Price dynamics in the import wooden bed market of the United States

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Abstract

Imports of wooden beds by the United States have been over one billion dollars since 2005. With a combined market share of more than 60%, China and Vietnam have dominated the import market in recent years. This trade surge has become a serious threat to the domestic furniture industry. Threshold cointegration and an asymmetric error correction model are used to analyze the price dynamics between China and Vietnam. China has been the price leader of the market and its price has been evolving more independently. The transmission between the prices of China and Vietnam has been asymmetric in both the long term and short term. Furniture firms in China and Vietnam adjust prices about two times faster when the price margin is squeezed than stretched.

Keywords: Asymmetric price transmission; China; Furniture; Market integration; Threshold cointegration

1. Introduction

The United States has experienced a rapid growth in the consumption of furniture in recent decades. Total domestic furniture retail sales have been steadily rising in recent years, exceeding $100 billion in 2003 (Gazo and Quesada, 2005). However, an increasing share of the furniture demand has been met by imports from foreign countries. Traditionally, the United States has imported furniture from Canada, Italy and Taiwan. As global trade liberalization accelerates, the furniture market has changed considerably over the last decade. Furniture from China and Vietnam has substituted for that from traditional suppliers and has begun to dominate the US import market in recent years. In particular, the International Trade Commission (ITC) reports that in the import market of wooden beds, China has been the largest supplier and accounted for a 42% share of the total imports over 2002 – 2009 (U.S. ITC, 2010). Consequently, this import surge has become a serious threat to the domestic furniture manufacturing industry and aroused wide concerns in the United States.

The rising liberalization of global markets has been accompanied by an increasing number of studies on market and price relationships. Price transmissions, market integration and price leadership have been often examined to help understand price interaction and dynamics. Particularly, Asymmetric Price Transmission (APT) has received considerable attention in the literature because price transmission may differ according to whether prices are increasing or decreasing. Depending on the issue and study purpose, APT has been classified and analyzed in several ways. One typical classification is positive versus negative APT. If one price (e.g., price of bread) reacts more fully or rapidly to an increase in another price (e.g., price of wheat) than to a decrease, then the price transmission is referred to as positive asymmetry (Meyer and von Cramon-Taubadel, 2004). More generally, with positive APT, price movement that squeezes the margin is transmitted more rapidly or completely than the equivalent movement that stretches the margin. Conversely, APT is negative when price movements that stretch the margin are transmitted more rapidly or completely than movements that squeeze it. In addition, APT also can be classified as vertical or spatial. A typical example of vertical APT is that consumers often feel increases in farm prices are more fully and rapidly transmitted to retail levels than equivalent decreases

(Kinnucan and Forker, 1987). As an example of spatial APT, a rise of the US export price for wheat can cause a more pronounced reaction in the Canadian export price than a corresponding reduction of the same magnitude (Ghoshray, 2007).

Various sources of APT have been discussed in the literature (Frey and Manera, 2007). Most publications on APT refer to noncompetitive market structures as an explanation for asymmetry, and furthermore, many studies conclude that market power leads to positive APT. Another major explanation for APT is adjustment or menu costs that arise when firms change quantities and prices of inputs or outputs. If these costs are asymmetric with respect to increases or decreases in quantities and prices, APT becomes the natural result. A related cause of spatial APT often cited is the asymmetric flow of information between central and peripheral markets (Abdulai, 2000). Prices at a central market, by virtue of its size and the fact that it is at the center or a network of information, may tend to be less responsive to price changes in individual peripheral market than vice versa. Other causes of APT include political intervention and inventory management (Meyer and von Cramon-Taubadel, 2004).

A number of econometric models have been employed to analyze APT in the past. There have been four major types of models. The first one is the classical specification developed in earlier studies by Wolffram (1971), Houck (1977), and Ward (1982). In a representative study of this type, Kinnucan and Forker (1987) analyze the farm-retail price transmission for four dairy products in the United States, using monthly data from January 1971 to December 1981. Results reveal that the farm-retail price transmission process in the dairy sector is asymmetric. Retail dairy product prices adjust more rapidly and more fully to increases in the farm price of milk than to decreases. The second type of model considers the nonstationary property of data and incorporates cointegration concept into the analysis. As one of the seminal studies of this type, for instance, von Gramon-Taubadel (1998) uses cointegration and error correction representation to analyze the transmission between producer and wholesale pork prices in northern Germany. The analyses demonstrate that the price transmission is asymmetric and the margin is corrected more rapidly when it is squeezed relative to its long-term level, than when it is stretched. More recently, Koutroumanidis et al. (2009) use the similar technique in analyzing the asymmetric price transmission between the producer and consumer prices in the sector of forest products in Greece.

The third type of APT studies utilizes the regime switching model. It improves the specification further by adding a threshold autoregression mechanism to a standard error correction model. This assumes that the price relationship as a whole depends on a state variable, which can be one of the explanatory variables. Generally, the level of the state variable, relative to a threshold value, describes different states of the world, or regimes, hence the name of regime switching models. This type of model is especially helpful in considering transaction costs between spatial markets. For example, Goodwin and Piggott (2001) evaluate daily price linkages among several corn and soybean markets in North Carolina. The results confirm the presence of thresholds and generate a strong support for market integration, though adjustments following shocks may take many days to be completed. The model further suggests much faster adjustments in response to deviations from equilibrium than is the case when the threshold behavior is ignored. The final type of APT studies exploits the error correction model with threshold cointegration. The rationale is that if the true long-term relationship between two prices is asymmetric, a test for cointegration based on a symmetric long-term equilibrium may result in misleading findings. A solution to this problem is proposed by Enders and Granger (1998), who introduce Threshold Autoregressive (TAR) and Momentum Threshold Autoregressive (MTAR) cointegration. Correspondingly, the error correction terms are revised based on threshold cointegration. With this methodology, Ghoshray (2007) examines the relationship between Canadian and US durum wheat prices. The US price responds to restore the equilibrium relationship with the corresponding Canadian price, while the Canadian price evolves independently.

Given the evolution of the US import wooden bed market and the current status of empirical analyses of price transmission, the objective of this study is to examine the dynamics between the import prices of wooden beds from China and Vietnam. The latest error correction model with threshold cointegration is employed for the price analysis. At the beginning, linear cointegration analyses, including Johansen and Engle-Granger two-step approaches, are applied to evaluate the cointegration relationship.

Then the analysis is extended to nonlinear threshold cointegration. At the end, an asymmetric error correction model with threshold cointegration is utilized to analyze the short-term relationship. The data used are monthly import prices of wooden beds for China and Vietnam from January 2002 to January 2010. This study will help us understand the progressively more integrated international market for wooden bedroom furniture where the interaction of two major supplying countries, i.e., China and Vietnam, has been prominent in recent years.

The rest of the article is organized as follows. In the second section, a review of the US wooden bed market is presented. The emphasis is on the expansion of China and Vietnam in the import wooden bed market of the United States in the past ten years. In the third section, the methodology employed is presented, including the linear and threshold cointegration approaches and the asymmetric error correction model. Following that, the data used and the empirical results are presented. Finally, the section concludes with closing remarks.

2. Import wooden bed market in the United States

The wooden furniture industry contains both locally craft-based firms and large volume producers (Kaplinsky et al., 2003). Under increasing global competition, furniture manufacturers have retained some segments primarily for high-end, local, expensive, or design-led products (e.g., kitchen cabinet). Nevertheless, mass-producing furniture has become a viable manufacturing strategy with the advent of ready-to-assemble packing technology. This product innovation has paved the way for firms to design, manufacture and ship products in large quantities for both local and export markets. Firms that massively produce flat-pack furniture usually target low- to medium-price markets. The following brief review focuses on the growth of global market for mass-producing wooden furniture in recent years. In particular, the expansion of China and Vietnam in the import wooden bed market of the United States is highlighted.

2.1. Factors behind China's export growth

The globalization of the furniture industry has been accelerated since the 1980s. A primary reason behind that has been the establishment of global production networks using Chinese subcontractors (Drayse, 2008). The explosive export growth of China in the wooden furniture market is attributable to its comparative advantages and constructive actions in several aspects. First of all, many segments of the wooden furniture industry are labor intensive. The wage rate in the Chinese furniture industry has been about 5% to 10% of that in developed economies so China has great advantage in labor cost. Second, furniture manufacturing is resource intensive too. China has become one of the major timber imports worldwide in recent years. Major suppliers to China's timber market have been Indonesia, Malaysia and Russia (Katsigris et al., 2004). Combined with its cheap labor, the abundant supply of timber worldwide has allowed China to gain considerable advantages in reducing production costs.

Furthermore, the Chinese furniture industry has benefited greatly from Taiwanese investment. When China started its economic reform in 1978, Taiwanese entrepreneurs crossed the strait to take advantage of much lower production costs (Drayse, 2008; Tung, 2009). This has been a critical factor behind the growing furniture industry in China, especially during its early stage of development. These furniture firms from Taiwan have provided infusions of capital and technological know-how, and even more remarkably, connections with customers in the United States and other developed countries.

Finally, in exploring international markets, many Chinese firms have formed alliance with distributors and retailers in the United States or other importing countries. Thus, it has been very common for the furniture firms in the United States to import wooden bedroom furniture from China and supplement their domestic production. Specifically, U.S. ITC (2004) reports that one-third of wooden bedroom furniture imports from China over 2001 – 2003 have been imported by US manufacturing firms. This type of sourcing and joint ventures shifts the corporate power in the global furniture industry to large transnational and branded manufactures. At the same time, the strategy allows China as a developing country to gain market share rapidly in the United States.

Overall, the globalization of the furniture industry since the 1980s has accelerated through a convergence of technological innovations, the implementation of economic development strategies and regulatory regimes favoring global investment and trade, and the emergence of furniture manufacturers and retailers with the capacity to develop global production and distribution networks (Drayse, 2008). China has been one of the key players in this wave of globalization of the furniture market. In particular, the growth of the modern Chinese furniture industry has benefited from the cheap labor and abundant capital in coastal cities, plentiful timber supplies, supportive government policies, and dynamic local and external networks established through various alliances and strategies.

2.2. Antidumping investigation against wooden bedroom furniture from China

US imports of wooden bedroom furniture have been growing steadily in recent years. In particular, as a representative product, wooden beds also have experienced rapid import expansion with a total import of $140.8 million in 1996 and $895.2 million in 2003 (U.S. ITC, 2010). Historically, the major suppliers of wooden beds in the US import market are Canada, Indonesia, Italy and Taiwan. For example, the import value from Canada was $35.0 million in 1996, accounting for 24.8% of the import share; but in 2003, its corresponding numbers were $66.1 million and 7.4%. In contrast, newly industrialized Asian countries, especially China, have demonstrated their strong potential to expand in the US market. The import value of wooden beds from China was only $5.6 million in 1996, but quickly climbed to $197.4 million in 2001, and reached $477.3 million in 2003 (with a 53.3% market share). Overall, US import market structure of wooden beds has undergone dramatic changes, and by 2003, the rapid import growth has aroused wide concerns and resulted in strong reaction from a number of domestic firms (Drayse, 2008).

In October 2003, a group of furniture firms in the United States filed a petition with the ITC and the Department of Commerce. The petitioners alleged that wooden bedroom furniture from China has been dumped in the United States at less than fair value. After a 15-month investigation, several conclusions were reached in December 2004 by U.S. ITC (2004). The volume of subject imports was large both in absolute terms and relative to the consumption in the United States during the period of investigation (2001 – 2003). There was also a moderate to high degree of substitutability between domestic and imported wooden bedroom furniture, and therefore, price has been a critical factor for consumers to determine their purchases. In addition, the subject imports had a great adverse impact on the domestic industry. From 2001 to 2003, the capacity of the US furniture industry fell by 2.9% and its quantity of shipments declined by 9.8%. Therefore, the Department of Commerce and ITC concluded that the US furniture industry was materially injured by wooden bedroom furniture imports from China. Based on the damage to the US furniture industry, final antidumping duties ranging from 0.83% to 198.08% have been imposed on Chinese firms since January 2005 (U.S. ITC, 2004). Overall, the duty rates were low (i.e., less than 6%) for most firms but prohibitively high for small firms in the import market.

The long antidumping investigation resulted in some temporary decline of imports from China from July 2004 to early 2005. However, after that, imports from China have been stable or increasing for several years until 2008 (Fig. 1). From 2004 to 2009, the total imports of wooden beds from China were $480.5, 539.5, 581.6, 547.2, 418.1, 219.4 million; the corresponding share of China is 49.0%, 43.7%, 46.5%, 43.3%, 35.4% and 29.7%. The decrease of imports from China in 2008 and 2009 are seemingly related to the economic recession and reduced market demand in the United States, and more importantly, to the rapid rising of another competitor — Vietnam.

2.3. Vietnam's growth

In the import wooden bed market of the United States, Vietnam has grown up from a virtual nonentity at the start of the 21st century to currently a leading supplier in less than ten years. The growth of furniture shipments entering US ports from Vietnam has been staggering. Vietnam began its sporadic exports of wooden beds to the United States with a value of $0.5 million only in 2001, but reached $357.8 million in 2008 and $238.5 million in 2009 (Fig. 1). Imports from Vietnam have increased so fast that they account for 30.3% of the market share in 2008 and 32.3% in 2009. Furthermore, since January 2009, the monthly import value from Vietnam has been consistently larger than that from China. At present, in spite of the declining overall market demand and imports by the United States, Vietnam has passed China and become the leading supplier to the import wooden bed market.

Why has Vietnam become such a hot spot for furniture manufacturing? A major reason discussed in the literature has been that Vietnam is one of the few places in the world with lower labor costs than China now (Drayse, 2008). The political climate in Vietnam has been relatively stable in recent years and its workforce is generally diligent. Vietnam also has benefited from the US antidumping investigation that resulted in duties being placed on Chinese wooden bedroom furniture since 2005 (Wan et al., 2010a).

Survey studies revealed that a large number of Taiwanese furniture firms have moved their factories to Vietnam in recent years (Tung, 2009). These manufacturers have shifted bedroom furniture production to Vietnam, and consequentially, increased the export of wooden beds from Vietnam to the United States.

Overall, the import market of wooden beds in the United States has been driven by the rapid growth from China, and in more recent years, from Vietnam. As the monthly imports from Vietnam have exceeded these from China since 2009, this motivates us to investigate the interaction between China and Vietnam in this market through a price analysis with contemporary time series econometric models.

3. Methodology

Cointegration has been widely used to investigate relationship among price variables. The two major cointegration methods are Johansen and Engle-Granger two-step approaches. Both of them assume symmetric relationship between variables. In recent years, threshold cointegration has been increasingly used in price transmission studies. Balke and Fomby (1997) propose a two-step approach for examining threshold cointegration on the basis of the approach developed by Engle and Granger (1987). Enders and Granger (1998) and Enders and Siklos (2001) further generalize the standard Dickey-Fuller test by allowing for the possibility of asymmetric movements in time-series data. This makes it possible to test for cointegration without maintaining the hypothesis of a symmetric adjustment to a long-term equilibrium. Thereafter, the method has been widely applied to analyze asymmetric price transmission. Representative studies are analyses of Ghanaian maize market (Abdulai, 2000), Swiss pork prices (Abdulai, 2002), US and Canadian durum wheat prices (Ghoshray, 2007), stock market indexes (Shen et al., 2007), and rice export prices between Thailand and Vietnam (Ghoshray, 2008). In this study, linear cointegration, threshold cointegration, and an asymmetric error correction model are employed to examine the price dynamics in the import wooden bed market in the United States.

3.1. Linear cointegration analysis

The focus variables in this study are import prices of wooden beds for two supplying countries. Let V_t represent the import price at month *t* for Vietnam, and similarly, H_t for China. As usual, their properties of nonstationarity and order of integration can be assessed using the Augmented Dickey-Fuller (ADF) Test (Dickey and Fuller, 1979). If both the price series appear to have a unit root, then it is appropriate to conduct cointegration analysis to evaluate their interaction. Two cointegration methods widely used are the Johansen approach and Engle-Granger two-step approach (Enders, 2004). The Johansen approach is a multivariate generalization of the Dickey-Fuller test (Johansen, 1988; Johansen and Juselius, 1990). It concentrates on the relationship between the rank of a matrix and its characteristic roots in a vector autoregression. The Johansen approach starts with a vector autoregressive model and then reformulates it into a vector error correction model as follows:

$$X_t = \pi_1 X_{t-1} + \cdots + \pi_k X_{t-K} + \varepsilon_t \tag{1a}$$

$$\Delta X_t = \sum_{i=1}^{K-1} \Gamma_i \Delta X_{t-i} + \Pi X_{t-K} + \varepsilon_t \tag{1b}$$

where X_t is a vector of the price series V_t and H_t, K is the number of lags, and ε_t is the error term. The relationship among the coefficients for the two equations is $\Gamma_i = -I + \sum_{j=1}^{i} \pi_j$ and $\Pi = -I + \sum_{h=1}^{K} \pi_h$, where I is an identity matrix. Two types of tests, i.e., the trace and maximum eigenvalue statistics, can be used to detect the number of cointegrating vectors, r, among the variables in X_t.

The Engle-Granger two-stage approach focuses on the time series property of the residuals from the long-term equilibrium relationship (Engle and Granger, 1987). For this study, it can be expressed as:

$$V_t = \alpha_0 + \alpha_1 H_t + \xi_t \tag{2}$$

$$\Delta \hat{\xi}_t = \rho \hat{\xi}_{t-1} + \sum_{i=1}^{P} \phi_i \Delta \hat{\xi}_{t-i} + \mu_t \tag{3}$$

where α_0, α_1, ρ, and ϕ_i are coefficients, ξ_t is the error term, $\hat{\xi}_t$ is the estimated residuals, Δ indicates the first difference, μ_t is a white noise disturbance term, and P is the number of lags. In the first stage of estimating the long-term relationship among the price variables V_t and H_t, the price of China is chosen to be placed on the right side and assumed to be the driving force. This considers the fact that China has been the leading supplier in the import wooden bed market of the United States for most years over the study period from 2002 to 2010. In the second stage, the estimated residuals $\hat{\xi}_t$ are used to conduct a unit root test (Engle and Granger, 1987). The number of lags is chosen so there is no serial correlation in the regression residuals. It can be selected using the Akaike Information Criterion (AIC), Bayesian Information Criterion (BIC), or Ljung-Box *Q* test. If the null hypothesis of $\rho = 0$ is rejected, then the residual series from the long-term equilibrium is stationary and the focal variables of V_t and H_t are cointegrated.

3.2. Threshold cointegration analysis

The above cointegration tests assume symmetric price transmission. Enders and Siklos (2001) propose a two-regime threshold cointegration approach to entail asymmetric adjustment in cointegration analysis. The alternative model modifies Equation (3) such that:

$$\Delta\hat{\xi}_t = \rho_1 I_t \hat{\xi}_{t-1} + \rho_2 (1 - I_t)\hat{\xi}_{t-1} + \sum_{i=1}^{P} \varphi_i \Delta\hat{\xi}_{t-i} + \mu_t \tag{4}$$

$$I_t = 1 \text{ if } \hat{\xi}_{t-1} \geq \tau \text{, 0 otherwise; or} \tag{5a}$$

$$I_t = 1 \text{ if } \Delta\hat{\xi}_{t-1} \geq \tau \text{, 0 otherwise} \tag{5b}$$

where I_t is the Heaviside indicator, P the number of lags, ρ_1, ρ_2 and φ_i the coefficients, and τ the threshold value. The lag *P* is specified to account serially correlated residuals and it can be selected using AIC, BIC, or Ljung-Box *Q* test.

The Heaviside indicator I_t can be specified with two alternative definitions of the threshold variable, either the lagged residual ($\hat{\xi}_{t-1}$) or the change of the lagged residual ($\Delta\hat{\xi}_{t-1}$). Equations (4) and (5a) together have been referred to as the Threshold Autoregression (TAR) model, while Equations (4) and (5b) are named as the Momentum Threshold Autoregression (MTAR) model. The TAR model is designed to capture potential asymmetric deep movements in the residuals (Enders and Granger, 1998; Enders and Siklos, 2001). The MTAR model is useful to take into account steep variations in the residuals; it is especially valuable when the adjustment is believed to exhibit more momentum in one direction than the other. Negative deepness (i.e., $|\rho_1| \leq |\rho_2|$) of the residuals implies that increases tend to persist, whereas decreases tend to revert quickly towards to equilibrium.

The threshold value τ can be specified as zero, given the regression deals with the residual series. Alternatively, Chan (1993) proposes a search method for obtaining a consistent estimate of the threshold value. A super consistent estimate of the threshold value can be attained with several steps. First, the process involves sorting in ascending order the threshold variable, i.e., $\hat{\xi}_{t-1}$ for the TAR model or the $\Delta\hat{\xi}_{t-1}$ for the MTAR model. Second, the possible threshold values are determined. If the threshold value is to be meaningful, the threshold variable must actually cross the threshold value (Enders, 2004). Thus, the threshold value τ should lie between the maximum and minimum value of the threshold variable. In practice, the highest and lowest 15% of the values are excluded from the search to ensure an adequate number of observations on each side. The middle 70% values of the sorted threshold variable are used as potential threshold values. Third, the TAR or MTAR model is estimated with each potential threshold value. The sum of squared errors for each trial can be calculated and the relationship between the sum of squared errors and the threshold value can be examined. Finally, the threshold value that minimizes the sum of squared errors is deemed to be the consistent estimate of the threshold.

Given these considerations, a total of four models are entertained in this study. They are TAR — Equation (5a) with $\tau = 0$; consistent TAR — Equation (5a) with τ estimated; MTAR — Equation (5b) with $\tau = 0$; and consistent MTAR — Equation (5b) with τ estimated. Since there is generally no presumption on which specification is used, it is recommended to choose the appropriate adjustment mechanism via model selection criteria of AIC and BIC (Enders and Siklos, 2001). A model with the lowest AIC and BIC will be used for further analysis.

Insights into the asymmetric adjustments in the context of a long-term cointegration relation can be obtained with two tests. First, an *F*-test is employed to examine the null hypothesis of no cointegration ($H_0 : \rho_1 = \rho_2 = 0$) against the alternative of cointegration with either TAR or MTAR threshold adjustment. The test statistic is represented by Φ. This test does not follow a standard distribution and the critical values in Enders and Siklos (2001) should be used. The second one is a standard *F*-test to evaluate the null hypothesis of symmetric adjustment in the long-term equilibrium ($H_0 : \rho_1 = \rho_2$). Rejection of the null hypothesis indicates the existence of an asymmetric adjustment process.

3.3. Asymmetric error correction model with threshold cointegration

The Granger representation theorem (Engle and Granger, 1987) states that an error correction model can be estimated where all the variables in consideration are cointegrated. The specification assumes that the adjustment process due to disequilibrium among the variables is symmetric. Two extensions on the standard specification in the error correction model have been made for analyzing asymmetric price transmission. Granger and Lee (1989) first extend the specification to the case of asymmetric adjustments. Error correction terms and first differences on the variables are decomposed into positive and negative components. This allows detailed examinations on whether positive and negative price differences have asymmetric effects on the dynamic behavior of prices. The second extension follows the development of threshold cointegration (Balke and Fomby, 1997; Enders and Granger, 1998). When the presence of threshold cointegration is validated, the error correction terms are modified further.

The following asymmetric error correction model with threshold cointegration is developed in this study:

$$\Delta H_t = \theta_H + \delta_H^+ E_{t-1}^+ + \delta_H^- E_{t-1}^- + \sum_{j=1}^{J} \alpha_{Hj}^+ \Delta H_{t-j}^+ + \sum_{j=1}^{J} \alpha_{Hj}^- \Delta H_{t-j}^- + \sum_{j=1}^{J} \beta_{Hj}^+ \Delta V_{t-j}^+ + \sum_{j=1}^{J} \beta_{Hj}^- \Delta V_{t-j}^- + \vartheta_{Ht} \tag{6a}$$

$$\Delta V_t = \theta_V + \delta_V^+ E_{t-1}^+ + \delta_V^- E_{t-1}^- + \sum_{j=1}^{J} \alpha_{Vj}^+ \Delta H_{t-j}^+ + \sum_{j=1}^{J} \alpha_{Vj}^- \Delta H_{t-j}^- + \sum_{j=1}^{J} \beta_{Vj}^+ \Delta V_{t-j}^+ + \sum_{j=1}^{J} \beta_{Vj}^- \Delta V_{t-j}^- + \vartheta_{Vt} \tag{6b}$$

where ΔH and ΔV are the import prices of China and Vietnam in first difference, E error correction terms, θ, δ, α and β coefficients, and ϑ error terms. The subscripts ^{H} and ^{V} differentiate the coefficients by country, *t* denotes time, and *j* represents lags. All the lagged price variables in first difference (i.e., ΔH_{t-j} and ΔV_{t-j}) are split into positive and negative components, as indicated by the superscripts $^+$ and $^-$. For instance, ΔV_{t-1}^+ is equal to $(V_{t-1} - V_{t-2})$ if $V_{t-1} > V_{t-2}$ and equal to 0 otherwise; ΔV_{t-1}^- is equal to $(V_{t-1} - V_{t-2})$ if $V_{t-1} < V_{t-2}$ and equal to 0 otherwise. The maximum lag *J* is chosen with the AIC statistic and Ljung-Box Q test so the residuals have no serial correlation. The two error correction terms are defined as $E_{t-1}^{+} = I_t \hat{\xi}_{t-1}$ and $E_{t-1}^{-} = (1 - I_t)\hat{\xi}_{t-1}$, which in turn are constructed from the threshold cointegration regressions in Equations (4) and (5). Note this definition of the error correction terms not only considers the possible asymmetric price in response to positive and negative shocks to the deviations from long-term equilibrium, but also incorporates the impact of threshold cointegration through the construction of Heaviside indicator in Equation (5).

A simple inspection of the estimated coefficients can offer a first insight on the presence of asymmetric price behavior and can reveal the response of individual variables to the disequilibrium in the previous periods. Note the price of China is assumed to be the driving force and the long-term disequilibrium is measured as the price spread between Vietnam and China. Thus, the expected signs for the error correction terms should be positive for China (i.e., $\delta_H^{+} > 0,\ \delta_H^{-} > 0$) and negative for Vietnam (i.e., $\delta_V^{+} < 0,\ \delta_V^{-} < 0$).

Furthermore, single or joint hypotheses can be formally formed. In this study, four types of hypotheses and F-tests are examined (Frey and Manera, 2007). The first one is Granger causality test. Whether the Chinese price Granger causes its own price or the Vietnamese price can be tested by restricting all the Chinese prices to be zero and then employing a F-test ($H_{01}: \alpha_i^{+} = \alpha_i^{-} = 0$ for all lags i simultaneously). Similarly, the test can be applied to the Vietnamese price ($H_{02}: \beta_i^{+} = \beta_i^{-} = 0$ for all lags). The second type of hypothesis is concerned with the distributed lag asymmetric effect. At the first lag, for instance, the null hypothesis is that the Chinese price has symmetric effect on its own price or the Vietnamese price ($H_{03}: \alpha_1^{+} = \alpha_1^{-}$). This can be repeated for each lag and both countries (i.e., $H_{04}: \beta_4^{+} = \beta_4^{-}$). The third type of hypothesis is cumulative asymmetric effect. The null hypothesis of cumulative symmetric effect can be expressed as $H_{05}: \sum_{i=1}^{J} \alpha_i^{+} = \sum_{i=1}^{J} \alpha_i^{-}$ for China and $H_{06}: \sum_{i=1}^{J} \beta_i^{+} = \sum_{i=1}^{J} \beta_i^{-}$ for Vietnam. Finally, the equilibrium adjustment path asymmetry can be examined with the null hypothesis of $H_{07}: \delta^{+} = \delta^{-}$ for each equation estimated.

4. Data and variables

Under the Harmonized Tariff Schedule (HTS) system, the commodity of wooden beds is classified as HTS 9403.50.9040. The period covered in this study is from January 2002 to January 2010. Vietnam has been exporting wooden beds to the United States since 2001. However, the trade volumes in 2001 for Vietnam are thin and its monthly import prices are volatile. Thus, the starting period is set at January 2002. The end period reflects data availability when the data are collected. The monthly cost-insurance-freight values in dollar and quantities in piece are reported by country from (U.S. ITC, 2010). The unit price ($/piece) is calculated for China and Vietnam and used in all the analyses.

5. Empirical results

5.1. Descriptive statistics and unit root test

The descriptive statistics for the import prices of China and Vietnam are reported in Table 1. The trend of the monthly import prices is demonstrated through Fig. 2. From January 2002 to January 2010, the average import price is 148.791 ($/piece) for China and 115.526 for Vietnam; the Chinese price is 28.8% higher. The correlation coefficient is 0.25 between the two prices over the whole study period. In recent years, however, the price margin between the two prices has become more stable, and the correlation coefficient has been improved to be over 0.5. At the same time, the import value from Vietnam has grown steadily and has passed that from China since January 2009 (Fig. 1). Overall, the price of China has been usually higher than that for Vietnam, and with the increasing trading volume from Vietnam, the two prices have evolved to be closer and more correlated.

The nonstationarity properties of the two prices are examined using the ADF test. The lag length for the ADF test is determined by the AIC statistic and Ljung-Box *Q* test. The procedures in Enders (2004) are followed in making the decision of whether a trend or constant should be included in the regression. As reported in Table 1, the statistics reveal that unit roots cannot be rejected at the 1% level for the level forms of both the price variables but rejected for the first difference form. Thus, it is concluded that both the import prices for China and Vietnam are integrated of order one.

5.2. Results of the Linear Cointegration Analysis

Linear cointegration analyses are conducted using both the Johansen and Engle-Granger approach. First of all, implementation of the Johansen approach requires the determination of a lag length for the model. Based on the lowest AIC and BIC, five lags are used in the regression. Without prior information, three model specifications with trend, constant, or no intercept are entailed (Table 2). For instance, with only a constant included, the Johansen maximum eigenvalue statistic (λ_{max}) is 14.304 for the null hypothesis of no cointegrating vector between the prices of Vietnam and China. This is significant at the 10% level so the null hypothesis is rejected. However, for the null hypothesis of one cointegrating vector, the λ_{max} statistic decreases to 4.461, which is not significant at all. Thus, the maximum eigenvalue statistic concludes that there is one cointegrating vector. Similarly, the Johansen trace statistic also supports the conclusion that the prices of China and Vietnam are cointegrated.

The Engle-Granger cointegration test is executed through two steps. In the first step, the long-term relationship between the imports prices of Vietnam and China is estimated, as specified in Equation (2). The estimate for the coefficient on the Chinese price (i.e., α_1) is 0.214 with a *p*-value of 0.02. In the second step, the residual is used to conduct a unit root test with the specification in Equation (3). As reported in Table 3, the AIC and Ljung-Box *Q* statistics indicate that one lag is sufficient to address the serial correlation. The statistic from the unit root test is -0.407 and it is significant at the 1% level. Thus, the Engle-Granger approach also confirms that the import prices of Vietnam and China are cointegrated.

5.3. Results of the threshold cointegration analysis

The nonlinear cointegration analysis is conducted using the threshold autoregression models. Four models (i.e., TAR, MTAR and their consistent counterparts) are examined and the results are reported in Table 3. In selecting an appropriate lag to address possible serial correction in the residual series, a maximum lag of 12 is specified and tried at the beginning. Diagnostic analyses on the residuals through AIC, BIC and Ljung-Box *Q* statistics all reveal that a lag of three is sufficient.

In estimating the threshold values for consistent TAR and MTAR, the method by Chan (1993) is followed. It turns out that different lag specifications in the models have little impact of the final threshold values selected. The variation of the sum of squared errors by threshold value for consistent MTAR with a lag of three is presented in Fig. 3. The whole range of $\Delta\hat{\xi}_{t-1}$ is from -9.941 to 6.784. Around the values of zero and -8, the sum of squared errors is relatively low. The lowest sum of squared errors for the consistent MTAR model is 5056.542 at the threshold value of -0.451. Similarly, the best threshold value with the lowest sum of squared errors is estimated to be -8.041 for the consistent TAR model. Finally, while the four nonlinear threshold cointegration models have similar results (Table 3), the consistent MTAR model has the lowest AIC statistic of 636.430 and BIC statistic of 651.495, and therefore, is deemed to be the best.

Focusing on the results from the consistent MTAR model, the *F*-test for the null hypothesis of no cointegration has a statistic of 19.053 and it is highly significant at the 1% level. Thus, the import prices of Vietnam and China are cointegrated with threshold adjustment. Furthermore, the *F* statistic for the null hypothesis of symmetric price transmission has a value of 8.256 and it is also significant at the 1% level. Therefore, the adjustment process is asymmetric when the prices of Vietnam and China adjust to achieve the long-term equilibrium.

The point estimate for the price adjustment is -0.251 for positive shocks and -0.754 for negative shocks. Positive deviations from the long-term equilibrium resulting from increases in the Vietnamese price or decreases in the Chinese price ($\Delta\hat{\xi}_{t-1} \geq -0.451$) are eliminated at 25.1% per month. Negative deviations from the long-term equilibrium resulting from decrease in the Vietnamese price or increases in the Chinese price ($\Delta\hat{\xi}_{t-1} < -0.451$) are eliminated at a rate of 75.4% per month. In other words, positive deviations take about four months ($1/0.251 = 3.98$ months) to be fully digested while negative deviations take 1.3 months only. Therefore, there is substantially slower convergence for positive (above threshold) deviations from long-term equilibrium than negative (below threshold) deviations.

5.4. Results of the error correction model

Given the consistent MTAR model is the best among these from the threshold cointegration analyses, the error correction terms are constructed using Equations (4) and (5b). The asymmetric error correction model with threshold cointegration is estimated and the results are reported in Table 4. Diagnostic analyses on the residuals with AIC, BIC and Ljung-Box *Q* statistics select a lag of four for the model. In the equation for China, there are two coefficients significant at the 10% level or better (i.e., α_{H1}^{+}, α_{H2}^{-}). In equation for Vietnam, there are five significant coefficients (i.e., θ_V, α_{V4}^{-}, β_{V4}^{+}, δ_V^{+}, δ_V^{-}). There are two additional coefficients significant at the 15% level in each of the equations. The R^2 statistic is 0.342 for China and 0.525 for Vietnam. The AIC statistic is 690.417 for China and 641.193 for Vietnam. Overall, the model specification has a better fit on Vietnam than on China.

The hypotheses of Granger causality between the prices are assessed with *F*-tests. The *F*-statistic of 2.422 and the *p*-value of 0.02 reveal that the price of China does Granger cause the price of Vietnam. However, the *F*-statistic of 0.414 indicate that the price of Vietnam does not Grange cause the price of China. Similarly, the *F*-statistics of 2.855 for China and 1.825 for Vietnam disclose that the lagged price series have significant impacts on its own price. Thus, in the short term, the price of China has been evolving more independently while the price of Vietnam has been dependant on the price of China in the previous periods.

Several types of hypotheses are examined for asymmetric price transmission. The first one is the distributed lag asymmetric effect. In each price equation, the equality of the corresponding positive and negative coefficients for each of the four lags is tested; in total, there are eight *F*-tests for this hypothesis. It turns out that two of them are significant at the 10% level. Distributed lag asymmetric effect is found for China for its own price at lag two and for Vietnam for its own price at lag four. Furthermore, the cumulative asymmetric effects are also examined. The largest *F*-statistic is 1.931 but none of the four statistics are significant at the conventional level. Thus, there have been some distributed lag asymmetric effect but cumulative effects are symmetric.

The final type of asymmetry examined is the momentum equilibrium adjustment path asymmetries. For China, the *F*-statistic is 3.538 with a *p*-value of 0.06. The point estimates of the coefficients for the error correction terms are -0.187 for positive error correction term and 0.315 for the negative one. While the sign is wrong for the first term, both of them are not significant at the conventional level (the corresponding *p*-value for 0.315 is 0.12). Therefore, it seems that in the short term the price of China has some different responding speed to positive and negative deviations but the difference is weak. In contrast, for Vietnam, the *F*-statistic is 3.656 with a *p*-value of 0.06. Thus, there is momentum equilibrium adjustment asymmetry. The point estimates is -0.287 with a *p*-value of 0.09 for positive deviation and -0.678 with a *p*-value of less than 1% for negative deviation. The magnitude suggests that in the short term the price of Vietnam responds to the positive deviations by 28.7% in a month but by 67.8% to negative deviations. Measured in response time, positive deviations take about three months to be fully digested while negative deviations take one and half months only. Therefore, in the short term, the price of Vietnam has a much slower reaction to positive deviations from long-term equilibrium than negative deviations.

6. Conclusions

Imports of wooden beds by the United States have been over one billion dollars since 2005. With a combined market share of more than 60%, China and Vietnam have dominated the import market in recent years. The fast import growth has become a serious threat to the domestic furniture manufacturing industry and aroused wide concerns in the United States. In this study, the price dynamics between China and Vietnam is examined using threshold cointegration. The price adjustment in the short term is also analyzed through an asymmetric error correction model with threshold cointegration incorporated.

Several conclusions can be drawn from the analyses. First of all, China has been the price leader of the market and its price has been evolving more independently. This is revealed by the Granger causality test and the insignificant response to long-term price deviation by China. This result is consistent with the trade pattern of the two countries over the study period. The monthly import value from China has been over \$20 million from the beginning of the study period (i.e., January 2002) while Vietnam has only passed that since May 2005. The import price of China has been fluctuated between \$140 and \$160 per piece, which is about 29% higher than the price of Vietnam. Nonetheless, the monthly import value from Vietnam has passed that from China since January 2009. China may lose its price leadership to Vietnam over next several years if the current trade trend continues.

The transmission between the prices of China and Vietnam has been asymmetric in both the long term and short term. The threshold cointegration analysis reveals that in the long term positive deviations of the price spread between the two countries take about four months to be fully digested while negative deviations take less than one and half months. Similarly, in the short term, the error correction model reveals that Vietnamese furniture firms need three months to fully digest positive price shocks but one and half months only for negative shocks. This result is consistent with the conclusions of positive asymmetry found in the majority of spatial price studies (Frey and Manera, 2007). Overall, furniture firms are more sensitive and act more promptly when the price margin is squeezed than stretched. Several large transnational manufactures with businesses in China, Vietnam, and United States have been the major traders of wooden beds, and their market power may be the source of these asymmetric price adjustments.

At present, the import of wooden beds has declined dramatically because of the economic recession in the United States starting in late 2008. The antidumping investigation against wooden bedroom furniture from China in 2004 and subsequent duty imposition seemingly have generated some temporary trade depression effect on China and trade diversion effect to Vietnam (Wan et al., 2010b). Furthermore, with rapid growth in less than ten years and even lower prices, it appears that Vietnam has been following China’s step in the import market of wooden bedroom furniture in the United States. The result from this study discloses in detail the price relationship between these two major competitors over the last ten years. To assist the US furniture industry to compete in a global market, effective trade policies will need to take into consideration the price dynamics and competition of major supplying countries, which has been constantly evolving over time.

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Table 1 Descriptive statistics and unit root test results for the import prices of China and Vietnam.

Statistic	Import price of China		Import price of Vietnam	
	Level (H_t)	1st Diff (ΔH_t)	Level (V_t)	1st diff (ΔV_t)
Mean	148.791	—	115.526	—
Std. Dev.	11.461	—	9.882	—
Minimum	119.618	—	99.335	—
Maximum	177.675	—	150.721	—
Total obs.	97		97	
ADF with trend	-2.956 [3]	-7.394 [3]***	-2.936 [12]	-5.777 [10] ***
ADF with drift	-2.422 [3]	-7.195 [3]***	-1.161 [11]	-5.740 [10]***

Notes: The critical values are -4.04, -3.45, and -3.15 for ADF test with trend, and -3.51, -2.89, -2.58 for ADF test with a drift at the 1%, 5%, and 10% level, respectively (Enders, 2004). *, **, and *** denote significance at the 10%, 5%, and 1% level, respectively. The numbers in the bracket are lags used in the test.

Table 2 Results of the Johansen cointegration tests on the import prices of China and Vietnam.

Test	Specification	Statistic	Critical value		
			10%	5%	1%
Johansen λ_{max}					
$r = 1$	Trend	10.001	10.49	12.25	16.26
$r = 0$	Trend	20.253**	16.85	18.96	23.65
$r = 1$	Constant	4.461	7.52	9.24	12.97
$r = 0$	Constant	14.304*	13.75	15.67	20.20
$r = 1$	none	4.438	6.50	8.18	11.65
$r = 0$	none	14.300*	12.91	14.90	19.19
Johansen λ_{trace}					
$r \leq 1$	Trend	10.001	10.49	12.25	16.26
$r = 0$	Trend	30.254**	22.76	25.32	30.45
$r \leq 1$	Constant	4.461	7.52	9.24	12.97
$r = 0$	Constant	18.765*	17.85	19.96	24.60
$r \leq 1$	None	4.438	6.50	8.18	11.65
$r = 0$	None	18.738**	15.66	17.95	23.52

Notes: r is the number of cointegrating vectors. *, **, and *** denote significance at the 10%, 5%, and 1% level, respectively. The critical values are from Enders (2004).

Table 3 Results of the Engle-Granger and threshold cointegration tests.

Item	Engle-Granger	TAR	Consistent TAR	MTAR	Consistent MTAR
Estimate					
Threshold	—	0	-8.041	0	-0.451
ρ_1 †	-0.407***	-0.397***	-0.350***	-0.263**	-0.251**
	(-4.173)	(-3.590)	(-3.462)	(-2.196)	(-2.130)
ρ_2	—	-0.615***	-0.846***	-0.732***	-0.754***
	—	(-3.984)	(-4.843)	(-5.788)	(-5.957)
Diagnostics					
AIC	669.627	651.299	646.312	637.621	636.430
BIC	677.351	666.430	661.443	652.686	651.495
Q_{LB} (4)	0.773	0.974	0.983	0.869	0.868
Q_{LB} (8)	0.919	0.995	0.993	0.979	0.980
Q_{LB} (12)	0.239	0.260	0.305	0.366	0.425
Hypotheses					
Φ (H_0: $\rho_1 = \rho_2 = 0$)	—	13.734***	16.922***	19.034***	19.851***
C.V (1%)	—	8.820	9.880	8.460	8.910
C.V (5%)	—	6.280	7.410	6.200	6.560
F (H_0: $\rho_1 = \rho_2$)	—	1.385	6.309**	7.355***	8.585***
	—	[0.242]	[0.014]	[0.008]	[0.004]

Notes: † For the Engle-Granger cointegration test, ρ_1 refers to ρ in Equation (3). *, **, and *** denote significance at the 10%, 5%, and 1% level, respectively. For the Engle-Granger cointegration test, the critical value is -3.087, -3.398, and -4.008 at the 10%, 5%, and 1% level, respectively (Enders, 2004). TAR refers to the threshold autoregressive model and MTAR is the momentum threshold autoregressive model. Q_{LB} (p) denotes the significance level for the Ljung-Box Q statistic; it tests serial correlation based on p autocorrelation coefficients (p = 4, 8, 12). Φ is the threshold cointegration test with the critical values from Ender and Siklos (2001). F is a standard F-test on the asymmetry of the price transmission and the numbers in the brackets are p-values.

Table 4 Results of the asymmetric error correction model with threshold cointegration.

Item	China		Vietnam	
	Estimate	*t*-ratio	Estimate	*t*-ratio
θ	-0.146	-0.052	-3.853*	-1.777
α_1^+	-0.622***	-2.755	-0.155	-0.897
α_2^+	0.082	0.344	-0.144	-0.795
α_3^+	-0.282	-1.264	0.146	0.854
α_4^+	-0.324	-1.403	-0.193	-1.091
α_1^-	-0.314†	-1.464	-0.105	-0.641
α_2^-	-0.584***	-2.651	0.085	0.508
α_3^-	-0.041	-0.182	-0.101	-0.588
α_4^-	-0.082	-0.408	-0.413***	-2.700
β_1^+	-0.076	-0.281	-0.067	-0.322
β_2^+	0.105	0.436	0.262†	1.426
β_3^+	0.118	0.502	0.281†	1.564
β_4^+	0.129	0.588	0.425**	2.526
β_1^-	-0.052	-0.192	0.033	0.157
β_2^-	-0.211	-0.814	-0.017	-0.086
β_3^-	-0.027	-0.110	0.343*	1.787
β_4^-	0.182	0.764	-0.117	-0.639
δ^+	-0.187	-0.849	-0.287*	-1.703
δ^-	0.315†	1.535	-0.678***	-4.315
R^2	0.342	—	0.525	—
AIC	690.417	—	641.193	—
BIC	740.853	—	691.629	—
Q_{LB} (4)	0.926	—	0.521	—
Q_{LB} (8)	0.584	—	0.717	—
$H_{01}: \alpha_i^+ = \alpha_i^- = 0$ for all lags	2.855***	[0.01]	2.422**	[0.02]
$H_{02}: \beta_i^+ = \beta_i^- = 0$ for all lags	0.414	[0.91]	1.825*	[0.09]
$H_{03}: \alpha_2^+ = \alpha_2^-$	3.008*	[0.09]	0.613	[0.44]
$H_{04}: \beta_4^+ = \beta_4^-$	0.019	[0.89]	3.493*	[0.07]
$H_{05}: \sum_{i=1}^{4} \alpha_i^+ = \sum_{i=1}^{4} \alpha_i^-$	0.060	[0.81]	0.224	[0.64]
$H_{06}: \sum_{i=1}^{4} \beta_i^+ = \sum_{i=1}^{4} \beta_i^-$	0.383	[0.54]	1.931	[0.17]
$H_{07}: \delta^+ = \delta^-$	3.538*	[0.06]	3.656*	[0.06]

Notes: *, **, and *** denote significance at the 10%, 5%, and 1% level, respectively. † denote significant at the 15% level. Numbers in brackets are *p*-values. See Table 3 for Q_{LB}. For the hypotheses, H_{01} and H_{02} are Granger causality tests, H_{03} and H_{04} evaluate distributed lag asymmetric effect, H_{05} and H_{06} assess the cumulative asymmetric effect, and H_{07} is about equilibrium adjustment path asymmetric effect.

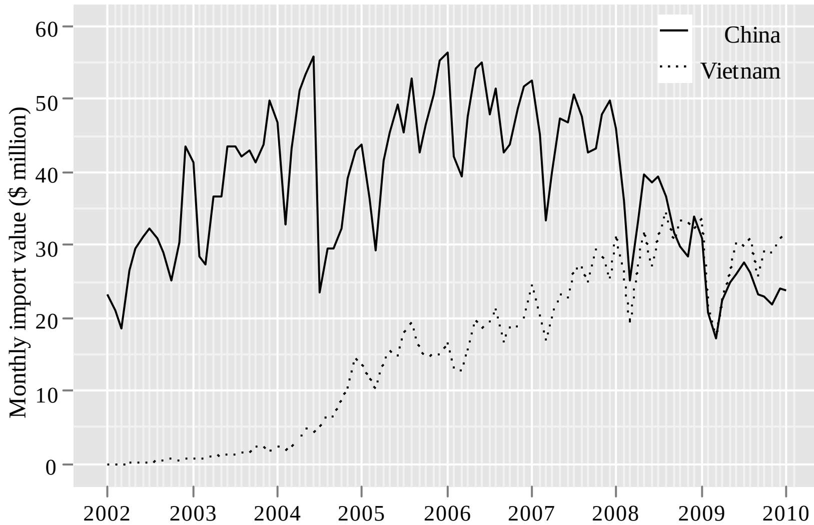

Fig. 1. Monthly import values of wooden beds from China and Vietnam (Jan. 2002 – Jan. 2010).

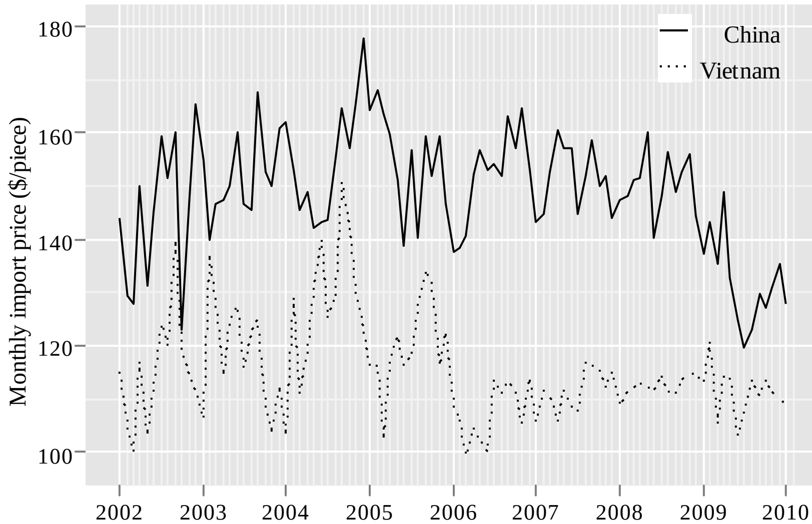

Fig. 2. Monthly import prices of wooden beds from China and Vietnam (Jan. 2002 – Jan. 2010).

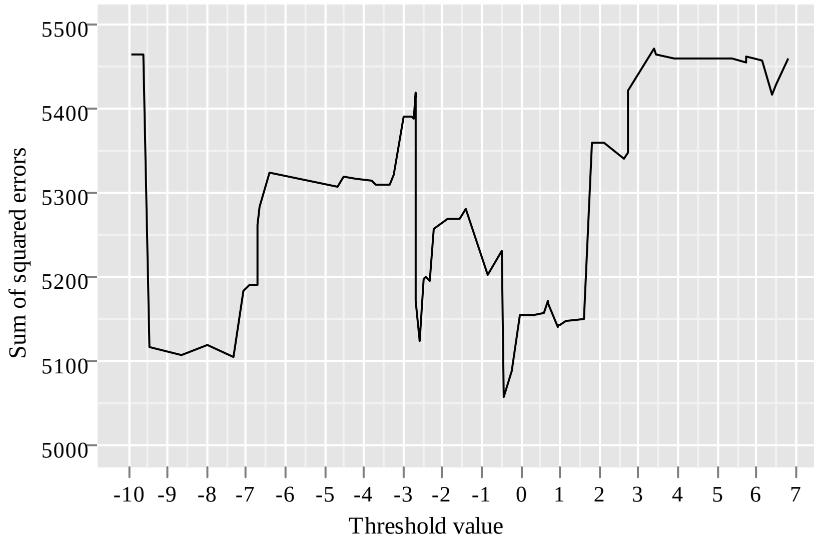

Fig. 3. Sum of squared errors by threshold value from the momentum threshold autoregression.