepsilon^*_{\Delta \mathrm{CDS},t} \end{pmatrix} = \begin{pmatrix} \varepsilon_{\Delta \mathrm{RR},t} \\ \varepsilon_{\Delta \mathrm{CDS},t} \end{pmatrix} + \begin{pmatrix} \alpha_1 \\ \alpha_2 \end{pmatrix} (-\xi_t \cdot \mathrm{CDS}_t)$. As $\varepsilon_{\Delta \mathrm{RR},t}$, $\varepsilon_{\Delta \mathrm{CDS},t}$, and ξ_t are normally distributed, $\varepsilon^*_t = (\varepsilon^*_{\Delta \mathrm{RR},t}, \varepsilon^*_{\Delta \mathrm{CDS},t})'$ is bivariate normal with mean zero and conditional variance–covariance matrix $H_t = E(\varepsilon^*_t\varepsilon^{*\prime}_t | \mathfrak{I}_{t-1})$. The specification in Eqs. (8) and (9) can be estimated by the maximum likelihood method (MLE). Following the procedure in Fong and Li (2004), the conditional likelihood function is defined as,

$$l(\theta) = -\frac{1}{2}\sum_{t=1}^{T}\left\{ \log|H_t| + \varepsilon^{*\prime}_t H_t \varepsilon^*_t \right\} \tag{10}$$

where θ is a vector of all parameters in the model. Given the past information $\mathfrak{I}_{t-1}$, the MLE estimates and standard errors are derived from the first derivative of the likelihood function and estimated Hessian matrix. In our bivariate case, the first and second derivatives of the likelihood function are found to be:

$$\frac{\partial l}{\partial \theta} = -\sum_{t=1}^{T} U^{*}_{t-1} H_t^{-1} \varepsilon^{*}_t \tag{11}$$

and

$$\frac{\partial^2 l}{\partial \theta' \partial \theta} = -\sum_{t=1}^{T} \left[\frac{\partial U^{*}_{t-1}}{\partial \theta'} H_t^{-1} \varepsilon^{*}_t + \left(U^{*}_{t-1} H_t^{-1} \otimes I_1 \right) \frac{\partial \varepsilon^{*}_t}{\partial \theta'} \right] \tag{12}$$

where $U_{t-1}^{*} = [(A' \otimes J' Z_{t-1})', I_2 \otimes \tilde{U}'_{t-1}]'$ and $\tilde{U}_{t-1} = [(BZ_t)', \Delta Z_{t-1}', \ldots, \Delta Z_{t-p+1}']'$, $Z_t' = (RR_t, CDS_t)$, $A' = (\alpha_1, \alpha_2)$, $B = (1, -\beta)$, $J' = (0, 1)$, $\otimes$ is the Kronecker product, I_2 is a 2×2 identity matrix. The asymptotic representation of $\hat{\theta} - \theta$ can be obtained by the first-order Taylor expansion of Eq. (11) at $\hat{\theta}$, i.e.,

$$\frac{\partial l}{\partial \theta} = \frac{\partial l}{\partial \theta}\Big|_{\theta=\hat{\theta}} + \left(\theta - \hat{\theta}\right) \frac{\partial^2 l}{\partial \theta' \partial \theta}\Big|_{\theta=\hat{\theta}} + o_p(1). \tag{13}$$

Since $\partial l / \partial \theta|_{\theta=\hat{\theta}} = 0$ in estimation, Eq. (13) becomes

$$\hat{\theta} - \theta = \left(\frac{\partial^2 l}{\partial \theta' \partial \theta}\Big|_{\theta=\hat{\theta}} \right)^{-1} \frac{\partial l}{\partial \theta} + o_p(1). \tag{14}$$

Following Ahn and Reinsel (1990) and Li, Ling, and Wong (2001), Fong and Li (2004) suggest that $\hat{\theta}$ is a consistent estimator and converges to a non-degenerating random process specified in terms of the Brownian processes.

Compared with a conventional error-correction model, the model used in this paper generalizes the assumption of a constant cointegration (i.e., β) to that of a random cointegration (i.e., β_t). In other words, when the variance of β_t (i.e., σ_β^2) in this specification is zero, the model is reduced to the conventional error-correction model. When this variance is different from zero, the cointegration relationship (i.e., β_t) is allowed to vary at the constant cointegration (i.e., β) with the variance σ_β^2 over time, which suggests that a stationary combination of integrated variables is no longer linear since it allows non-linear adjustment to the long-run equilibrium. Thus, a larger variance can facilitate a larger deviation of the cointegration relationship from the constant cointegration.[11] The motivation of this extension is consistent with Hansen (1992)'s cointegration model in which the errors display non-stationary variances; the regime-shifting models in the long-run equation in Gregory and Hansen (1996), Kejiriwal and Perron (2008, 2010), Esteve, Navarro-Ibanez, and Prats (2013), and Zhang and Li (2014); and the threshold autoregression models in Balke and Fomby (1997), Hansen and Seo (2002), and Al-Abri and Goodwin (2009) in which the long-run equilibrium error is allowed to move freely within a given range but mean-reverting outside the range.

Therefore, this simple extension can serve two purposes. First, the new long-run specification provides more flexibility than the conventional one when modeling multivariate financial time series linked by non-linear long-run relationships. In addition to the observable patterns recognized by the literature (e.g., regime switching or structural break due to interventions), any unobservable non-linear patterns are reflected in the variance of the random cointegration. Secondly, this model helps one to understand how the two markets interact in short-run over time in the presence of a non-linear long-run relationship. This is achieved by deriving a time-varying conditional correlation from the conditional variance–covariance matrix H_t in the new specification. Specifically, the conditional correlation at time t is obtained by

$$Corr_t = \frac{\hat{h}_{12t}}{\sqrt{\hat{h}_{11t}\hat{h}_{22t}}} \tag{15}$$

where $\hat{h}_{ijt}$ ($i = 1{,}2$ and $j = 1{,}2$) are elements of the estimated variance covariance matrix H_t which are in terms of the speeds of adjustment (i.e., α_1 and α_2), variance of random coefficient (i.e., σ_β^2) and past prices in the two markets. Its characteristic is thus similar to those of Bauwens, Deprins, and Vandeuren (1997), Silvapulle and Podivinsky (2007), and Kurita (2013) who specify their cointegrated VAR models with generalized autoregressive conditional heteroskedasticity (GARCH) and ARCH errors.

To characterize the contribution of each market to price discovery, we compute a ratio using the speeds of adjustment of α_1 and α_2. According to Gonzalo and Granger (1995) and further explored by Ammer and Cai (2011) and Palladini and Portes (2011), if one market always lags the other, then the speed of the adjustment of the lagging market should be different from zero. This is because the lagging market takes time to remove pricing errors, while the leading market has already reflected all information in the prices.

[11] One example is that when the prices deviate from the long-run relationship drastically and persistently, the variance of our model will be non-zero so that the model can capture non-linear change due to deviation of the cointegration relationship from the constant cointegration for a period of time. We have elaborated this in the revised paper.

If both coefficients are different from zero with the expected signs, the relative magnitude of the two coefficients reveals which market leads the price discovery process more. Following Gonzalo and Granger (1995), the ratio, denoted by GG, is defined as:

$$GG = \frac{\alpha_2}{\alpha_2 - \alpha_1} \tag{16}$$

with a lower bound of 0 and upper bound of 1. This ratio can be interpreted as the fraction of overall price movements in the currency options market relative to overall movements in the sovereign CDS market. When the currency option market leads and the sovereign CDS market follows in price discrepancy corrections, the GG measure will be closer to 1. When the sovereign CDS market leads in price discovery, the GG measure will be closer to 0.

3. Data descriptions

3.1. Data, summary statistics and cointegration tests

For the purpose of comparison, we use the perspectives of USD-based and euro-based investors respectively in the analysis. We obtain weekly over-the-counter 25-delta risk reversals of USD–yen, USD–euro, USD–Swiss franc, euro–yen, and euro–Swiss franc options at the 3-month maturity, and the 5-year sovereign CDS spreads of the eurozone, Japan, Switzerland and the US from July 2007 to May 2013.[12] The tenors of these derivative instruments are commonly used as benchmarks of their respective markets. The eurozone CDS spread is the median sovereign CDS spread of the eleven eurozone countries.[13] As currency option prices reflect market expectations of the exchange rate between two economies' currencies, CDS spreads are thus expressed as the differences (in short "relative CDS spreads") between the CDS spreads of the US and individual economies (i.e., the CDS spread of the US minus that of another country) from the perspective of USD-based investors. Similarly, from the perspective of euro-based investors, the relative CDS spreads are expressed as differences between the eurozone and those of individual economies. Note that the USD-euro risk reversal and the US–eurozone relative CDS spread can offer both USD- and euro-based investors' perspectives at the same time.

The data are depicted in Fig. 1. The risk reversals and relative CDS spreads move similarly in trend during most of the time. In some cases, they could however deviate to an extent that the deviation is almost enough to break the long-run cointegrating relationship between the two markets. To see this more clearly, a scatter plot of them, which reflects their unconditionally linear long-run relationship, is depicted in Fig. 3. For examples, the linear relationship between the USD–yen risk reversal and the US relative CDS spread (Fig. 3a) appears to be strong, but it is undermined by some outliers on the top right hand corner to some extent. The linear relationship between the euro–yen risk reversal and eurozone relative CDS spread (Fig. 3d) appears to be undermined by the outliers (on the top left hand corner) to a larger extent, but the linear relationship remains noticeable from most of the observations. Between the USD–Swiss franc and the US relative CDS spreads (Fig. 3b), the relationship seems to be ambiguous given that the correlation between the two markets is negative.

Table 1 provides summary statistics for these time series of the data in levels and changes. Over the sample period, the average sovereign CDS spreads of the US are 31.30 bps, 83.40 bps, and 8.53 bps lower than those of Japan, the eurozone, and Switzerland respectively, while the average risk reversals of the yen, euro, and Swiss franc against the USD are +1.88% points, −1.30% points, and −0.28% points respectively. The average sovereign CDS spread of the eurozone are 48.21 bps and 92.07 bps higher than those of Japan and Switzerland respectively, while the average risk reversals of the yen and Swiss franc against the euro were 3.24% points and 1.07% points respectively. Table 1 also reports the Augmented Dickey–Fuller (ADF) and Phillip–Perrons (PP) test results. Using a maximum of four lags (i.e., a 4-week period), both tests fail to reject at the 5% level the presence of a unit root for all levels of the risk reversals and relative CDS spreads. In addition, both tests for the first differences are significant at the 5% level. Thus, the levels of risk reversals and relative CDS spreads appear to be non-stationary while the changes appear to be stationary. This suggests that all pairs of risk reversals and relative CDS spreads are $I(1)$ (i.e., integrated of same order 1), which satisfies the requirement for the variables being cointegrated.

To test the cointegration between the relative sovereign CDS spreads and risk reversals, we use the Engle–Granger single-equation test and the Johansen cointegration trace test. The methodology proposed by Engle and Granger (1987) is regarded as an easy and super-consistent method of estimation, which determines whether the residuals of the linear combination among the cointegrated variables estimated from the ordinary least squares method are stationary. The Johansen cointegration test allows testing for cointegration with unknown cointegrating vectors, estimating both the cointegrating vectors and determining the number of cointegrating relationships. It is noted that there can be at most one independent linear combination that is stationary in a two-variable case.

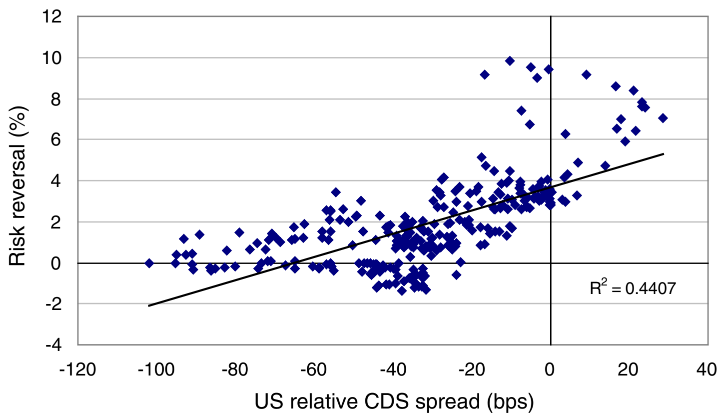
Table 2 reports both the cointegration tests between the risk reversals and relative CDS spreads. In the Engle–Granger test, we employ the ADF and PP tests to check whether the residuals of the regression of the risk reversals on relative CDS spreads are stationary. The critical values of the tests are based on MacKinnon (1996) and the lag length is determined by the Schwartz criterion. The results for the Swiss franc against the USD and euro are significant at the 10% level, while the others are statistically significant at the 5% level.

[12] The CDS data are from Bloomberg.

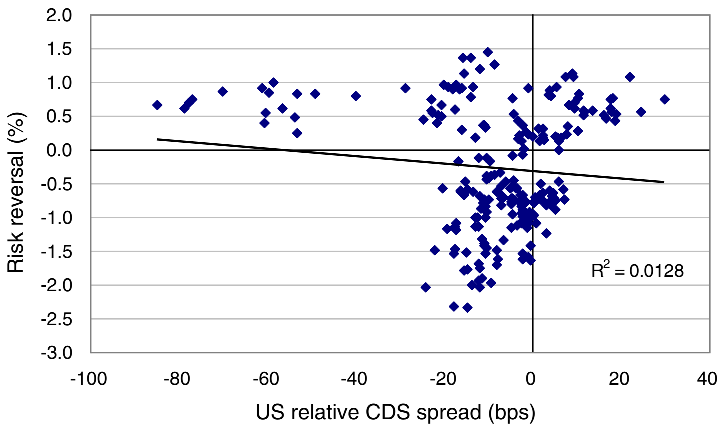
[13] These countries include Austria, Belgium, Finland, France, Germany, Greece, Ireland, Italy, Portugal, the Netherlands and Spain. There is no active sovereign CDS on Cyprus, Luxembourg, Malta, Slovakia, and Slovenia.

Thus, we reject the null hypothesis that the relative CDS spreads and risk reversals are not cointegrated in favor of the alternative hypothesis that there is at least one cointegrating vector. In the Johansen test, we find that only the cointegrating relationship between the USD–euro risk reversal and the US relative CDS spreads is conclusive since only one cointegrating vector exists between the two

a. Difference between US and Japanese sovereign CDS spread and risk reversal of USD-yen



b. Difference between US and Swiss sovereign CDS spread and risk reversal of USD-Swiss franc



c. Difference between US and eurozone sovereign CDS spread and risk reversal of USD-euro[2]

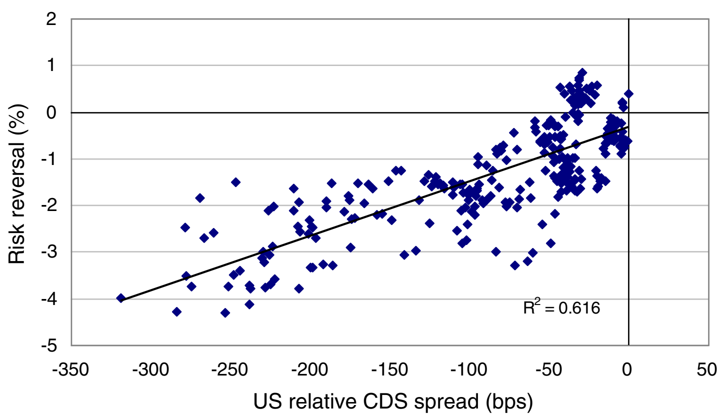
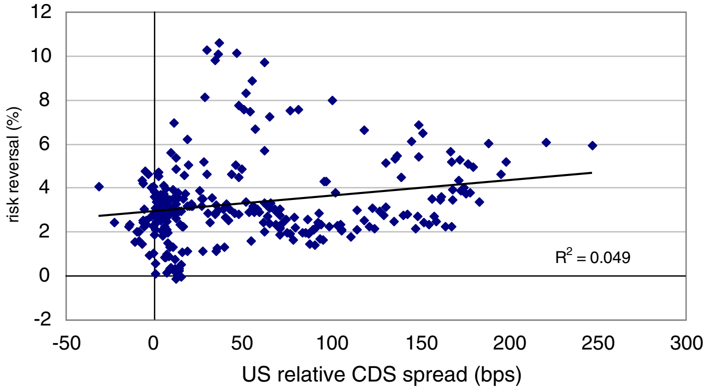


Fig. 3. Scatter plot of relative sovereign CDS spreads and risk reversals.
Notes:

1. The risk reversal is the implied volatility of an out-of-the-money USD (euro) put minus that of an out-of-the-money USD (euro) call at the 25% delta.
2. The USD-euro risk reversal and the US-eurozone sovereign CDS spreads can offer perspectives of both USD- and euro-based investors at the same time.

d. Difference between eurozone and Japanese sovereign CDS spread and risk reversal of euro-yen



e. Difference between eurozone and Swiss sovereign CDS spread and risk reversal of euro-Swiss franc

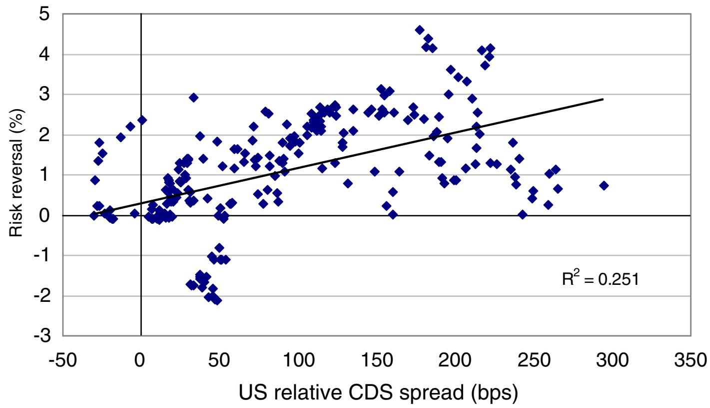


Fig. 3 (*continued*).

markets (i.e., rejecting the null hypothesis of none but not rejecting the null of having at most one cointegrating vector at the 5% significant level). The Johansen test results of the other pairs however are found to be inconclusive (i.e., rejecting or not rejecting both hypotheses of none and at most one cointegrating vector at the same time). The two tests' results suggest different conclusions, attributable to the fact that some long-run relationships between the risk reversals and sovereign relative CDS spreads are undermined by the outliers observed during the crisis periods as shown in Fig. 3. Attfield (2003) finds that the Johansen cointegration test usually over-rejects the null hypothesis of no cointegration between multivariate time series in the presence of structural break or heteroskedasticity. Maki (2013) suggests that the Engle–Granger approach also suffers the same problem, however, it is generally regarded as one of the less problematic tests when the residual has a GARCH effect. His simulation results show that the empirical size is only slightly higher than the nominal size when using the Engle–Granger approach. In view of this, we do not reject the hypothesis that cointegration exists in all our data and proceed to the step of estimation. If this assumption is too strong, the generalized error-correction model, despite controlling for impacts of outliers, will show insignificant coefficients of the long-run cointegration.

To test for the constancy of the cointegrating coefficients, we employ the score-based test derived by Fong and Li (2004).[14] The test primarily works on scores which are the first derivative of the likelihood function with respect to the parameters. Theoretically, when the scores with respect to all parameters are close to zero, the fitted model can be regarded as adequate for describing the data. In this study, the null hypothesis of the test is $\sigma_\beta^2 = 0$ against the alternative hypothesis of $\sigma_\beta^2 \neq 0$. The test results are summarized in Table 2. At a 10% level of significance, all test statistics are larger than the critical value,[15] indicating that the null hypothesis of constant cointegration is rejected in all of our examples.

[14] The score test is an alternative to the likelihood ratio test and Wald test. It is regarded as the locally most powerful test when the estimated parameters are close to their true values. The main advantage of the test is that it does not require an estimate of the information under the alternative hypothesis or unconstrained maximum likelihood. This makes the test feasible when the unconstrained maximum likelihood estimate is a boundary point in the parameter space.

[15] The critical value can be found in Fong and Li (2004). In our examples, the critical values are 2.348 and 1.677 at 5% and 10% levels of significance respectively.

Table 1
Descriptive statistics for risk reversals and relative sovereign CDS spreads.

	25-delta 3-month risk reversal against USD						25-delta 3-month risk reversal against euro			
	Japanese yen		Swiss franc		Euro		Japanese yen		Swiss franc	
Statistic	Level	Change	Level	Change	Level	Change	Level	Change	Level	Change
Mean	1.88	−0.01	−0.28	−0.01	−1.30	0.00	3.24	−0.01	1.07	−0.02
Median	1.51	−0.04	−0.47	0.01	−1.31	0.01	2.96	−0.02	1.13	−0.01
Maximum	9.82	1.72	1.45	0.81	0.85	1.22	10.61	2.12	4.60	1.53
Minimum	−1.35	−1.15	−2.34	−1.08	−4.30	−0.82	−0.14	−1.70	−2.12	−1.63
Std. dev.	2.17	0.34	0.90	0.24	1.11	0.27	1.82	0.44	1.40	0.35
Skewness	1.30	0.65	−0.07	−0.59	−0.43	0.55	1.46	0.34	−0.18	−0.02
Kurtosis	5.22	6.80	1.96	6.22	2.75	5.74	6.30	6.51	2.97	7.27
ADF test statistic	−1.68	−12.35**	−2.70	−11.25**	−2.22	−8.82**	−2.10	−12.26**	−2.33	−11.05**
Phillips–Perron test statistic	−1.80	−12.02**	−2.52	−11.25**	−2.18	−14.00**	−2.48	−12.62**	−1.95	−10.71**
Observations	285	284	227	226	285	284	306	305	227	226
	Difference between US and country's sovereign CDS spreads						Difference between Eurozone and country's sovereign CDS spreads			
	Japan		Switzerland		Eurozone (median)		Japan		Switzerland	
Statistic	Level	Change	Level	Change	Level	Change	Level	Change	Level	Change
Mean	−31.30	−0.13	−8.53	0.16	−83.40	−0.11	48.21	0.00	92.07	−0.01
Median	−31.61	−0.05	−4.81	−0.11	−48.99	−0.08	19.35	0.04	74.31	−0.54
Maximum	28.45	24.55	29.68	20.36	0.08	44.31	247.33	43.38	294.44	−42.04
Minimum	−101.84	−30.41	−84.94	−37.68	−318.79	−49.77	−31.63	−49.14	−30.31	−34.77
Std. dev.	25.25	5.75	18.20	4.95	74.97	11.36	57.25	11.47	78.54	11.54
Skewness	−0.38	−0.41	−1.81	2.13	−1.10	−0.29	1.13	−0.07	0.53	0.26
Kurtosis	3.07	7.64	7.55	21.08	3.13	6.61	3.26	6.06	2.21	5.04
ADF test statistic	−2.12	−16.05**	−2.42	−14.29**	−1.66	−13.81**	−2.00	−14.28**	−1.48	−12.69**
Phillips–Perron test statistic	−2.23	−16.05**	−2.48	−14.37**	−1.55	−13.70**	−1.85	−14.19**	−1.18	−12.66**
Observations	285	284	227	226	285	284	306	305	227	226

Notes:
1. ** and * indicate significance at levels of 5% and 10% respectively.
2. Both tests check the null hypothesis of unit root existence in the time series, assuming non-zero mean in the test equation.

3.2. Exogenous macro-financial variables

Recent research finds that sovereign credit risk interacts strongly with global and regional financial risk factors. Longstaff, Pan, Pedersen, and Singleton (2011) show that sovereign CDS spreads are primarily driven by common factors, including the US stock and high-yield bond markets and global risk premiums, whereas Pan and Singleton (2008) find that the spreads are related to investors' risk appetite associated with global event risk, financial market volatility and macroeconomic policy. Therefore, it is important to identify whether the sovereign CDS spreads and risk reversals of the four economies in this study remain cointegrated in the presence

Table 2
Tests for cointegration between risk reversals and relative sovereign CDS spreads.

	USD-based risk reversal			Euro-based risk reversal	
	Japanese yen	Euro	Swiss franc	Japanese yen	Swiss franc
Engle–Granger single-equation test					
(null hypothesis: residual has an unit root)					
ADF test statistic	−2.96**	−3.79**	−2.77*	−2.95**	−2.77*
Phillips–Perron test statistic	−2.90**	−2.94**	−2.53	−2.96**	−2.86*
Johansen cointegration trace test					
Null hypothesis: zero cointegrating vector	4.39	13.88**	28.37**	7.55	7.73
Null hypothesis: at most 1 cointegrating vector	1.36	1.12	5.11**	2.66	0.67
Score-based test					
Null hypothesis: $\sigma_\beta^2 = 0$	10.72**	1.70*	50.47**	8.31**	21.27**

Notes:
1. ** and * indicate significance at levels of 5% and 10% respectively.
2. The cointegration test uses the Augmented Dickey–Fuller and Phillips–Perron tests to check the null hypothesis that the residuals of the regression of a risk reversal on a relative CDS spread are non-stationary assuming non-zero mean in the test equation. The critical value of the test is obtained from MacKinnon (1996).
3. The score-based test checks the constancy of cointegrating coefficients. For a series length of 300, the critical values are 2.348 and 1.677 at 5% and 10% levels of significance respectively.

of other macro-financial factors. To address this issue, we employ a set of macro-financial factors as control variables, including the following factors[16]:

(i) *Interest rate differential*. In a currency carry trade, an investor borrows in a low yielding currency and invests in a high yielding currency. To realize the carry in the trade, investors are required to hold the position for some time. The risk to the carry trade is an adverse price movement in the level of the exchange rate, i.e., a currency crash. As found in Brunnermeier, Nagel, and Pedersen (2009), the expected exchange rate movements between high-interest-rate and low-interest-rate currencies are negatively skewed due to the crash risk of sudden unwinding of carry trades and reflected in the risk reversals of their out-of-the-money currency options. Therefore, if the interest rate differential of the two currencies ($r_{\text{USD}} - r_{\text{other}}$) or ($r_{\text{euro}} - r_{\text{other}}$) increases, the risk reversal is expected to increase as a result of hedging against the crash risk. We use the 1-month LIBORs for the interest rate differential.

(ii) *US dollar volatility*. The implied volatility of an exchange rate is essentially linked to the anticipated uncertainty on the values of both currencies in the pair. Therefore, we use the US dollar index (DXY), a weighted average of the dollar's value relative to a basket of foreign currencies, to capture the actual volatility attributable to the dollar factor. We proxy the volatility of the US dollar (R^2_{USD}) as the ex-post squared return of the index.

(iii) *Global risk appetite*. We use the market volatility index (*VIX*) of the US S&P 500 index (EURO STOXX 50 index) to gauge the global risk appetites of USD-based (euro-based) investors in the financial market.[17] An increase in the VIX index is usually associated with heightened volatility across different asset classes in particular equities. Currency option-implied volatility shares commonality with the VIX index as a measure of investors' aversion to volatility exposure and hence their willingness to put capital at risk.

(iv) *Funding liquidity constraint*. Another potential determinant of the risk reversals is the sudden unwinding of currency carry trades. We follow Brunnermeier et al. (2009) and use the USD (euro) TED spread (TED), the difference between the 3-month interbank rate and the yield of the 3-month Treasury bill (German bunds), to capture traders' funding liquidity constraint. When funding liquidity is tight, as reflected by a widened TED spread, traders are forced to repatriate funds to a safer currency.

(v) *Macro-financial condition*. To capture the broad changes in the macro-financial condition, we include two measures from the stock and bond markets that have been used by Collin-Dufresne, Goldstein, and Martin (2001) and Cao et al. (2010). Regarding the stock market variables, we use the weekly returns of the S&P 500 index (*US stock return*), STOXX European 600 index (*EU stock return*), Swiss Market Index, and Nikkei 225 (*Country's stock return*). Conventionally, a negative stock market return indicates a weaker economic outlook and puts downward pressure on the corresponding currency. For the bond market variables, we use the term spreads between 10-year and 2-year yields of the US Treasury bonds (*US term spread*), German bunds (*EU term spread*), Switzerland government bonds, and Japan government bonds (*Country's term spread*). Collin-Dufresne et al. (2001) interpret the term spread (i.e., the slope of a yield curve) as a proxy for the overall state of an economy. An upward sloping yield curve indicates future economic growth, whereas a flattening yield curve reflects a poor economic prospect.

4. Empirical results

In the MLE estimation, all risk reversals and relative CDS spreads are first normalized (i.e., $z = (x - \mu)/\sigma$) because the risk reversals and relative CDS spreads have very different scales and therefore this standardization makes the pass-through between the risk reversal and relative CDS spreads among selected currency pairs comparable. Initial values of the estimates are derived from the least squares method. An autoregressive order of four (i.e., the parameter K) is chosen, in other words, it assumes that all autoregressive impacts of risk reversals and relative CDS spreads are insignificant after four weeks. All coefficients of each error correction model are estimated in one-step but the estimation results for the long-run and short-run are reported separately in Tables 3 and 4.

To check the models' adequacy, we first standardize the residuals in Eqs. (8) and (9) by the time-varying volatility estimated from the heteroskedastic error terms. We then conduct the Ljung–Box test for zero autocorrelations and cross-correlations between the two time series of standardized residuals. Since the reference distribution for the test statistics is not known in our case, the test results are benchmarked against the Chi-squared distribution. As reported in Table 4, diagnostic test statistics show that all estimated models are generally adequate. Except for the euro–yen's risk reversal, all test statistics are less than the benchmark at the 10% level of significance, suggesting no significant autocorrelations in the standardized residuals and no significant cross-correlations between risk reversals and relative CDS spreads. The exception one, whose residual autocorrelation is 27.656, is only slightly more than the benchmark (i.e. 23.5), so we still consider this estimated model marginally acceptable.

Table 3 reports the estimated cointegrating vectors between the risk reversals and relative CDS spreads. For USD-based investors, the coefficients β for Japan, the eurozone and Switzerland are 0.6766, 0.9560, and 4.4072 respectively, in which the last coefficient for Switzerland is insignificant.[18] Their corresponding variance estimates (i.e., σ^2_β) are 3.0037^2, 1.6735^2, and 4.4639^2 respectively. For

[16] We obtain data for these additional variables from Bloomberg.

[17] Collin-Dufresne et al. (2001) and Zhang, Zhou, and Zhu (2009) use the US's VIX index as a measure of market-level volatility and find a strong relationship with firm-level credit spreads. Pan and Singleton (2008) view the VIX index as a measure of investors' risk aversion for the event risk in credit markets.

[18] The coefficient for USD–Swiss franc is found to be very large, suggesting that the risk reversal could be very responsive to the relative CDS spread in some occasions. This coefficient, however, is found to be statistically insignificant. Consistent with this insignificance, the linear relationship between the two markets shown in the scatter plot appears to be ambiguous (see Fig. 3b). This may be arising from the fact that most Swiss franc's investors are euro-based and they tend to hold more euro than USD so the long-run relationship between the relative sovereign CDS and currency options markets may not be always strong.

Table 3
Estimates of cointegrating vectors (i.e., the long-run part of Eqs. (8) and (9)).

	USD-based risk reversal (i.e. the long-run part of Eq. (8))			Euro-based risk reversal (i.e. the long-run part of Eq. (9))	
	Japan/yen	Eurozone/euro	Switzerland/Swiss franc	Japan/yen	Switzerland/Swiss franc
Relative CDS spread (β)	0.6766**	0.9560**	4.4072	0.7444*	1.0239*
σ_β	3.0037**	1.6735*	4.4639*	7.9722*	3.9888*

Note: ** and * indicate significance at levels of 5% and 10% respectively.

euro-based investors, the coefficient β for Japan and Switzerland is 0.7444 and 1.0239 respectively. The corresponding variance estimates are 7.9722^2 and 3.9888^2 respectively. The results in Table 3 have two interesting implications. First, the sovereign CDS and currency option markets are found to be positively cointegrated (i.e., coefficient β is positive), implying that, other things being equal, a sharp rise in the sovereign risk of one economy would weaken its exchange rate expectation eventually. Secondly, among the significant coefficients, the magnitude of estimated β ranges from two-third to unity, implying that the pass-through of shock from sovereign CDS to currency option markets is substantial in the long run. These findings support the first hypothesis that, in these developed economies, the relative sovereign credit risk has a considerable impact on the market expectations of their exchange rates in the long run.

In the short run, the interconnectivity between the prices of the two markets measured by the conditional correlations of the two markets is found to be time varying. As shown in Fig. 4, the correlation is generally positive, which is consistent with the intuition that an increase in the relative CDS spread of the US (the eurozone) will raise the risk reversals such that high option premiums (i.e., high option-implied volatilities) are paid for USD-puts (euro-puts). This implies that the currency option prices of the currency pairs in this study incorporate sovereign credit risks. Furthermore, in times of uncertainty, the positive correlation between the two markets tends to fluctuate drastically. For example, the conditional correlation of US–Japan (see Fig. 4a) falls sharply to negative and then rises back

Table 4
Estimation results of the bivariate vector error-correction model in Eqs. (8) and (9).

	USD-based risk reversal/US relative CDS spread						Euro-based risk reversal/eurozone CDS spread			
	Yen	Japan	Euro	Eurozone	Swiss franc	Switzerland	Yen	Japan	Swiss franc	Switzerland
	ΔRR_t	ΔCDS_t	ΔRR_t	ΔCDS_t	ΔRR_t	ΔCDS_t	ΔRR_t	ΔCDS_t	ΔRR_t	ΔCDS_t
Short-term variable										
Speed of adjustment η_{t-1} (i.e. α_1 and α_2)	−0.0275**	0.0077	−0.0029	0.0432*	−0.0000	0.0438*	0.0033	0.0151*	−0.0271*	0.0171*
ΔRR_{t-1}	0.2792**	−0.0128	0.2995*	−0.0010	0.3539*	0.0387	0.2805*	0.0057	0.3426*	0.0480*
ΔRR_{t-2}	0.0981**	−0.0335	−0.1564*	0.0381	−0.1726*	0.0301	0.0828*	0.0335	−0.0464	−0.0548*
ΔRR_{t-3}	−0.1087**	0.1171*	0.2086*	0.0845*	0.1852*	−0.0078	0.0059	0.0101	−0.0612	0.0750*
ΔRR_{t-4}	−0.0872**	0.0819	−0.1286*	−0.0108	−0.0925*	0.0737*	−0.1405*	0.0592*	−0.0809*	−0.0180
ΔCDS_{t-1}	−0.0035	0.0693*	−0.0936	0.2450*	−0.0339	0.1420*	0.0926*	0.1539*	0.2104*	0.2429*
ΔCDS_{t-2}	0.0689**	−0.0354	−0.1532*	−0.1170*	0.0775*	−0.0678	0.0031	−0.0363	−0.3246*	−0.1058*
ΔCDS_{t-3}	−0.0571**	0.1044**	0.0795	−0.0334	−0.0165	−0.1301*	0.0516	−0.0245	0.1076	0.0426
ΔCDS_{t-4}	−0.0371*	−0.0438	−0.0075	−0.0302	−0.0246	−0.1398*	−0.0171	−0.0160	−0.0841	−0.0421
Const	−0.0069	−0.0020	−0.0041	0.0078	−0.0115	0.0268*	0.0008	0.0047	−0.0002	0.0096*
Interest rate differential	0.5597	−0.6787	−0.0264	0.1691	−0.3714	−6.4639*	2.6122*	1.4315*	5.1011*	2.6763*
Dollar squared return	−0.0038	−0.0229	−0.0901*	−0.0005	−0.0966*	−0.0155	0.0363*	−0.0031	0.0280	−0.0049
US or EU volatility index	0.0147	−0.0108	−0.0022	−0.0650*	0.0354	0.0042	0.0703*	−0.0571*	0.1397*	−0.0647*
US or EU Ted spread	−0.1077**	0.0499**	−0.0154	0.1102*	−0.0436	−0.0183	−0.1466*	0.0166	−0.2649*	−0.0243
US or EU stock return	−0.8728**	0.8978	0.8149	−0.0933	−1.8774	−5.9743*	−1.1388	−1.1074	−0.4570	−2.6000*
US or EU term spread	0.0170**	0.0239	0.0301*	−0.0077	−0.0097	−0.1034*	−0.0244*	0.0262*	0.0104	−0.0099
Country's stock return	0.0143	−0.0130	−0.0049	−0.0293*	−0.0001	−0.1580*	−0.0100	0.0229	−0.1420*	0.0274
Country's term spread	0.0083	0.0142	0.0299	0.0270*	−0.0341	0.0095	−0.0090	0.0008	0.1595*	−0.0423*
Variance related										
σ_{CDS} and σ_{RR}	0.0403	0.0979**	0.4481*	0.2117*	0.0283	0.2392*	0.2025*	0.1414*	0.1900*	0.1078*
r	0.2156**		0.1106*		0.1006*		0.0543		0.4304*	
Ljung–Box test statistic										
Residuals' autocorrelation	4.5902	6.0496	18.2470	0.8065	3.3170	2.3064	27.6560	14.1250	17.8770	7.6968
Residuals' cross-correlation	16.55		19.40		11.04		17.61		15.87	
GG statistic	0.2188		0.9371		0.9999		0.8265		0.3869	

Notes: 1. ** and * indicate significance at levels of 5% and 10% respectively. 2. The two Ljung–Box tests check the null hypotheses of jointly zero residuals' autocorrelations and jointly zero residuals' cross-correlations respectively up to lag 4 (i.e. the fourth week). The test statistics can be roughly compared to the Chi-squared distribution. With the degree of freedom of 16 (=22 × 4), the critical value is 23.5 at the 10% confidence level.

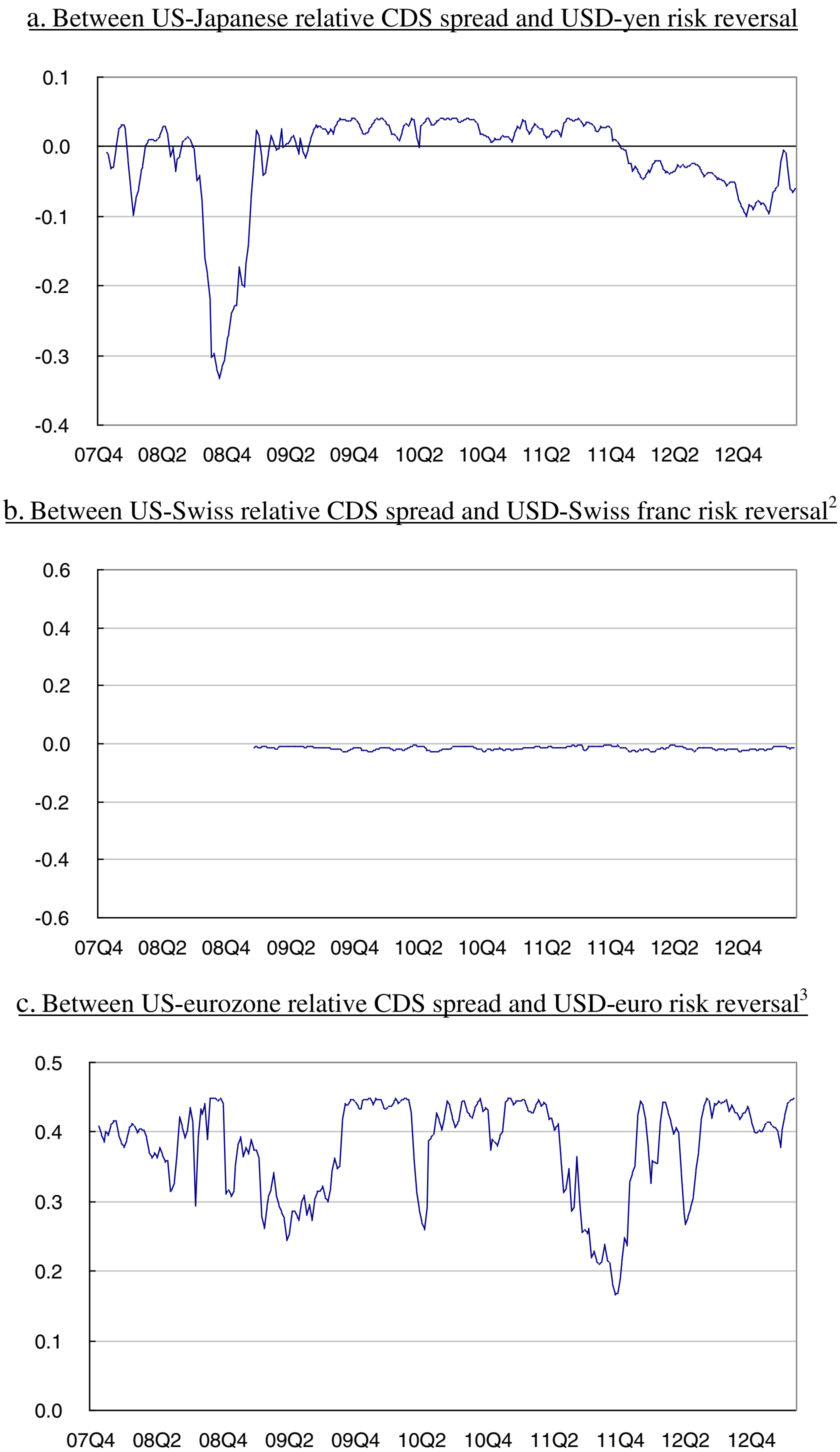
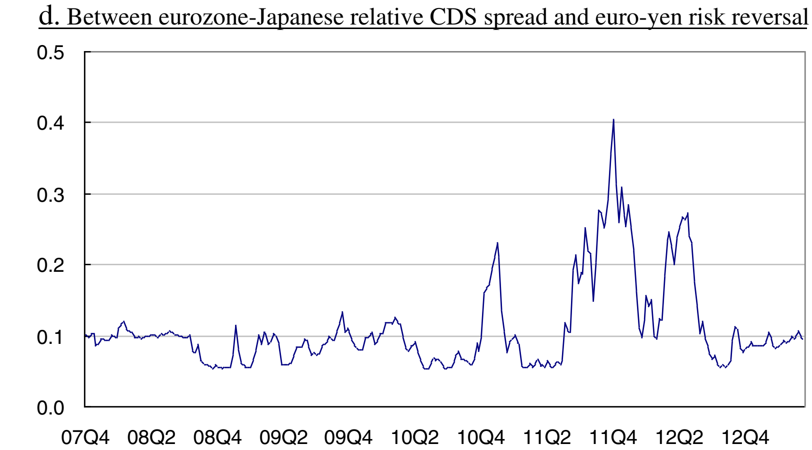


Fig. 4. Estimated conditional correlation between risk reversals and relative CDS spreads in short-run[1].
Notes:

1. The risk reversal is the implied volatility of an out-of-the-money USD (euro) put minus that of an out-of-the-money USD (euro) call at the 25% delta.
2. The estimated conditional correlation begins in January 2009 because the Switzerland's sovereign CDS spreads is only available since January 2009 in the data source (JP Morgan Chase).
3. The USD-euro risk reversal and the US-eurozone sovereign CDS spreads can offer perspectives of both USD- and euro-based investors at the same time.



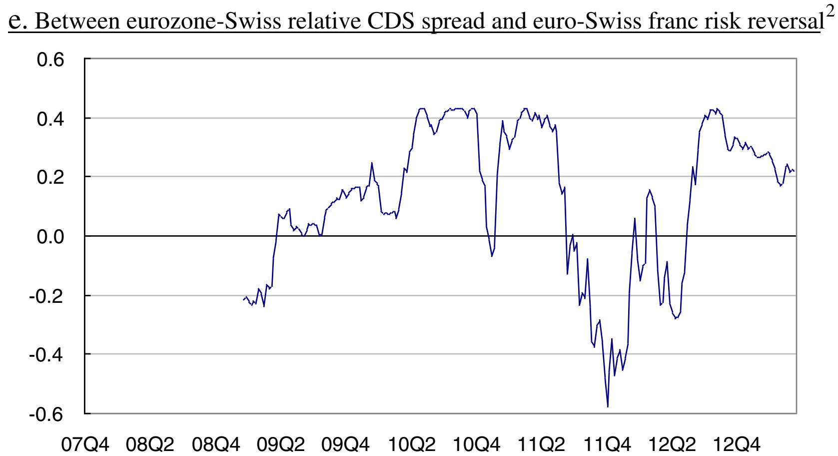


Fig. 4 (*continued*).

to the positive values in the period of 2008 Q3–2009 Q1, when the global financial crisis intensified after the Lehman default.[19] As shown in Fig. 1a, the risk reversal had surged for a short period of time in 2008 Q4, implying that market participants anticipated a stronger yen against the USD during the period. However, the relative US CDS spread decreased during the same period. The result reflects that the short-term relation between the relative CDS spread and risk reversal for the USD–yen pair is relatively weak and the positive relationship may not hold during market turbulence.

Another example is that the conditional correlation of eurozone–Japan (see Fig. 4d) rises substantially when the European sovereign debt crisis deepened in mid-2011. This suggests that the positive relationship between the two markets tends to strengthen during the period. Meanwhile, the conditional correlations of US–eurozone (see Fig. 4c) and eurozone–Switzerland (see Fig. 4e) weaken noticeably. The former case is attributable to the unresolved US debt-ceiling discussions which raised the specter of a potential technical default by the federal government (failure to pay the interest and/or principal of US treasury securities on time). [20] The latter case is arising from the fact that a sharp risk re-appraisal triggered a selloff in risky assets and prompted investors to seek safe havens amid worries over the deepened European sovereign debt crisis during the summer in 2011. As a result, the Swiss franc came under tremendous upward pressure. The rising risk reversal shows that the market had never been so one-sided in betting on a stronger Swiss franc (see Fig. 1e). Hence, in view of the potential impact on the economy and increasing risk of deflation, the SNB decided on 6 September 2011 to curb further appreciation of the currency. A series of interventions in the spot, forward, and option markets brought the Swiss franc below 1.2 as a cap against the euro and forced the volatility of the exchange rate sharply lower. Such measure pushed the conditional correlation to negative values for three quarters from 2011 Q4 to 2012 Q2. These examples support the second

[19] In other periods, the conditional correlation moves only within a narrow range of −0.1 and 0.1 (see Fig. 4a), reflecting that the short-term relation between the relative CDS spread and risk reversal for the USD–yen pair is relatively weak compared with the other currency pairs.

[20] In the summer of 2011, the credit rating agencies including the Fitch, Moody's and Standard & Poor's (S&P's) expressed concerns about very large budget deficits and rising indebtedness of the US government. Despite a settlement to raise the borrowing limit, S&P's lowered the credit rating of the US from AAA to AA+ on 5 August 2011, deciding pessimistic about its fiscal outlook. As a result, both the USD and euro were anticipated by market participants having risk of depreciation, such that their risk reversal became less correlated to the relative CDS spread for three quarters.

hypothesis that the interconnectivity between the sovereign CDS and currency option markets changes drastically and persistently in times of uncertainty.

The lead–lag relationship in price discovery is examined by studying the GG measures derived from the speed of adjustment. As reported in Table 4, the estimates (i.e., α_1) of US–Japan, US–eurozone, US–Switzerland, eurozone–Japan, and eurozone–Switzerland are 0.0077, 0.0432, 0.0438, 0.0151, and 0.0171 respectively. They are found all positive, suggesting that the relative CDS spreads will subsequently adjust upward to restore the long-run equilibrium when a deviation occurs. For the risk reversals, the estimates α_2 of the USD–yen, USD–euro, USD–Swiss franc, euro–yen, and euro–Swiss franc are -0.0271, -0.0000, $+0.0033$, -0.0275, -0.0029 respectively. They are found all negative except for the euro–yen case, demonstrating that the risk reversals will subsequently adjust downward to restore the long-run equilibrium in general.[21] Using these speeds of adjustment in calculation, the GG measures in Eq. (16) are found to be closer to zero in the cases of US–Japan (0.2188) and eurozone–Switzerland (0.3869), but closer to one in the cases of US–eurozone (0.9371), US–Switzerland (0.9999), and eurozone–Japan (0.8265). These suggest that the sovereign CDS market tends to lead the currency option market in price discovery in the cases of US–Japan and eurozone–Switzerland, but lag the currency option markets in other cases.[22] This may be due to the fact that the relative sovereign CDS spreads of US–Japan and eurozone–Switzerland are both negatively correlated with their respective risk reversals during the crisis periods (i.e., the two markets experience a breakdown from the positive long-run relationship as discussed in the second hypothesis). This results in a longer process of determining the new prices in the currency option market during the periods. As a result, the GG measures support the third hypothesis that the sovereign credit risk market can move ahead of the currency option market in price discovery in some economies.

5. Conclusion

The sovereign CDS spreads and exchange rates of the developed economies including the US, Japan, Switzerland and the eurozone with the first three countries' currencies conventionally considered as safe-haven varied in a wide range when the global financial crisis emerged in late 2007 and the European sovereign debt crisis began in late 2009. This raises the question of any interconnectivity between the anticipated sovereign credit risks of these economies and the market expectations of their exchange rates during the crises. This paper uses a bivariate vector error-correction model with random coefficients to show evidence of cointegration between the prices in the currency option and sovereign CDS markets of these four economies. The estimation results find that both the relative sovereign credit risks and the exchange rate expectations are governed by a long-run relationship, while the USD–Swiss franc exchange rate is an exception statistically.

Disequilibrium from the long-run relationship triggers significant interactions between the two markets in price discovery. In particular, the relative sovereign CDS spreads of the US (the eurozone) move ahead of the risk reversals of the USD–yen (the euro–Swiss franc) in the short-run. We also find that, the price changes in the sovereign CDS and currency option markets are positively correlated in general after controlling for the non-linear cointegrating relationship and the exogenous macro-financial factors, implying that the currency option prices of these economies incorporate their sovereign credit risks. However, during the crisis periods, the positive correlations for some economies weaken noticeably, resulting in the drastic and persistent price deviations from the long-run equilibrium amid central banks' monetary measures and market turbulence.

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[21] The coefficient of the euro–yen case however is not significant.

[22] During the four-quarter period of 2009 Q3–2010 Q2, Hui and Chung (2011) find that there was a one-way information flow from the eurozone sovereign CDS market to the USD–euro currency option market. This may reflect the price dynamics of the two markets during a short period of time right after the European sovereign debt crisis that emerged in September 2009.

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