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Abstract Variation in the price of steel is an important factor to take into consideration when discussing cost control and management decisions in the construction industry. We employ various conventional and advanced econometrics methods to examine the interrelationships of steel prices in three related markets during the time period June 2002 to May 2010: Mainland China (CH), Taiwan (TW), and the United States (US). We adopt the Gregory and Hansen (GH) test and regime-switching (RS) model for cointegration, both of which accommodate endogenous structural break(s), to produce a more accurate analysis of a period in the presence of structural change(s). The empirical result of the RS cointegration test with respect to multiple structural breaks suggests a long-run equilibrium relationship among the three variables considered. This finding differs from the result of the GH test but confirms the result of the conventional Johansen test. Furthermore, the results of the Granger causality test indicate that both CH and US steel prices have great influence on the TW steel price;

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the Taiwanese steel market is closely linked with China and US steel markets in the long run.

Keywords Steel price · Cointegration · Structural change · Granger causality · Construction management

JEL Classification C22 · F15 · F36

1 Introduction

Steel price variation is a major factor to consider with respect to cost control and management decisions in the construction industry. Previous research has focused on price prediction based on statistical analysis of the time series of a single market, disregarding the possibility of external impacts and long-term relationships with related markets. Our analysis empirically investigates the effects of long-term cointegration and causal relationships, within a period of structural changes, among steel prices of Mainland China (CH), the United States (US), and Taiwan (TW).

Prior to the end of WWII, the US almost completely dominated global crude steel production and had a monopoly on price. After the 1997 Asian economic crisis, US steel prices fell to a 20-year low, due to decreased prices and falling demand for domestic products. Fortunately, however, with the rapid growth of the China market in recent years, global steel production has once again entered a state of rapid growth. This rapid development of China's steel industry and its influence on the world steel market attract an increasing amount of global attention. China's demand for imported steel and the production capacity of global crude steel continues to expand.

Despite the fluctuation of prices for various steel products in the Chinese domestic market, prices have remained quite high as a whole in recent years. In China, steel demand has continued on an upward trend concomitant with the robust growth of other steel-consuming industries, such as the automobile and construction sectors. Meanwhile, the global steel output continued, spawning a substantial increase in the volume of international steel trade with frequent record highs in price index and a phenomenal rise in raw material prices. In 2004, China's government implemented a macroeconomic control policy to temper its overheated economy. These measures included establishing stable production and orderly sales in the global steel industry to prevent short-sighted and damaging expansion.

In Taiwan, commencing in the mid-1970s, the "Ten Major Construction Projects" initiated by the government established the integrated steel plant operated by China Steel Corporation in Taiwan and the steel industry entered a phase of rapid growth. Subsequently, in the 1980s, China Steel Corporation conducted its second and third phase plant expansion projects, launching the Taiwanese steel industry into its major growth stage. The 1990s was a volatile decade for the industry as it saw cross-strait political instability and the rapid growth of Taiwanese investment in China. Taiwan officially became a WTO member on January 2, 2002, immediately following China's admittance into the WTO on December 11, 2001. As both sides gradually eased general trade and investment barriers, further cross-strait integration and specialization seemed inevitable.

The recent 2007–2008 financial market turmoil in the US incited by the subprime mortgage crisis has grown. Through 2008, the crisis caused panic in financial markets and encouraged investors to take their money out of risky mortgage bonds and shaky equities and instead put their money into commodities. Financial speculation in commodity futures following the collapse of the financial derivatives markets resulted in the world food price crisis and escalation of oil prices. Along with world inflation, steel prices jumped as well. This suggests an important structural change in world steel markets. Asia, and in particular, China, led global output trends, where surging demand set the pace of growth. In 2008, however, world crude steel production fell largely due to the slump in growth in Asia, according to the World Steel Association. Steel prices dropped due to the decreasing demand for steel in 2009.

Hence, this study investigates steel market cointegration and price transmissions of CH, US, and TW steel prices. The adoption of the Gregory and Hansen (GH) test and regime-switching (RS) model for cointegration, both of which accommodate endogenous structural breaks, produces a more accurate picture when examining a period in the presence of structural change(s). In addition, the Granger causality (GC) test is utilized to examine the dynamic interactive relationship among the three variables in consideration.

The remainder of the paper is organized as follows: Section 2 reviews some relevant literature; Sect. 3 describes the collected data; Sect. 4 deals with methods of research and presents the empirical results; and Sect. 5 concludes this paper.

2 Literature review

Conventionally, the volatility of steel prices is a function of the market. In particular, general equilibrium theory explains the price variation from the perspective of the relationship between market supply and demand. However, price variations in one market can spread easily and instantaneously to another market, especially following the economic globalization and the process of liberalization it has generated in recent years that have accelerated the integration of world markets.

Taylor and Bowen (1987) and Runeson (1988) developed a methodology for price-level forecasting in the construction industry. By including steel price as a major component of building or construction price, Akintoye and Skitmore (1994) analyzed the accuracy of three different models produced by the Building Cost Information Service (BCIS) system. These models were used to forecast the movement in macro building prices in the UK. These models, however, did not take into account external circumstances such as economic shock, e.g., a severe economic recession may decrease the accuracy of forecasts.

In later research, Fitzgerald and Akintoye (1995) used a number of quantitative methods to assess the accuracy of the construction tender price index (TPI). Wang and Mei (1998) used a time series model, the autoregressive integrated moving average (ARIMA) method to build a Taiwan Construction Cost Index (TCCI) model for the short-term forecasting of construction cost indices in Taiwan. Huang et al. (2002)

added a structural dummy variable to the existing TCCI model, which increased the forecasting accuracy of the TCCI model by accounting for the impact of external intervention. Ng et al. (2004) then used an integrated regression analysis and time series model to forecast construction TPI movements. This method entailed a judgmental projection of future market conditions, including inflation. The contributions of the aforementioned past research focused mainly on short-term forecasting for construction price movement and different methods aimed at improving forecasting accuracy; however, the methodologies limited the researchers to an estimation of some short-term dynamic processes in a single country or market and did not allow testing and explanation of the action of long-term equilibria and interactions with other related markets.

Liebman (2006) examined the factors of the US safeguard protection (steel safeguard tariffs in 2002) on steel imports contributing to the price of steel in the US. This study utilized a panel data set of monthly product-level price observations between 1997 and March 2005. Little evidence was found that the safeguards affected steel prices in the US. Instead, declining production capacity, improved macroeconomic conditions, and a falling dollar helped return prices to healthier levels. Liebman's findings further suggested that China's increasing demand for steel played an important role in the market conditions of the US steel industry. Vandenbussche and Zarnic (2008) also investigated the effects of the US safeguard protection of steel imports in 2002 on the markups of the EU steel firms. The results indicated smaller adverse effects on multi-product EU firms. The US safeguard protection also resulted in some diversion of EU steel, especially toward China, and ultimately resulted in the Chinese trade protection of steel imports

3 Data

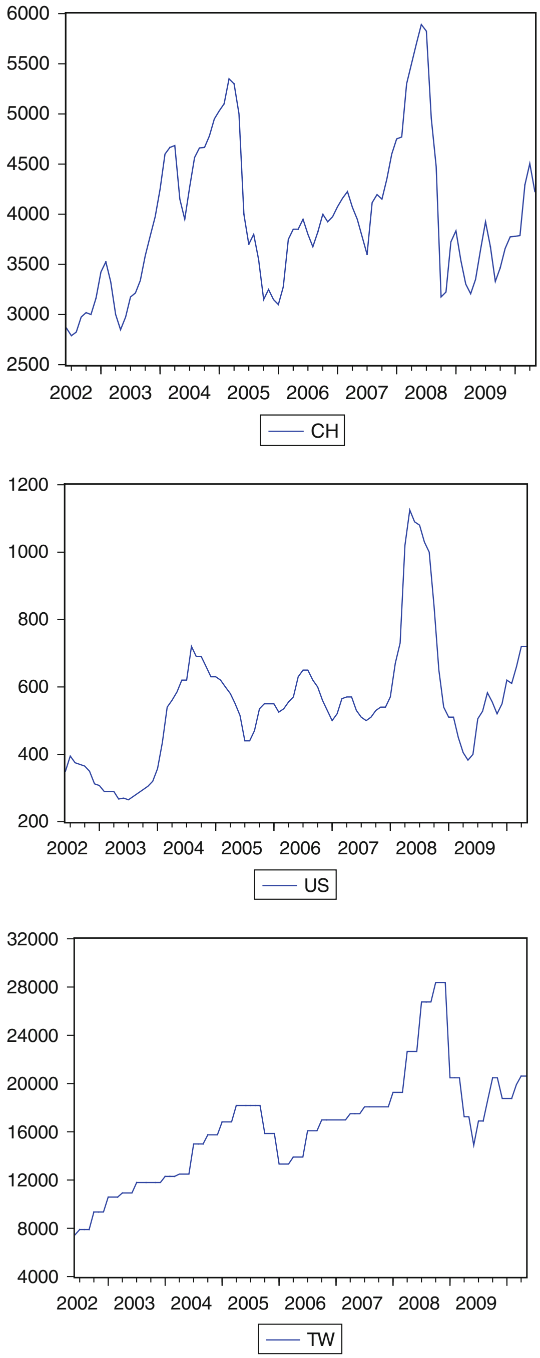
The sampling period of this study runs from June 2002 to May 2010; a total of 96 monthly observations are obtained for each variable. The source of data for CH and US steel prices is the Bloomberg website,[1,2] while the source of data for TW steel prices is the Cmoney data bank in Taiwan.[3] All prices are based on hot-rolled coil steel prices. Figure 1 illustrates the trends of CH, TW, and US steel prices. Trends from June 2002 to May 2010 are up-and-down cycles. It is clear that in 2008 the price of steel in the US experienced a dramatic increase from April to September. The price movement in CH steel prices exhibited an apparent rise as well. During this period, however, the fluctuation of steel price in Taiwan was only moderate. It appears that there exists a possible structural break in the trends of CH, US, or TW steel prices around the year 2008.

[1] http://www.bloomberg.com

[2] The code for CH is SBBSCDHR Index and for US is SBBSNAHR Index. Both price indices are compiled by Steel Business Briefing (SBB). It is a journalistic assessment of the range of prices for this product category market. The price assessment is made following off-the-record telephone conversations with industry participants such as steel mills, traders, processors, and end-users.

[3] http://www.cmoney.com.tw

Fig. 1 The steel price movements of CH, US, and TW



4 Methodologies and empirical results

In this study, various econometric methods are used to examine the long-term cointegration and dynamic short-term relationships of steel prices in the three related markets of CH, TW, and US spanning the period from June 2002 to May 2010. Also considered is the impact of structural changes (breaks) on the econometric models as failure to detect and account for parameter changes may lead to erroneous inferences and poor forecasting performances (Clements and Hendry 1999). This is especially relevant for cointegration analysis, since it normally involves data spanning long periods of time.

4.1 Unit root tests

4.1.1 Conventional linear unit root tests

This paper first tests for "stationarity" of each variable by employing three traditional unit root test techniques: ADF (Dickey and Fuller 1981), PP (Phillips and Perron 1988), and NP (Ng and Perron 2001).[4]

According to Schwert (1989), the ADF test with long lags is superior to the others. The three differencing AR models of ADF are expressed as follows:

$$\Delta y_t = \phi\, y_{t-1} + \sum_{i=1}^{p-1} \beta_i \Delta y_{t-i} + \varepsilon_t \tag{1}$$

$$\Delta y_t = \alpha + \phi\, y_{t-1} + \sum_{i=1}^{p-1} \beta_i \Delta y_{t-i} + \varepsilon_t \tag{2}$$

$$\Delta y_t = \alpha + \phi\, y_{t-1} + \gamma + \sum_{i=1}^{p-1} \beta_i \Delta y_{t-i} + \varepsilon_t \tag{3}$$

Model (1) is a pure random walk with lag terms. Model (2) possesses a drift. Model (3) has a drift and a time trend.

Hamilton (1994) argues that a strategy is necessary to determine which of the three ADF models should be employed in conducting the unit root test. In this paper, we follow the determining rule by Dolado, Jenkinson, and Sosvilla-Rivero (DJS) (1990) to ascertain the appropriate model for each variable.[5]

Phillips and Perron (1988) proposed an alternative (non-parametric) method of controlling for serial correlation when testing for a unit root. In subsequent research, Ng and Perron (2001) further constructed four tests statistics (MZa, MZt, MSB, and MPT), known as the NP unit root test.

[4] The test statistic of NP test is MZ_t in this paper.

[5] The determining rule by DJS (1990) is to test for the significance of trend coefficient in the third model first, followed by testing for the significance of the drift coefficient in the second model. If both outcomes result in insignificant, the first model is selected.

Table 1 Summary statistics for CH, TW, and US steel prices

	CH	TW	US
Mean	3960.599	16396.02	546.8490
Maximum	5890.000	28369.00	1125.000
Minimum	2790.000	7461.000	265.0000
Std. dev.	729.5005	4598.323	183.2209
Skewness	0.618198	0.452649	1.033935
Kurtosis	2.851117	3.415984	4.789486
Jarque-Bera	6.203362	3.970420	29.91339
Probability	0.044974	0.137352	0.000000
Observations	96	96	96

CH, TW, and US are the symbols for Mainland China, Taiwan, and USA raw steel prices, respectively
The unit price for CH is China domestic hot-rolled coil RMB/ metric ton; for US is north America domestic hot-rolled coil (FOB US Midwest mill) USD/short ton; for TW is Taiwan domestic hot-rolled coil NTD/ metric ton

Based on the "principle of parsimony," our unit root tests utilize modified Akaike information criterion (MAIC) (suggested by Ng and Perron (2001) for ADF and NP, and the Bartlett kernel-based criterion proposed by Newey and West (1994) for PP to, respectively, determine the optimal number of lags and the optimal bandwidth.

The summary statistics are presented in Table 1. And the results of various unit root tests (ADF, PP, and NP) are summarized in Table 2. They show that the null of non-stationarity cannot be rejected at any level (except for the ADF test for the US at a very low 10% significance level). After first differencing, however, the null is rejected at the 1% significance level for all series. We thus conclude that all the variables considered in this paper are I(1) series.

4.1.2 Advanced non-linear unit root test with a structural break—ZA test

One major drawback of conventional unit root tests is the implicit assumption that the deterministic trend is correctly specified. Perron (1989) developed a unit root model with an exogenous structural break, and argued that if there is a break in the deterministic trend, the unit root test may lead to the misleading conclusion that there is a unit root when in fact there is not. Work by Zivot and Andrews (1992) threw out the assumed exogenous break point resulting in a unit root test with an endogenous structural break. This test has been regarded as more suitable for the stationarity of series. The three models of ZA tests are expressed as follows:

$$\text{Model A}:\ \Delta Y_t = \mu_1^{\mathrm{A}} + \gamma_1^{\mathrm{A}} t + \mu_2^{\mathrm{A}} DU_t(\lambda) + \alpha^{\mathrm{A}} Y_{t-1} + \sum_{j=1}^{k-1} \beta_j \Delta Y_{t-j} + \varepsilon_t \quad (4)$$

$$\text{Model B}:\ \Delta Y_t = \mu_1^{\mathrm{B}} + \gamma_1^{\mathrm{B}} t + \gamma_2^{\mathrm{A}} DT_t^{*}(\lambda) + \alpha^{\mathrm{B}} Y_{t-1} + \sum_{j=1}^{k-1} \beta_j \Delta Y_{t-j} + \varepsilon_t \quad (5)$$

Table 2 Results of various unit root tests

		CH	TW	US
ADF	Level	−2.882 [2] (2)	−1.806 [2] (6)	−3.251* [3] (1)
	First difference	−6.387*** [1] (0)	−4.006*** [1] (2)	−5.370*** [1] (0)
PP	Level	−2.540 [2] (1)	−2.057 [2] (5)	−2.334 [2] (5)
	First difference	−6.089*** [1] (7)	−9.792*** [1] (5)	−5.362*** [1] (1)
NP	Level	−2.160 [3] (1)	−2.484 [3] (5)	−2.517 [3] (3)
	First difference	−3.707*** [3] (8)	−5.501*** [3] (5)	−3.981*** [3] (1)

1. CH, TW, and US are the symbols for Mainland China, Taiwan, and USA raw steel prices, respectively
2. ***, ** , and * denote significant at 1, 5, and 10% levels, respectively
3. The critical values for 1, 5, and 10% levels of ADF, PP and NP, are (−4.059; −3.458; −3.155), (−3.501; −2.892; −2.583) and (MZt: −3.42; −2.91; −2.62), respectively. The critical values for the ADF t statistics are from the Mackinnon (1996) table
4. The test statistic of the NP test is MZ_t
5. The numbers in the brackets indicate model specification. The numbers in the parentheses of ADF and NP are the appropriate lag lengths selected by MAIC, whereas the numbers in the parentheses of PP is the optimal bandwidths decided by the Bartlett kernel of Newey and West (1994)
6. The appropriate models selected for the level and the first difference are based on the decision procedures suggested by Dolado et al. (1990)

$$\text{Model C}: \Delta Y_t = \mu_1^{\mathrm{C}} + \gamma_1^{\mathrm{C}} t + \mu_2^{\mathrm{C}} DU_t(\lambda) + \gamma_2^{\mathrm{C}} DT_t^*(\lambda) + \alpha^{\mathrm{C}} Y_{t-1} + \sum_{j=1}^{k-1} \beta_j \Delta Y_{t-j} + \varepsilon_t \quad (6)$$

where $DU_t(\lambda)$ is 1 and $DT_t^*(\lambda) = t - T\lambda$ if $t > T\lambda$, 0 otherwise. $\lambda = T_{\mathrm{B}}/T$ and T_{B} represents a possible break point. Model A allows for a change in the level of the series while Model B allows for a change in the slope of trend of a series and Model C combines both changes in the level and slope of trend. Since the appropriate model and the optimal lag lengths are crucial in interpreting the results of the tests, we adopt the findings from the ADF tests to select the model and the lag lengths for the ZA tests.

The results of the ZA tests are presented in Table 3. All three variables carry unit root in the significance level and reject the null of "non-stationarity" in the first difference. This is consistent with and insures the I(1)-type series for the three variables considered. The same order of integration among non-stationary series found from both linear and non-linear tests enables us to go further with cointegration tests.

4.2 Cointegration test

4.2.1 Johansen's cointegration test

Various methods of estimating cointegration have been applied to capture the long-run equilibrium relationship among the variables. Of these, Johansen's methodology, based on the likelihood ratio with non-standard asymptotic distributions involving

Table 3 Results of Zivot and Andrews unit root test with a structural break

	Break	Level ($t(\hat{\lambda}_{\inf})$)	Break	First difference ($t(\hat{\lambda}_{\inf})$)
CH	4/2005	−4.121724	7/2008	−4.938685*
US	1/2008	−4.465961	8/2008	−5.194436**
TW	3/2008	−4.347958	12/2008	−6.499189***

1. The symbols *** and ** represent the significant at 1 and 5% levels, respectively
2. The critical value for (1, 5, and 10%) levels are (−5.57, −5.08, and −4.82) for Model C from Zivot and Andrews (1992)
3. The characters within the parenthesis indicate the appropriate model according to the results from ADF test

integrals of Brownian motions, is found to be the best method to proceed with cointegration estimation by Gonzalo (1994).[6]

The elaborate work developed by Johansen (1988, 1994) is distilled into five VAR models with ECM, which are presented as follows:[7]

$$\begin{aligned}
\mathrm{H}_0(r): \Delta X_t &= \Gamma_1 \Delta X_{t-1} + \cdots + \Gamma_{k-1}\Delta X_{t-(k-1)} \\
&\quad + \alpha\beta' X_{t-1} + \Psi D_t + \in_t \qquad (1988) && (7)\\
\mathrm{H}_1^*(r): \Delta X_t &= \Gamma_1 \Delta X_{t-1} + \cdots + \Gamma_{k-1}\Delta X_{t-(k-1)} + \alpha(\beta', \beta_0)(X'_{t-1}, 1)' \\
&\quad + \Psi D_t + \in_t \qquad (1990) && (8)\\
\mathrm{H}_1(r): \Delta X_t &= \Gamma_1 \Delta X_{t-1} + \cdots + \Gamma_{k-1}\Delta X_{t-(k-1)} + \alpha\beta' X_{t-1} + \mu_0 + \Psi D_t \\
&\quad + \in_t \qquad (1990) && (9)\\
\mathrm{H}_2^*(r): \Delta X_t &= \Gamma_1 \Delta X_{t-1} + \cdots + \Gamma_{k-1}\Delta X_{t-(k-1)} + \alpha(\beta', \beta_1)(X'_{t-1}, t)'\mu_0 \\
&\quad + \Psi D_t + \in_t \qquad (1994) && (10)\\
\mathrm{H}_2(r): \Delta X_t &= \Gamma_1 \Delta X_{t-1} + \cdots + \Gamma_{k-1}\Delta X_{t-(k-1)} + \alpha\beta' X_{t-1} + \mu_0 \\
&\quad + \mu_1 t + \Psi D_t + \in_t \qquad (1994) && (11)
\end{aligned}$$

To analyze the deterministic term, Johansen decomposed the parameters μ_0 and μ_1 in the directions of α and $\alpha_\perp$ as $\mu_i = \alpha\beta_i + \alpha_\perp\gamma_i$. Thus, we have $\beta_i = (\alpha'\alpha)^{-1}\alpha'\mu_i$ and $\gamma_i = (\alpha_\perp{}'\alpha_\perp)^{-1}\alpha_\perp{}'\mu_i$. The nested sub-models of the general model of null hypothesis $\Pi = \alpha\beta'$ are, therefore, defined as follows:

$$\begin{aligned}
&\mathrm{H}_0\,(r): Y = 0\\
&\mathrm{H}_1^*\,(r): Y = \alpha\beta_0\\
&\mathrm{H}_1\,(r): Y = \alpha\beta_0 + \alpha_\perp\gamma_0\\
&\mathrm{H}_2^*\,(r): Y = \alpha\beta_0 + \alpha_\perp\gamma_0 + \alpha\beta_1 t\\
&\mathrm{H}_2\,(r): Y = \alpha\beta_0 + \alpha_\perp\gamma_0 + (\alpha\beta_1 + \alpha_\perp\gamma_1)t
\end{aligned}$$

[6] Gonzalo (1994) compared several methods of estimating cointegration, which include ordinary least squares, non-linear least squares, maximum likelihood in an error correction model, principle components, and canonical correlations.

[7] The 1990 equation form is from Johansen and Juselius (1990).

Table 4 Determination of cointegration rank in consideration of a linear trend and a quadratic trend

Rank	Model 1 (H_0)		Model 2 (H_1^*)		Model 3 (H_1)		Model 4 (H_2^*)		Model 5 (H_2)	
	$T_0(r)$	$C_0(5\%)$	$T_1^*(r)$	$C_1^*(5\%)$	$T_1(r)$	$C_1(5\%)$	$T_2^*(r)$	$C_2^*(5\%)$	$T_2(r)$	$C_2(5\%)$
$r = 0$	37.62	24.28	48.97	35.19	47.68	29.80	67.38	42.92	66.24	35.01
$r \leq 1$	8.44	12.32	19.12	20.26	18.41	15.49	25.60	25.87	25.03	18.40
$r \leq 2$	0.00	4.13	7.00	9.16	6.29	3.84	10.24	12.52	9.69	3.84

1. $T_0(r)$, $T_1^*(r)$, $T_1(r)$, $T_2^*(r)$, and $T_2(r)$ denote the LR test statistics for all the nulls of $H(r)$ versus the alternative of $H(p)$ of Johansen's five models
2. $C_0(5\%)$, $C_1^*(5\%)$, $C_1(5\%)$, $C_2^*(5\%)$, and $C_2(5\%)$ are the 5% LR critical values for Johansen's five models, which are extracted from Osterwald-Lenum (1992)
3. The model selection follows Nieh and Lee's (2001) decision procedure, diagnosing models one by one until the model that cannot be rejected for the null
4. VAR length is 2 for all the models, which is selected based on the smallest number of AIC

Johansen (1994) emphasized the role of the deterministic term, $Y = \mu_0 + \mu_1 t$, which includes constant and linear terms in the Gaussian VAR. A decision procedure, like that of Nieh and Lee (2001), for the hypotheses $H(r)$ and $H^*(r)$ of the above five different models is presented as follows:[8]

$$
\begin{aligned}
&H_0\,(0) \rightarrow H_1^*\,(0) \rightarrow H_1\,(0) \rightarrow H_2^*\,(0) \rightarrow H_2\,(0) \rightarrow H_0\,(1) \rightarrow H_1^*\,(1) \rightarrow \\
&H_1\,(1) \rightarrow H_2^*\,(1) \rightarrow H_2\,(1) \rightarrow \cdots \rightarrow \cdots \rightarrow H_0\,(p-1) \rightarrow H_1^*\,(p-1) \rightarrow \\
&H_1\,(p-1) \rightarrow H_2^*\,(p-1) \rightarrow H_2\,(p-1)
\end{aligned}
$$

In this procedure, we evaluate models one by one until we reach the model whose null hypothesis cannot be rejected. The empirical findings for the long-run relationship among CH, TW, and US steel prices are modeled by a linear trend and a quadratic trend and are summarized in Table 4. VAR length is 2 for all the models, which was selected based on the smallest number of AIC. The empirical results show that one cointegrating vector exists among these three steel prices. Based on the decision procedure described above, the result is Johansen's first model with neither linear nor quadratic trend.

4.2.2 Advanced non-linear cointegration test with a structural break—GH test

The GH test is considered as a multivariate extension of the univariate ZA unit root test. The GH test is a residual-based test, which tests the null of no cointegration against the alternative of cointegration in the presence of a possible regime shift.[9] This shift

[8] Johansen (1992, 1994) developed a testing procedure based on the ideas developed by Pantula (1989) to determine the number of cointegrating rank in the presence of linear trend (Johansen 1992) and quadratic trend (Johansen 1994).

[9] The residuals are tested to determine if they are I(0) or I(1) processes. Rejection of the null implies an I(0) residual and, therefore, indicates the presence of cointegration.

Table 5 GH cointegration test with a structural break

	Break	ADF*	Break	Z_t^*	Break	Z_α^*
C	7/2005	−4.3671	7/2005	−3.5266	6/2005	−22.6556
C/T	7/2005	−4.4408	8/2005	−3.6348	8/2005	−24.2917
C/S	6/2008	−4.4699	4/2008	−4.0027	4/2008	−28.5019

Note Critical values for model C, C/T, and C/S are −4.92, −5.29, and −5.50 for ADF* and Z_t^*, and −46.98, −53.92, −58.33 for Z_α^*, respectively (see Table 1 in Gregory and Hansen 1996, p. 109)

is a single break in the intercept and/or slope coefficients vector. Four models of GH test are proposed in Gregory and Hansen (1996). Model 1 is the standard Engle and Granger (1987) two-step approach. The other three models are expressed as follows:

$$\text{Model 2 (C)}: y_t = \mu_1 + \mu_2 D_t(\lambda) + \beta_1 X_t + e_t \tag{12}$$

$$\text{Model 3 (C/T)}: y_t = \mu_1 + \mu_2 D_t(\lambda) + \gamma t + \beta_1 X_t + e_t \tag{13}$$

$$\text{Model 4 (C/S)}: y_t = \mu_1 + \mu_2 D_t(\lambda) + \beta_1 X_t + \beta_2 X_t D_t(\lambda) + e_t \tag{14}$$

where y_t is an univariate variable, and X_t is a variate vector. If $t > T\lambda(t = 1, \ldots, T)$, $D_t(\lambda)$ equals 1, otherwise 0. $\lambda = T_B/T$ and T_B represents a possible break point.

The residuals are then tested to determine whether they are I(0) or I(1) processes. If the residuals are I(0) processes, this implies the presence of cointegration. The results of the GH test are shown in Table 5. As these figures are not statistically significant, no cointegration is found among the three variables in the presence of a possible structural break or regime shift.

4.2.3 RS cointegration test

The RS test, proposed by Gabriel et al. (2002), further extends previous GH model to enable testing over periods of time with multiple structural breaks.

The RS model allows for RS in the intercept, regression parameters, and volatility, which are given as follows:

$$y_t = \beta(s_t)'x + \theta(s_t)u_t \tag{15}$$

$$\hat{u}_t = (\hat{\pi}_1\hat{\theta}_1^2 + \hat{\pi}_2\hat{\theta}_2^2)^{-0.5}(y_t - \hat{\pi}_1\hat{\beta}_1 x_t - \hat{\pi}_2\hat{\beta}_2 x_t) \tag{16}$$

where $\hat{\pi}_i$ is the smoothed transition probability.[10]

[10] See Appendix B in Davies (2006).

Both Davies (2006) and Lucey and Voronkova (2008) assume $\theta(s_t) = 1$ in the model of Gabriel et al. (2002). As a result, their models allow for RS only in the intercept and regression parameters.

Davies (2006) specifies the model as the following:

$$y_t = \beta'_{s_t} x_t + u_t \tag{17}$$

$$\hat{u}_t = y_t - \hat{\pi}_1 \hat{\beta}_1 x_t - \hat{\pi}_2 \hat{\beta}_2 x_t = \hat{\pi}_1 e_1 + \hat{\pi}_2 e_2 \text{ where } e_{s_t} = y_t - \hat{\beta}_{s_t} x_t \tag{18}$$

On the other hand, the model of Lucey and Voronkova (2008) model is represented by the following equations:

$$y_t = \alpha_{s_t} + \textstyle\sum_{i=1}^{n} \beta_{i,s_t} x_{i,t} + u_t \tag{19}$$

$$\hat{u}_t = y_t - \hat{\pi}_1 \hat{\beta}_1 x_t - \hat{\pi}_2 \hat{\beta}_2 x_t = \hat{\pi}_1 e_1 + \hat{\pi}_2 e_2 \text{ where } \beta_1 = (\alpha_1 \ \beta_{1,1} \ \ldots \ \beta_{n,1}) \tag{20}$$

Gabriel et al. (2002) argued that the residual-based two-step approach by Engle and Granger (1987) can be employed in the cointegration test with multiple structural breaks. In this paper, we apply three methods, ADF, PP, and NP unit root tests on $\hat{u}_t$ for detecting the long-run relationship between x and y of Eq. (20).[11]

There are several findings from the RS cointegration test which takes multiple structural breaks into consideration. First, as illustrated in Table 6, the rejection of a unit root from all three approaches suggests that there exists a long-run equilibrium relationship among the three variables considered. This finding differs from the result of the GH test but confirms the previous result of the conventional Johansen test. Second, as seen in Table 7 (the estimated value of parameters), all the parameters are significant. Third, when we investigate the regime characteristics of the Markov switching model, the transition matrix of Table 7 illustrates that there is a phenomenon of regime clustering. The first regime (regime 1) can be considered to be the high steel price period, while the second regime (regime 2) can be looked upon as the regular steel price period. For instance, when steel price is quite high as in regime 1 for the first month, the probability that the price will continue to be high in the next month is approximately 97%. Furthermore, the regime probability shown in Table 7 indicates that the probabilities that the price of steel will remain at a higher price in regime 1 and remain at a regular price in regime 2 are 71 and 29%, respectively. The average duration of regime 1 and regime 2 are approximately 33 and 13.5 months, respectively.[12]

Finally, from our results of the RS cointegration test, the smoothed probability based on regime 1 shows that there are four structural breaks that occurred within our sample period. These breaks occurred in October 2003, June 2005, October 2007, and September 2008 (Fig. 2).

11 The appropriate model for each variable again follows the determining rule by Dolado et al. (1990).

12 The periods of regime 1 are from 2002:6 to 2003:10, 2005:6 to 2007:10, and 2008:9 to 2010:5. On the contrary, the periods of regime 2 are from 2003:11 to 2005:5 and 2007:11 to 2008:8.

Table 6 RS cointegration tests with multiple structural breaks

ADF	PP	NP
−4.235841*** [1] (0)	−4.111845*** [1] (5)	−24.334209*** [1] (0)

1. ***, **, and * denote significant at 1, 5, and 10% levels, respectively
2. The critical values for 1, 5, and 10% levels of ADF, PP, and NP, are (−4.044; −3.452; −3.151), (−4.037; −3.448; −3.149), and (MZ_a: −13.800; −8.100; −5.700), respectively. The critical values for the ADF statistics are from the Mackinnon (1996) table
3. The test statistic of the NP test is MZ_a
4. The numbers in the brackets indicate model specification. The numbers in the parentheses of ADF, and NP are the appropriate lag lengths selected by MAIC, whereas the numbers in the parentheses of PP is the optimal bandwidths decided by the Bartlett kernel of Newey and West (1994)
5. The appropriate models selected for the level and the first difference are based on the decision procedures suggested by Dolado et al. (1990)

Table 7 Estimated results and regime characteristics of markov switching model

Parameter estimation	α		β	
Regime 1	5.7930***	(0.3560)	0.2477***	(0.0369)
Regime 2	4.8081***	(0.8126)	0.3773***	(0.0838)
Regime characteristics	Transition matrix		Regime probability	Average duration
	Regime 1	Regime 2		
Regime 1	0.9697	0.0303	0.7103	33.0033
Regime 2	0.0743	0.9257	0.2897	13.4590

1. Numbers within the parentheses are standard deviations
2. *** denotes significant at 1% level

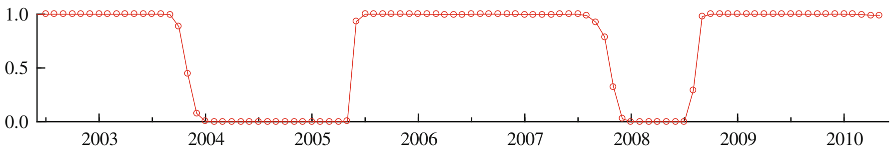


Fig. 2 Smoothed probability of RS

4.3 GC test

If all variables are non-stationary and cointegrated, the appropriate method to examine the causal relations is the vector error correction model (VECM) (Granger 1988).[13] Otherwise a vector autoregression (VAR) model is used in the case that no cointegration relationship is found between the variables (Granger 1969).

[13] This is due to the reason that the Engle and Granger (1987) "Granger representation theorem" argued that error correction and cointegration are equivalent representations.

Table 8 Results of pairwise GC test

Null hypothesis	F statistic	Probability
TW does not Granger Cause CH	0.03803	0.9627
CH does not Granger Cause TW	8.71341***	0.0004
US does not Granger Cause CH	0.11968	0.8874
CH does not Granger Cause US	12.7432***	0.0000
US does not Granger Cause TW	14.1049***	0.0000
TW does not Granger Cause US	0.60616	0.5477

1.CH, TW, and US are the symbols for Mainland China, Taiwan and USA raw steel prices, respectively
2. ***, **, and * denote significant at 1, 5, and 10% levels, respectively
3. The null hypothesis, H_0, is for "no causal relation"
4. Lag length is 2 selected by AIC

The standard GC test then examines whether there exists feedback (bi-directional) or one-way causality between variables.[14] Considering two series, X_t and Y_t, the GC test is as follows:

$$\Delta X_t = \alpha_1 + \sum_{i=1}^{n_1} \alpha_{11}(i)\, \Delta X_{t-i} + \sum_{j=1}^{m_1} \alpha_{12}(j)\, \Delta Y_{t-j} + \varepsilon_{Xt} \tag{21}$$

$$\Delta Y_t = \alpha_2 + \sum_{i=1}^{n_2} \alpha_{21}(i)\, \Delta X_{t-i} + \sum_{j=1}^{m_2} \alpha_{22}(j)\, \Delta Y_{t-j} + \varepsilon_{Yt} \tag{22}$$

where ε_{X_t} and ε_{Y_t} are stationary random processes intended to capture other pertinent information not contained in lagged values of X_t and Y_t. The lag lengths, n and m, are determined by AIC in our study. The series Y_t fails to cause X_t if $\alpha_{12}(j) = 0 (j = 1, 2, \ldots, m_1)$ and the series X_t fails to cause Y_t if $\alpha_{21}(i) = 0 (i = 1, 2, \ldots, n_2)$.

Since the results from the former Johansen and GH tests indicate a cointegration relationship existing among the three steel prices, we apply VECM (Granger 1988) to test the causal relationships between each set of paired variables. The results of the GC test, shown in Table 6, indicate that one-way causality between each pair of variables reveals that CH steel prices affects TW and US and US steel prices affects TW. However, there is no significant lead–lag (causal) relationship found in the reverse order. This indicates that both CH and US steel prices have great influence on the TW steel price. TW is obviously an exogenous variable to the remaining variables (Table 8).

[14] The pairwise GC test is actually employed to examine the "precedence," not causality, between two variables.

5 Conclusions

The variation in the price of steel is an important factor to take into account when making cost control and management decisions in the construction industry; hence, various conventional and advanced econometric methods were used to examine the interrelationships of steel prices in the three related markets of CH, TW, and the US during the period June 2002 to May 2010.

Several conclusions can be drawn from our empirical tests. First, we find that all variables, from both linear (conventional ADF, PP, and NP) and non-linear (ZA) tests, are insured I(1) non-stationary time series. Second, Johansen's maximum likelihood cointegration test shows that there exists a long-term equilibrium relationship among CH, TW, and US steel prices. Third, the findings of the RS cointegration test that accounts for multiple structural breaks suggest that there exists a long-run equilibrium relationship among the three variables considered. This finding differs from the result of the GH tests but confirms the previous result of the conventional Johansen test. The four structural breaks occur in October 2003, June 2005, October 2007, and September 2008. Furthermore, there seems to be a phenomenon of regime clustering. When steel prices are quite high as in the first regime for the first month, the probability that the price will remain at that high level in the next month is approximately 97%. Fourth, the results of the GC test reveal that CH steel prices affect TW and US and US steel prices affect TW. This implies that CH affects US first and then TW. This is consistent with Liebman (2006) finding that China's increasing demand for steel played an important role in the market conditions of the US steel industry (China is the number one exporter to the US). China's dominant role in the global steel market in recent years played an important role as well.

The evidence from the GC tests indicate that the Taiwanese steel market is closely linked with that of the US and China in the long run. The variations in market steel prices of China and the US can substantially affect Tawianese steel prices. Thus, construction firms and material suppliers in Taiwan should establish a sound long-term plan and close bilateral trade relationship with China to further cross-strait integration and cooperation. On the other hand, Taiwan does not have too much bargaining power with respect to China or the US since the US has also acted as a major steel material exporter to Taiwan over a long period of time.

As steel trade between the two sides of the Taiwan Strait has been brisk in recent years, steel-related industries such as the automobile and construction industries are expected to benefit while steel demand in China remains in the high growth stage. As prices fluctuate, the orders from construction firms in Taiwan can be placed in advance to ensure an adequate supply of steel products required for construction. The establishment of coordination mechanisms for cross-strait industry trade is therefore very important.

References

Akintoye SA, Skitmore RM (1994) A comparative analysis of the three macro price forecasting models. Constr Manage Econ 12(3):257–270

Clements MP, Hendry DF (1999) On Winning Forecasting Competitions in Economics. Span Econ Rev 1:123–160
Davies A (2006) Testing for international equity market integration using regime switching cointegration techniques. Rev Financ Econ 15(4):305–321
Dickey DA, Fuller WA (1981) Likelihood ratio statistics for autoregressive time series with a unit root. Econometrica 49:1057–1072
Dolado JJ, Jenkinson T, Sosvilla-Rivero S (1990) Cointegration and unit roots. J Econ Surv 4:249–273
Engle R, Granger C (1987) Co-integration and error correction representation, estimation and testing. Econometrica 55:251–267
Fitzgerald E, Akintoye A (1995) The accuracy and optimal linear correction of UK construction tender price index forecasts. Constr Manage Econ 13(6):493–500
Gabriel VJ, Psaradakis Z, Sola M (2002) A simple method of testing for cointegration subject to multiple regime changes. Econ Lett 76(2):213–221
Gonzalo J (1994) Five alternative methods of estimating long-run equilibrium relationships. J Economet 60:203–233
Granger CWJ (1969) Investigating causal relations by econometric models and cross-spectral methods. Econometrica 37:24–36
Granger CWJ (1988) Some recent developments in a concept of causality. J Economet 39:199–211
Gregory A, Hansen B (1996) Residual-based tests for cointegration in models with regime shifts. J Economet 70:99–126
Hamilton JD (1994) Time series analysis. Princeton University Press, Princeton
Huang YC, Yang HA, Wang CH (2002) Increasing forecast accuracy of Taiwan construction cost indices by structural dummy variable. J Chin Inst Civ Hydraul Eng 14(4):651–658
Johansen S (1988) Statistical analysis of cointegration vectors. J Econ Dyn Control 12:231–254
Johansen S (1992) Determination of cointegration rank in the presence of a linear trend. Oxf Bull Econ Stat 54(3):383–397
Johansen S (1994) The role of the constant and linear terms in cointegration analysis of nonstationary variables. Economet Rev 13(2):205–229
Johansen S, Juselius K (1990) Maximum likelihood estimation and inference on cointegration with applications to the demand for money. Oxf Bull Econ Stati 52(2):169–210
Liebman BH (2006) Safeguards, China, and the price of steel. Rev World Econ 142(2):354–373
Lucey BM, Svitlana V (2008) Russian equity market linkages before and after the 1998 crisis: Evidence from stochastic and regime-switching cointegration tests. J Int Money Finan 27:1303–1324
Mackinnon JG (1996) Numerical distribution functions for unit root and cointegration tests. J Appl Econ 11:601–618
Newey W, West K (1994) Automatic lag selection in covariance matrix estimation. Rev Econ Stud 61: 631–653
Ng S, Perron P (2001) Lag length selection and the construction of unit root tests with goof size and power. Econometrica 69:1519–1554
Ng ST, Cheung SO, Skitmore M, Wong TCY (2004) An integrated regression analysis and time series model for construction tender price index forecasting. Constr Manage Econ 22(5):483–493
Nieh CC, Lee CF (2001) Dynamic relationship between stock prices and exchange rates for G-7 countries. Q Rev Econ Finan 41:477–490
Osterwald-Lenum M (1992) Practitioner's corner—a note with quantiles of the asymptotic distribution of the maximum likelihood cointegration rank test statistics. Oxf Bull Econ Stat 54:461–472
Pantula SG (1989) Testing for unit roots in time series data. Econometric Theory 5:256–271
Perron P (1989) The great crash, the oil price shock and the unit root hypothesis. Econometrica 57: 1361–1401
Phillips PCB, Perron P (1988) Testing for a unit root in time series regression. Biometrika 75:335–346
Runeson KG (1988) Methodology and method for price-level forecasting in the building industry. Constr Manage Econ 6(1):49–55
Schwert GW (1989) Tests for unit roots: a Monte Carlo investigation. J Bus Econ Stat 7:147–159
Taylor RG, Bowen PA (1987) Building price-level forecasting: an examination of techniques and application. Constr Manage Econ 5(1):21–44
Vandenbussche H, Zarnic Z (2008) US safeguards on steel and the markups of European producers. Rev World Econ 144(3):458–490

Wang CH, Mei YH (1998) Model for forecasting construction cost indices in Taiwan. Constr Manage Econ 16:147–157

Zivot E, Andrews DWK (1992) Further evidence on the great crash, the oil price shock, and the unit root hypothesis. J Bus Econ Stat 10:251–270