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國際能源價格與台灣總體經濟變數之關聯性分析

A Linkage Analysis of International Energy Prices and Macroeconomic Variables in Taiwan

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計畫報告

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國際能源價格與台灣總體經濟變數之關聯性分析

中文摘要

台灣為一能源有限且能源進口依存度高達 99.32 %的海島型經濟體。隨著國際能源價格不斷攀高，國際能源價格的變動對台灣總體經濟活動之影響將值得深入研究。本論文利用線性與非對稱的架構去評估國際能源價格(原油、煤炭、天然氣)與台灣總體經濟變數(工業生產指數、股價、利率、失業率、進口值與出口值等)之間的關係。利用 Tsay (1998) 所提出的多變量門檻模型結合非對稱動態調整過程加以分析，以能源價格變動當作一個門檻變數區分為能源價格上漲與下跌狀態，檢視在不同狀態下國際能源價格衝擊對台灣總體經濟活動的影響，進一步利用衝擊反應分析與變異數分解去評估能源價格波動對台灣總體經濟之衝擊。研究結果顯示：(1)油價的最適門檻值 2.48%，其次為天然氣價格門檻值 0.87%，最小為煤價門檻值 0.22%；(2)當油價大於門檻值時，油價對於工業生產值上的解釋能力更甚於利率；(3)當天然氣價格小於門檻值時，天然氣價格分別在股價與失業率上面都具有較大的解釋能力；(4)煤價衝擊與天然氣價格衝擊均對於台灣總體經濟活動而言具有延遲的負面影響。

關鍵詞：能源價格衝擊、總體經濟活動、多變量門檻誤差修正模型、衝擊反應分析、變異數分解

A Linkage Analysis of International Energy Prices and Macroeconomic Variables in Taiwan

Abstract

This paper applies a linear and asymmetric model to estimate the effects of international energy price shocks (including oil, coal, and natural gas prices) on Taiwan's macroeconomic activity (such as industrial production, stock price, interest rate, unemployment rate, imports and exports). We apply a multivariate threshold error correction model by Tsay (1998) to analyze the empirical data. By separating energy price changes into the decrease and increase, the energy price change as a threshold variable can analyze different impacts of energy price changes and their shock on industrial production. The variance decomposition and the impulse response functions are also employed to analyze the short-run dynamics of the variables. The preliminary findings are: (1) The optimal threshold levels are that the highest level is oil price at 2.48%, the next highest is natural gas price at 0.87%, and the lowest level is coal price at 0.22%. (2) If the change is above the threshold levels, then a change in oil price explains industrial production better than the interest rate. (3) If the change is below the threshold levels, then it appears that the change in natural gas price better explains stock prices and the unemployment rate than the interest rate. (4) Both oil price shock and natural gas shock have a delayed negative impact on macroeconomic activities.

Keywords: Energy Price Shocks; Macroeconomic Activity; Multivariate Threshold Error Correction (MVTEC) Model; Impulse Response Analysis; Variance Decomposition

1 Introduction

1.1 Research Background

Energy price shocks are generally acknowledged to have important effects on both the economic activity and macroeconomic policy of industrial countries. Huge and sudden rises in energy prices increase inflation and reduce real money balances with negative effects on consumption and output. Since the 1970s, oil price in the world market has experienced fluctuations including sharp rises during the first and second oil crises. During the periods of 1973-1974 and 1978-1979, when the Organization of Petroleum Exporting Countries (OECD) first imposed an oil embargo and the Iranian revolution disrupted oil supplies, the prices of a barrel of oil increased from $3.4 to $30. In 1990, prices rapidly rose from $16 to $26 after the Gulf War. Recently, the WTI oil future prices averaged $76 per barrel in October 2009 on the New York Mercantile Exchange (NYMEX).

Higher prices of oil driving demand for other energy have made natural gas and coal more competitive. Because price-setting is based on production costs and applications for rate increases move slowly through the bureaucratic process, natural gas volatility is quite small. On average, natural gas price variability is 3%-4%. The monthly average coal price reached a record high of $208.55 per mcf in 2008. Coal prices saw severe fluctuations during 2003 to 2009.

Domestic energy price-setting is relatively affected by projected world oil prices in the past several years. Jiménez-Rodríguez and Sánchez (2005) report that oil price increases have a direct impact on economy activity for oil-importing countries. Indeed, rising oil prices are interpreted as an indicator of an increase in scarcity and that means oil will be less available on the domestic market. This phenomenon is expected to keep the domestic energy price at high levels over the near term.

In sharp contrast to the volume of studies examining the link between oil price shocks and macroeconomic variables, there is currently not much existing literature on quantitative analyses of coal-price or gas-price shocks. There are also few analyses on the relationship between oil price shocks and financial markets such as the stock market. Market participants want a framework that identifies how energy price changes affect the stock market and labor market.

1.2 Research Motivation and Purpose

As energy prices play a critical role in influencing economic growth and economic activities, this phenomenon excites the research interest of this dissertation to address a linkage analysis of international energy prices and macroeconomic variables in Taiwan with linear and non-linear frameworks. Our research is motivated by the following reasons.

First, most studies (e.g., Burbidge and Harrison, 1984; Gisser and Goodwin, 1986; Mork, 1989; Hooker, 1996; Hamilton, 1996; Bernanke et al., 1997; Hamilton, 2003; Hamilton and Herrera, 2004) show that oil price shocks have a significantly negative impact on industrial production. However, little is known about the relationship between other energy prices and economy activities. For this reason, researchers may refocus their attention on the issue of natural gas price and coal price and their impact on economic activities.

Second, some of the related research (e.g., Mork, 1989; Mork et al., 1994; Sadorsky, 1999; Papapetrou, 2001; Hu and Lin, 2008) already consider the asymmetrical relation in terms of the impact of an oil price change or its volatility on industrial production and stock returns. However, these studies arbitrarily use zero as a cutoff point and distinguish oil price changes into up (increase) and down (decrease). This shows that the traditional approaches using predetermined value(s) as a demarcation point are rather unreasonable. They neglect the asymmetrical relation to accurately gauge varying degrees of impacts of energy price change (or volatility) on macroeconomy. To solve the neglected phenomenon, we implement rigorous econometric methods to refine the true relation.

Third, early studies about the macroeconomic consequences of energy price shocks focus on developed economies. Recent studies examine other research samples such as European countries (e.g., Mork et al, 1994; Papapetrou, 2001; Cunado and Pérez de Gracia, 2003; Jiménez-Rodríguez, 2008; Bjørnland, 2009) and Asian countries (e.g., Chang and Wong, 2003; Cunado and Pérez de Gracia, 2005; Huang et al., 2005). However, few studies investigate the relationship between energy price and macroeconomy for Taiwan. In contrast to these studies, this dissertation assesses the dynamic effect of energy price shocks on the macroeconomy in Taiwan.

Based on the aforementioned argument, the purposes of this dissertation contain two parts: The first purpose is to examine the effects of energy price shocks (including crude oil, natural gas and coal) on Taiwan's industrial production from a linear perspective. Energy prices do not affect industrial production in isolation, but through the perceived effect on the macroeconomy. Therefore, we further analyze the dynamic relationship between energy price shocks and major macroeconomic variables (including stock price, interest rate, unemployment rate, exports and imports) by applying a vector error correction (VECM) model. Next, the variance decomposition (VDC) and the impulse response functions (IRF) are employed to capture the effects of energy price shocks on the macroeconomy. The results find how each variable responds to shocks by other variables of the system and explore the response of a variable to a shock immediately or with various lags.

The second purpose focuses on the impacts of an energy price change and the shock on the macroeconomy from an asymmetric perspective. According to Sadorsky (1999), the energy price adjustment may not immediately impact macroeconomic variables. An economic threshold for an energy price impact is the amount of price increase beyond which an economic impact on industrial production and stock prices is palpable. Huang et al. (2005) propose that a change in oil price explains the macroeconomic variables better than the shock caused by the oil price if an oil price change exceeds the threshold levels. Therefore, we apply the multivariate threshold error correction model by Tsay (1998) to analyze the relevant data. By separating energy price changes into decrease (down) and increase (up), the energy price changes as the threshold variable can analyze different impacts of energy price changes on industrial production. In particular, we assess the impact of energy price fluctuations on the Taiwan economy. The impulse response and the variance decomposition analysis now follow.

2 Literature Review

Since the 1970s many studies have examined the relationship between energy prices and the macroeconomy especially for oil price shocks. However, there is an inconsistent conclusion in the literature with different estimation procedures and data. According to the different energy prices used by researchers, previous studies can be divided into three streams of research: the impact of oil price on GDP, the impact of oil price on other macroeconomic variables, and the natural gas and coal price effect.

Hamilton (1983) using Granger causality examines the impact of oil price shocks on the United States economy, indicating that oil price increases partly account for every United States recession. Many researchers extend and reinforce Hamilton's basic findings using different estimation procedures on new data (e.g., Burbidge and

Harrison, 1984; Gisser and Goodwin, 1986; Mork, 1989; Hooker, 1996; Hamilton, 1996; Bernanke et al., 1997; Hamilton, 2003; Hamilton and Herrera, 2004). These studies conclude that there is a significant negative correlation between increases in oil prices and the subsequent recessions in the United States, but that oil price changes have different impacts on economies over time.

By separating oil price changes into negative and positive, Mork (1989) finds that there is an asymmetrical relationship between oil price and real output. When the oil price is increasing, the increase in the cost of production and the decrease in the cost of resource allocation often offset each other. Mory (1993) follows Mork's (1989) measures and presents that positive oil price shocks Granger-cause the macroeconomic variables. Mork et al. (1994) again confirm that oil price shock induced inflation reduces real balances for seven industrialized countries. Lee et al. (1995) find that an oil shock in a price stable environment is more likely to have greater effects on GDP growth than those occurring in a price volatile environment. Jiménez-Rodríguez and Sánchez (2005) find that oil price increase has a larger impact on GDP growth than oil price declines.

In addition to exploring the relationship between oil price shocks and GDP, some economists have emphasized the relationship between oil price shocks and other macroeconomic variables. The first part is related to the macroeconomic level. Several models (e.g., Rasche and Tatom, 1981; Bruno and Sachs, 1982, Hamilton, 1983) and diverse episodes for oil price shocks (e.g., Davis, 1986; Carruth et al., 1998; Ferderer, 1996) present that an oil shock is one of the important influences on the macroeconomy. The directions for the causal relationship between oil price and macroeconomy can be concluded in four parts. First, oil price changes significantly impact economic activity (e.g., Papapetrou, 2001; Ewing et al., 2006; Jiménez-Rodríguez, 2008; Farzanegan and Markwardt, 2009). Second, there is an asymmetric correlation between oil price and the macroeconomy (e.g., Loungani, 1986; Mork, 1989; Lee et al., 1995; Hamilton, 2003; Cunado and Pérez de Gracia, 2003; Cunado and Pérez de Gracia, 2005; Jiménez-Rodríguez, 2009). Third, some researchers show effects of oil price shocks at a disaggregate level.

The second part is related to stock markets. Kaneko and Lee (1995) find that Japanese stock prices are affected by oil price shocks. Jones and Kaul (1996) further investigate the reaction of stock prices to oil price shocks and what may justify these movements. By using a cash-flow/dividend valuation model (i.e., Campbell, 1991), they find that oil prices can predict stock returns and output on their own. Sadorsky (1999) discovers that oil price movements can explain more of the forecast error variance of stock returns than can interest rates. Some studies (e.g., Lo and MacKinlay, 1990; Kaul and Seyhum, 1990; Sadorsky, 2003; Park and Ratti, 2008) propose that an increased volatility of oil prices significantly depresses real stock returns. Bjørnland (2009) indicates that following a 10% increase in oil prices, Norway's stock returns increase by 2.5%. Apergis and Miller (2009) also find that different oil market structural shocks play a significant role in explaining the adjustments in stock returns.

The third part involves the labor market. A clear negative relationship between oil prices and employment is reported by Rasche and Tatom (1981), Hamilton (1983), Keane and Prasad (1996), Uri (1996), Raymond and Rich (1997), among others. Keane and Prasad (1996) further indicate that oil price increases reduce employment in the short run, but tend to increase total employment in the long run. An oil price decrease depresses demand for some sectors, and unemployed labor is not immediately shifted elsewhere (Hamilton, 2003). However, oil price changes impact unemployment when the changes in oil prices persist for a long time as adjustments in employment (Keane and Prasad, 1996). Carruth et al. (1998) present an asymmetrical relationship among unemployment, real interest rates, and oil prices, meaning that oil price increases cause employment growth to decline more than oil price decreases cause employment growth to increase. Davis and Haltiwanger (2001) find an oil price shock can explain 25% of the cyclical variability in employment growth from 1972 to 1988.

Most studies show the effect of oil price shocks, but rarely consider the effect of natural gas or coal price shocks. Coal and natural gas are the two main alternative sources of energy. There are three effects of changing natural gas price controls: on regional economic activity (e.g., Leone, 1982), on inflation (e.g., Ott and Tatom, 1982), and on the distribution of income between households and suppliers (e.g., Stockfisch, 1982). Hickman et al. (1987) examine the correlation between natural gas price and industrial production. They indicate that a 10% increase in natural gas price affects the same effect on real GDP growth. Jin et al. (2009) find that energy prices have significant negative effects on real economic growth and oil price shocks are greater than other resources. Lutz and Meyer (2009) observe that a stabilizing effect via international trade and domestic structural change on the GDP of oil importing countries with a permanent oil price increase occurs.

Researchers have begun to analyze the causality relationship between coal consumption and economic growth in recent years. Yang (2000) shows a causality relationship between coal consumption and economic growth in Taiwan. Yoo (2006) finds that bidirectional causality running from GDP to coal consumption exists in South Korea. Both Li et al. (2008) and Li et al. (2009) cover that there is unidirectional causality between coal consumption and GDP in China and Japan. However, there are few studies specifically addressing coal price with economic growth.

3 Methodology

3.1 Unit Root Tests

3.1.1 Augmented Dickey Fuller (ADF) Test

Dickey and Fuller (1979) consider a autoregressive process *AR*(1) model $y_t = a_1 y_{t-1} + \varepsilon_t$, where the disturbances, ε_t, are assumed to be white noise, conditional on past y_t, and the first observation, y_1, is assumed to be fixed. By subtracting y_{t-1} from both sides of the equation, we can rewrite the model as follows: $\Delta y_t = \gamma y_{t-1} + \varepsilon_t$, where $\gamma = a_1 - 1$. The unit root test is equivalent to testing $\gamma = 0$, that is, that there exists a unit root. The standard *t*-statistic for $\hat{\gamma}$ can be used to test $\gamma = 0$, but with the Dickey-Fuller critical values.

However, simple unit root test described above is valid only if the series is an *AR*(1) process. If the series is correlated at higher order lags, the assumption of white noise disturbances is violated. Dickey and Fuller (1981) make a parametric correction for higher order correlation by assuming that the $\{y_t\}$ follows an *AR*(*p*) process and extending model as follows:

$$y_t = a_0 + a_1 y_{t-1} + a_2 y_{t-2} + \ldots + a_{p-1} y_{t-p+1} + a_p y_{t-p} + \varepsilon_t \tag{1}$$

By adding and subtracting $a_p y_{t-p+1}$ from both sides of the equation then the differenced form is:

$$y_t = a_0 + a_1 y_{t-1} + a_2 y_{t-2} + \ldots + a_{p-2} y_{t-p+2} + (a_{p-1} + a_p) y_{t-p+1} - a_p \Delta y_{t-p+1} + \varepsilon_t \tag{2}$$

Next, add and subtract $(a_{p-1} + a_p) y_{t-p+2}$ to obtain:

$$y_t = a_0 + a_1 y_{t-1} + a_2 y_{t-2} + \ldots - (a_{p-1} + a_p) \Delta y_{t-p+2} - a_p \Delta y_{t-p+1} + \varepsilon_t \tag{3}$$

Continuing in this fashion, we get:

$$\Delta y_t = a_0 + \gamma y_{t-1} + \sum_{i=2}^{p} \beta_i \Delta y_{t-i+1} + \varepsilon_t , \tag{4}$$

where $\gamma = -\left(1 - \sum_{i=2}^{p} a_i\right)$ and $\beta_i = \sum_{j=i}^{p} a_j$

In Eq. (4), the coefficient of interest is γ. If $\gamma = 0$, the equation is entirely in first differences and so has a unit root. Three ADF test actually consider three different regression equations that can be used to test for the presence of a unit root:

$$\Delta y_t = \gamma y_{t-1} + \sum_{i=2}^{p} \beta_i \Delta y_{t-i+1} + \varepsilon_t \tag{5}$$

$$\Delta y_t = a_0 + \gamma y_{t-1} + \sum_{i=2}^{p} \beta_i \Delta y_{t-i+1} + \varepsilon_t \tag{6}$$

$$\Delta y_t = a_0 + \gamma y_{t-1} + a_2 t + \sum_{i=2}^{p} \beta_i \Delta y_{t-i+1} + \varepsilon_t . \tag{7}$$

The differences between the three regressions concerns the presence of the deterministic elements a_0 and $a_2 t$. Without an intercept and time trend belongs in Eq. (5); with only the intercept belongs in Eq. (6); and with both an intercept and trend belongs in Eq. (7). If the coefficients of a difference equation sum to one, at least one characteristic root is unity. If $\Sigma a_i = 1$ and $\gamma = 0$, the system has a unit root.

3.1.2 The Kwiatkowski, Phillips, Schmidt, and Shin (KPSS) Test

Kwiatkowski et al. (1992) propose a test of the null hypothesis that an observable series is stationary around a deterministic trend. The series is expressed as the sum of deterministic trend, random walk, and stationary error, and the test is the LM test of the hypothesis that the random walk has zero variance. The KPSS statistic is based on the residuals from the OLS regression of y_t on the exogenous variables x_t:

$$y_t = x_t' \delta + \varepsilon_t$$

The Lagrange Multiplier (LM) statistic can be defined as:

$$LM = \sum_{t=1}^{T} S_t^2 / \hat{\sigma}_\varepsilon^2 ,$$

where S_t is a cumulative residual function (i.e., $S_t = \sum_{i=1}^{t} \hat{\varepsilon}_i, \quad i = 1, 2, \ldots, T$). We point out that the estimator of δ used in this calculation differs from the estimators for δ used by detrended GLS since it is based on a regression involving the original data and not on the quasi-differenced data. Finite sample size and power are considered in a Monte Carlo experiment.

Prior to performing the Johansen co-integration method, we need to determine the appropriate number of lag length of the VAR model. The Bayesian information criterion (BIC) (Schwarz, 1978) is employed. The test criteria to determine appropriate lag lengths and seasonality are the multivariate generalizations of the BIC. The BIC criterion is a purely statistical technique and allows data themselves to select optimal lags. Given any two estimated models, the model with the lower value of BIC is the one to be preferred. The selection of lag order of Δy_{t-i} can be used by the Bayesian information criterion (BIC):

$$BIC=-2*ln(L)+k*ln(n) \qquad (8)$$

where n is the number of observations, k is the number of free parameters to be estimated and L is the maximized value of the likelihood function for the estimated model. The BIC penalizes free parameters more strongly than does the Akaike information criterion (AIC) (Akaike, 1969).

3.2 Cointegration Analysis

The Johansen co-integration method is provided by Johansen (1988) and Johansen and Juselius (1990). This procedure applying maximum likelihood estimators circumvent the low-power of using Granger two-step estimators and can estimate and test for the presence of multiple cointegrating vectors. Moreover, this test allows the researcher to test restricted versions of the cointegrating vectors and speed of adjustment parameters.

Let $\mathbf{y_t}$ denotes the $(n\times 1)$ vector $(y_{1t,}y_{2t,}\cdots,y_{nt})$. The maintained hypothesis is that y_t follows a VAR(P) in levels and all of the elements for y_t are $I(1)$ process.

$$\mathbf{y_t} = A_1\mathbf{y_{t-1}} + \mathrm{A}_2\mathbf{y_{t-2}} + \ldots + \mathrm{A}_p\mathbf{y_{t-p}} + \varepsilon_t, \quad t = 1, 2, \ldots, T \tag{9}$$

where $\varepsilon_t \overset{i.i.d.}{\sim} N(0,\Omega)$.

Eq. (10) can be put in a more usable form by subtracting y_{t-1} from each side to obtain:

$$\Delta\mathbf{y_t} = (A_1 - I)\mathbf{y_{t-1}} + \mathrm{A}_2\mathbf{y_{t-2}} + \ldots + \mathrm{A}_p\mathbf{y_{t-p}} + \varepsilon_t, \quad t = 1, 2, \ldots, T \tag{10}$$

Now add and subtract $(A_1 - I)\mathbf{y_{t-2}}$ to obtain:

$$\Delta\mathbf{y_t} = (A_1 - I)\mathbf{y_{t-1}} + (A_2 + A_1 - I)\mathbf{y_{t-2}} + \ldots + \mathrm{A}_p\mathbf{y_{t-p}} + \varepsilon_t, \quad t = 1, 2, \ldots, T \tag{11}$$

Continuing in this fashion, we obtain:

$$\Delta\mathbf{y_t} = \sum_{i=1}^{p-1} \pi\Delta\mathbf{y_{t-i}} + \pi\Delta\mathbf{y_{t-p}} + \varepsilon_t \tag{12}$$

where $\pi = -\left(I - \sum_{i=1}^{p} A_i\right)$

$$\pi_i = -\left(I - \sum_{j=1}^{i} A_j\right)$$

Suppose we obtained the matrix π and order the n characteristic roots such that $\lambda_1 > \lambda_2 > ... > \lambda_n$. If the variables in $\mathbf{y_t}$ are not cointegrated, the rank of π is zero and all these characteristic roots will equal zero. Similarly, since $\ln(1) = 0$, each of the expressions $\ln(1 - \lambda_i)$ will equal zero if the variables are not cointegrated.

Suppose that each individual variable $\mathbf{y_{it}}$ is $I(1)$ and linear combinations of $\mathbf{y_t}$ are stationary. That implies π can be shown as

$$\pi = \alpha\beta'$$

where β is the matrix of cointegrating parameters, and α is the matrix of the speed of adjustment parameters. The number of cointegrating relations relies on the rank of π, and the rank of π is:

(1) $\text{rank}(\pi) = n$, λ is full rank means that all components of y_t is a stationary process.

(2) $\text{rank}(\pi) = 0$, λ is null matrix meaning that there is no cointegration relationships.

(3) $0 < \text{rank}(\pi) = r < n$, the variables for y_t are cointegrated and the number of cointegrating vectors is r.

The test for the number of characteristic roots that are insignificantly different from unity can by conducted using the following two test statistics:

(1) Trace test:

$$\lambda_{trace}(r) = -T \sum_{i=r+1}^{n} \ln(1 - \hat{\lambda}_i),$$

$$H_0 : \text{rank}(\pi) \le r,$$

$$H_1 : \text{rank}(\pi) > r$$

where $\hat{\lambda}_i$ is the estimated values of the characteristic roots (also called eigenvalues) obtained from the estimated π matrix, r is the cointegrating vector, and T is the number of usable observations. The statistic tests the null hypothesis that the number of distinct cointegrating vectors is less than or equal to r against a general alternative. If there is no cointegrating vector, it should be clear that λ_{trace} equals zero when all $\hat{\lambda}_i = 0$. The further the estimated characteristic roots are from zero, the more negative is $\ln(1 - \hat{\lambda}_i)$ and the larger the λ_{trace} statistic.

(2) Maximum eigenvalues test:

$$\lambda_{\max}(r, r+1) = -T \ln(1 - \hat{\lambda}_{r+1})$$

H_0: there are r cointegrating vectors

H_1: there are $r+1$ cointegrating vectors

The statistic tests the null that the number of cointegrating vectors is r against the alternative of $r+1$ cointegrating vectors. If the estimated value of the characteristic root is close to zero, $\lambda_{\max}$ will be small. The critical values of the λ_{trace} and $\lambda_{\max}$ statistics follows a chi-square distribution in general.

3.3 Multivariate Threshold Error Correction (MVTEC) Model

At the beginning, we consider the univariate TAR model which is also referred to as SETAR (self-exciting TAR). The SETAR(1) can be formed as:

$$y_t = (\phi_{0,1} + \phi_{1,1} y_{t-1})(1 - I[z_{t-1} > c]) + (\phi_{0,2} + \phi_{1,2} y_{t-1}) I[z_{t-1} > c] + \varepsilon_t \tag{13}$$

where ε_t is a white noise process, $z_{t-1} = y_{t-1}$, and c represents the threshold value. $I(\cdot)$ is an index function, which equals to one if the relation in the brackets holds, and equals to zero otherwise. Eq. (13) can be treated as a multivariate threshold VAR(1). Consider a k-dimensional time series $y_t = (y_{1t}, \dots, y_{kt})'$ and assume there is a cointegration relationship among these variables, then y_t follows a multivariate threshold error correction model (MVTEC) with threshold variable z_t and delay d and can be expressed as:

$$\begin{aligned} \mathbf{y_t} = {} & (\alpha_1 + \beta_1 \theta_{t-1} + \sum_{i=1}^{p} \phi_{i,1} \mathbf{y_{t-i}})(1 - I[z_{t-d} > c]) + \\ & (\alpha_2 + \beta_2 \theta_{t-1} + \sum_{i=1}^{p} \phi_{i,2} \mathbf{y_{t-i}}) I[z_{t-d} > c] + \varepsilon_t \end{aligned} \tag{14}$$

where α_1 and α_2 are the constant vectors below and above the threshold value, respectively. p and d are the lag length of $\mathbf{y_t}$ and delay order of z_t, respectively. Both p and d are nonnegative integers. θ_{t-1} is an error correction term. The threshold variable is assumed to be stationary and have a continuous distribution. Model (14) has two regimes and is a piecewise linear model in the threshold space

z_{t-d}.

Given observations $\{\mathbf{y_t}, z_t\}$, where $t = 1, \ldots, n$, we have to detect the threshold nonlinearity of $\mathbf{y_t}$. Assuming *p* and *d* are known, the Eq. (14) can be re-written as:

$$\mathbf{y}'_\mathbf{t} = X'_t \Phi + \varepsilon'_t, \quad t = h+1, \ldots, n \qquad (15)$$

where $h = \max(p, d)$, $X_t = (1, y'_{t-1}, \ldots y'_{t-p}, \theta_{t-1})'$ is a $(pk+1)$-dimensional regressor, and $\mathbf{\Phi}$ denotes the parameter matrix. If the null hypothesis holds, then the least squares estimates of (15) are useful. On the other hand, the estimates are biased under the alternative hypothesis. Eq. (15) remains informative under the alternative hypothesis when rearranging the ordering of the setup. For Eq. (15), the threshold variable z_{t-d} assumes values in $S = \{z_{h+1-d}, \ldots z_{n-d}\}$. Consider the order statistics of *S* and denote the *i*th smallest element of *S* by $z_{(i)}$. Then the arranged regression based on the increasing order of the threshold variable z_{t-d} is

$$\mathbf{y}'_{\mathbf{t(i)+d}} = X'_{t(i)+d} \mathbf{\Phi} + \varepsilon'_{t(i)+d}, \quad i = 1, \ldots, n-h, \qquad (16)$$

where $t(i)$ is the time index of $z_{(i)}$. Tsay (1998) use the recursive least squares method to estimate (16). If $\mathbf{y_t}$ is linear, then the recursive least squares estimator of the arranged regression (16) is consistent, so the predictive residuals approach white noise. Consequently, predictive residuals are uncorrelated with the regressor $X_{t(i)+d}$.

Let $\mathbf{\Phi}_{\boldsymbol{m}}$ be the least squares estimate of $\mathbf{\Phi}$ of Eq. (16) with $i = 1, \ldots, m$; i.e., the estimate of the arranged regression using data points associated with the *m* smallest values of z_{t-d}. Tsay (1998) suggests a range of *m* (between $3\sqrt{n}$ and $5\sqrt{n}$). Different values of *m* can be used to investigate the sensitivity of the modeling results with respect to the choice. It should be noted that the ordered autoregressions are sorted by the variable z_{t-d}, which is the regime indicator in the MVTEC model. Let

$$\hat{\mathrm{e}}'_{t(m+1)+d} = \mathbf{y}_{\mathbf{t(m+1)+d}} - \hat{\mathbf{\Phi}}'_{\boldsymbol{\mu}} \mathrm{X}'_{t(m+1)+d} \tag{17}$$

and

$$\hat{\eta}_{t(m+1)+d} = \hat{\mathrm{e}}_{t(m+1)+d} \,/ \left[1 + \mathrm{X}'_{t(m+1)+d} \mathrm{V}_m \mathrm{X}_{t(m+1)+d}\right]^{1/2}, \tag{18}$$

where $\mathrm{V}_m = \left[\Sigma_{i=1}^{m} \mathrm{X}_{t(i)+d} \mathrm{X}'_{t(i)+d}\right]^{-1}$ is the predictive residual and the standardized predictive residual of regression (16). These quantities can be efficiently obtained by the recursive least squares algorithm. Next, consider the regression

$$\hat{\eta}_{t(l)+d} = \mathrm{X}'_{t(l)+d} \Psi + w'_{t(l)+d}, \quad l = m_0 + 1, \ldots, n-h, \tag{19}$$

where m_0 denotes the starting point of the recursive least squares estimation. The problem of interest is then to test the hypothesis $H_0 : \Psi = 0$ versus the alternative $H_1 : \Psi \neq 0$ in regression (19). The $C(d)$ statistic is therefore defined as:

$$C(d) = (n - p - m - kp - 1) \times \left\{ \ln\left|\mathrm{S}_o\right| - \ln\left|\mathrm{S}_1\right| \right\}, \tag{20}$$

where the delay *d* implies the test depends on the threshold variable z_{t-d}, and

$$\mathrm{S}_o = \frac{1}{n-h-m_0} \sum_{l=m_0+1}^{n-h} \hat{\eta}_{t(l)+d} \hat{\eta}'_{t(l)+d}$$

and

$$\mathrm{S}_1 = \frac{1}{n-h-m_0} \sum_{l=m_0+1}^{n-h} \hat{\mathrm{w}}_{t(l)+d} \hat{\mathrm{w}}'_{t(l)+d},$$

where $\hat{\mathrm{w}}_t$ is the least squares residual of regression (19). Under null hypothesis the y_t is linear and some regularity conditions, $C(d)$ is asymptotically a chi-squared random variable with $k(pk+1)$ degree of freedom.

3.4 Impulse Response Analysis

Consider the first-order structural VAR model with 7-variables:

$$\begin{bmatrix} 1 & b_{12} & b_{13} & \dots & b_{17} \\ b_{21} & 1 & b_{23} & \dots & b_{27} \\ \vdots & \vdots & \vdots & \ddots & \vdots \\ b_{71} & b_{72} & b_{73} & \dots & 1 \end{bmatrix} \begin{bmatrix} y_{1t} \\ y_{2t} \\ \vdots \\ y_{7t} \end{bmatrix} = \begin{bmatrix} b_{10} \\ b_{20} \\ \vdots \\ b_{70} \end{bmatrix} + \begin{bmatrix} \gamma_{11} & \gamma_{12} & \gamma_{13} & \cdots & \gamma_{17} \\ \gamma_{21} & \gamma_{22} & \gamma_{23} & \cdots & \gamma_{27} \\ \vdots & \vdots & \vdots & \ddots & \vdots \\ \gamma_{71} & \gamma_{72} & \gamma_{73} & \cdots & \gamma_{77} \end{bmatrix} \begin{bmatrix} y_{1t-1} \\ y_{2t-1} \\ \vdots \\ y_{7t-1} \end{bmatrix} + \begin{bmatrix} \varepsilon_{1t} \\ \varepsilon_{2t} \\ \vdots \\ \varepsilon_{7t} \end{bmatrix}$$

We can write the system in the compact form:

$$\mathbf{B}\mathbf{y}_{\mathbf{t}} = \mathbf{\Gamma}_{\mathbf{0}} + \mathbf{\Gamma}_{\mathbf{1}}\mathbf{y}_{\mathbf{t-1}} + \boldsymbol{\varepsilon}_{\mathbf{t}}$$

where

$$\mathbf{B} = \begin{bmatrix} 1 & b_{12} & b_{13} & \dots & b_{17} \\ b_{21} & 1 & b_{23} & \dots & b_{27} \\ \vdots & \vdots & \vdots & \ddots & \vdots \\ b_{71} & b_{72} & b_{73} & \dots & 1 \end{bmatrix}, \qquad \mathbf{y}_{\mathbf{t}} = \begin{bmatrix} y_{1t} \\ y_{2t} \\ \vdots \\ y_{7t} \end{bmatrix}, \qquad \mathbf{\Gamma}_{\mathbf{0}} = \begin{bmatrix} b_{10} \\ b_{20} \\ \vdots \\ b_{70} \end{bmatrix},$$

$$\mathbf{\Gamma}_{\mathbf{1}} = \begin{bmatrix} \gamma_{11} & \gamma_{12} & \gamma_{13} & \cdots & \gamma_{17} \\ \gamma_{21} & \gamma_{22} & \gamma_{23} & \cdots & \gamma_{27} \\ \vdots & \vdots & \vdots & \ddots & \vdots \\ \gamma_{71} & \gamma_{72} & \gamma_{73} & \cdots & \gamma_{77} \end{bmatrix}.$$

Premultiplication by B^{-1} can obtain the vector autoregressive (VAR) model in standard form:

$$\mathbf{y}_{\mathbf{t}} = \mathbf{B}^{\mathbf{-1}}\mathbf{\Gamma}_{\mathbf{0}} + \mathbf{B}^{\mathbf{-1}}\mathbf{\Gamma}_{\mathbf{1}}\mathbf{y}_{\mathbf{t-1}} + \mathbf{B}^{\mathbf{-1}}\boldsymbol{\varepsilon}_{\mathbf{t}} = \mathrm{A}_0 + \mathrm{A}_1\mathbf{y}_{\mathbf{t-1}} + e_t \qquad (21)$$

where $\mathrm{A}_0 = \mathbf{B}^{\mathbf{-1}}\mathbf{\Gamma}_{\mathbf{0}}$, $\mathrm{A}_1 = \mathbf{B}^{\mathbf{-1}}\mathbf{\Gamma}_{\mathbf{1}}$ and $\mathbf{e}_{\mathbf{t}} = \mathbf{B}^{\mathbf{-1}}\boldsymbol{\varepsilon}_{\mathbf{t}}$. For notional purposes, we can define a_{io} as element i of the vector A_0, a_{ij} as the element in row i and column j of the matrix A_1, and e_{it} as the element i of the vector e_t. Using this new notation, we can rewrite (21) in the equivalent form:

$$\begin{aligned} y_{1t} &= a_{10} + a_{11}y_{1t-1} + a_{12}y_{2t-1} + \ldots + a_{17}y_{7t-1} + e_{1t} \\ y_{2t} &= a_{20} + a_{21}y_{1t-1} + a_{22}y_{2t-1} + \ldots + a_{27}y_{7t-1} + e_{2t} \\ &\vdots \qquad\qquad\qquad\qquad \vdots \\ y_{7t} &= a_{70} + a_{71}y_{1t-1} + a_{12}y_{2t-1} + \ldots + a_{17}y_{7t-1} + e_{7t} \end{aligned} \qquad (22)$$

or

$$\begin{bmatrix} y_{1t} \\ y_{2t} \\ \vdots \\ y_{7t} \end{bmatrix} = \begin{bmatrix} a_{10} \\ a_{20} \\ \vdots \\ a_{70} \end{bmatrix} + \begin{bmatrix} a_{11} & a_{12} & \dots & a_{17} \\ a_{21} & a_{22} & \dots & a_{27} \\ \vdots & \vdots & \ddots & \vdots \\ a_{71} & a_{72} & \dots & a_{77} \end{bmatrix} \begin{bmatrix} y_{1t-1} \\ y_{2t-1} \\ \vdots \\ y_{7t-1} \end{bmatrix} + \begin{bmatrix} e_{1t} \\ e_{2t} \\ \vdots \\ e_{7t} \end{bmatrix}$$

(23)

In model (21), the stability condition is that A_0 be less than unity in absolute value. Using the backward method to iterate model (21), we can obtain:

$$\begin{aligned} \mathbf{y_t} &= A_0 + A_1(A_0 + A_1\mathbf{y_{t-2}} + \boldsymbol{\varepsilon}_\mathbf{t-1}) + \boldsymbol{\varepsilon}_\mathbf{t} \\ &= (I + A_1)A_0 + A_1^2\mathbf{y_{t-2}} + A_1\boldsymbol{\varepsilon}_\mathbf{t-1} + \boldsymbol{\varepsilon}_\mathbf{t} \end{aligned}$$

where $I = 7\times 7$ identity matrix.

Assuming the stability condition is met, so that we can write the particular solution for y_t as:

$$\mathbf{y_t} = \mu + \sum_{i=0}^{\infty} A_1^i \mathbf{e_{t-i}}. \tag{24}$$

It is important to note that the error terms (i.e., e_{1t}, e_{2t},…, e_{7t}) are components of the seven shocks e_{1t}, e_{2t},…, e_{7t}. Since $\mathbf{e_t} = \mathbf{B^{-1}}\boldsymbol{\varepsilon}_\mathbf{t}$, we can compute $\{e_{1t}\}, \{e_{2t}\}$,…, $\{e_{7t}\}$ as:

$$\begin{bmatrix} e_{1t} \\ e_{2t} \\ \vdots \\ e_{7t} \end{bmatrix} = \det(b) \begin{bmatrix} c_{11} & c_{12} & \dots & c_{17} \\ c_{21} & c_{22} & \dots & c_{27} \\ \vdots & \vdots & \ddots & \vdots \\ c_{71} & c_{72} & \dots & c_{77} \end{bmatrix} \begin{bmatrix} \varepsilon_{1t} \\ \varepsilon_{2t} \\ \vdots \\ \varepsilon_{7t} \end{bmatrix} \tag{25}$$

Using model (24), model (23) can be re-written as:

$$\begin{bmatrix} y_{1t} \\ y_{2t} \\ \vdots \\ y_{7t} \end{bmatrix} = \begin{bmatrix} \bar{y}_1 \\ \bar{y}_2 \\ \vdots \\ \bar{y}_7 \end{bmatrix} + \sum_{i=0}^{\infty} \begin{bmatrix} a_{11} & a_{12} & \dots & a_{17} \\ a_{21} & a_{22} & \dots & a_{27} \\ \vdots & \vdots & \ddots & \vdots \\ a_{71} & a_{72} & \dots & a_{77} \end{bmatrix}^i \begin{bmatrix} e_{1t-i} \\ e_{2t-i} \\ \vdots \\ e_{7t-i} \end{bmatrix} \tag{26}$$

Equation (26) expresses $(y_{1t}, y_{2t}, \dots, y_{7t})$ in terms of the $\{e_{1t}\}, \{e_{2t}\}, \dots, \{e_{7t}\}$

sequences. However, it is insightful to rewrite (26) in terms of the $\{\varepsilon_{1t}\},\{\varepsilon_{2t}\},\ldots,\{\varepsilon_{7t}\}$ sequences. Equations (25) and (26) can be combined to form:

$$\begin{bmatrix} y_{1t} \\ y_{2t} \\ \vdots \\ y_{7t} \end{bmatrix} = \begin{bmatrix} \bar{y}_{1t} \\ \bar{y}_{2t} \\ \vdots \\ \bar{y}_{7t} \end{bmatrix} + \det(b)\sum_{i=0}^{\infty} \begin{bmatrix} a_{11} & a_{12} & \ldots & a_{17} \\ a_{21} & a_{22} & \ldots & a_{27} \\ \vdots & \vdots & \ddots & \vdots \\ a_{71} & a_{72} & \ldots & a_{77} \end{bmatrix}^{i} \begin{bmatrix} c_{11} & c_{12}\ldots & c_{17} \\ c_{21} & c_{22}\ldots & c_{27} \\ \vdots & \ddots & \vdots \\ c_{71} & c_{72}\ldots & c_{77} \end{bmatrix} \begin{bmatrix} e_{1t} \\ e_{2t} \\ \vdots \\ e_{7t} \end{bmatrix}$$

Since the notation gets unwieldy, we can simplify by defining the 7×7 matrix ϕ_i with elements $\phi_{jk}(i)$:

$$\phi_i = \left[A_1^i / \det(b)\right] \begin{bmatrix} c_{11} & c_{12}\ldots & c_{17} \\ c_{21} & c_{22}\ldots & c_{27} \\ \vdots & \ddots & \vdots \\ c_{71} & c_{72}\ldots & c_{77} \end{bmatrix}$$

Hence, the moving average representation of (26) can be written in terms of the $\{\varepsilon_{y_1t}\}$, $\{\varepsilon_{y_2t}\},\ldots,$ $\{\varepsilon_{y_7t}\}$ sequences:

$$\begin{bmatrix} y_{1t} \\ y_{2t} \\ \vdots \\ y_{7t} \end{bmatrix} = \begin{bmatrix} \bar{y}_{1} \\ \bar{y}_{2} \\ \vdots \\ \bar{y}_{7} \end{bmatrix} + \sum_{i=0}^{\infty} \begin{bmatrix} \phi_{11}(i) & \phi_{12}(i) & \ldots & \phi_{17}(i) \\ \phi_{21}(i) & \phi_{22}(i) & \ldots & \phi_{27}(i) \\ \vdots & \vdots & \ddots & \vdots \\ \phi_{71}(i) & \phi_{72}(i) & \ldots & \phi_{77}(i) \end{bmatrix} \begin{bmatrix} \varepsilon_{y_1t-i} \\ \varepsilon_{y_2t-i} \\ \vdots \\ \varepsilon_{y_7t-i} \end{bmatrix}$$

or more compactly,

$$\mathbf{y_t} = \mu + \sum_{i=0}^{\infty} \phi_i \boldsymbol{\varepsilon}_{\mathbf{t-i}} \,. \tag{27}$$

The coefficients of ϕ_i can be used to generate the effects of $\left[\varepsilon_{y_1t}, \varepsilon_{y_2t}, \ldots, \varepsilon_{y_7t}\right]$ shocks on the entire time paths of the $\{y_{1t}\},\{y_{2t}\},\ldots,\{y_{7t}\}$ sequences. It should be clear that forty-nine elements $\phi_{jk}(0)$ are impact multiplier. For instance, the coefficient $\phi_{12}(0)$ is the instantaneous impact of a one-unit change in ε_{zt} on y_t. In the same way, the elements $\phi_{11}(1)$, $\phi_{12}(1)$, …, $\phi_{17}(1)$ are the one period responses of unit changes in $\left[\varepsilon_{y_1t-1}, \varepsilon_{y_2t-1}, \ldots, \varepsilon_{y_7t-1}\right]$ on y_{1t}, respectively. Updating by one period indicates that $\phi_{11}(1)$, $\phi_{12}(1)$, …, $\phi_{17}(1)$ also represent the effects of unit changes in $\left[\varepsilon_{y_1t}, \varepsilon_{y_2t}, \ldots, \varepsilon_{y_7t}\right]$ on y_{1t+1}.

The accumulated effects of unit impulses in $\left[\varepsilon_{y_1t}, \varepsilon_{y_2t}, \ldots, \varepsilon_{y_7t}\right]$ can be obtained by the appropriate addition of the coefficients of the impulse response functions. Note that after *n* periods, the effect of ε_{y_2t} on the value of y_{1t+n} is $\phi_{12}(n)$. Thus, the cumulated sum of the effects of ε_{y_2t} on the $\{y_{1t}\}$ sequence is:

$$\sum_{i=0}^{n} \phi_{12}(i)$$

Letting *n* approach infinity yields the long-run multiplier. Since the $\{y_{1t}\}$, $\{y_{2t}\}$, …, $\{y_{7t}\}$ sequences are assumed to be stationary, it must be the case that for all *j* and *k*, $\sum_{i=0}^{\infty} \phi_{jk}^2(i)$ is finite. The sets of coefficients $\phi_{11}(i)$, $\phi_{12}(i)$, …, $\phi_{77}(i)$ are called the impulse response functions. We can plot the impulse response functions (i.e., plotting the coefficients of $\phi_{jk}(i)$ against *i*) is a practical manner to visually present the behavior of the $\{y_{1t}\}$, $\{y_{2t}\}$, …, $\{y_{7t}\}$ series in response to the various shocks.

3.5 Variance Decomposition

If we use the equation (27) to conditionally forecast y_{t+1}, the one-step ahead forecast error is $\phi_0 \varepsilon_{t+1}$. In general,

$$\mathbf{y_{t+n}} = \mu + \sum_{i=0}^{\infty} \phi_i \boldsymbol{\varepsilon}_{\mathbf{t+n-i}},$$

So that the *n*-period forecast error $\mathbf{y_{t+n}} - E_t \mathbf{y_{t+n}}$ is

$$\mathbf{y_{t+n}} - E_t \mathbf{y_{t+n}} = \sum_{i=0}^{n-1} \phi_i \boldsymbol{\varepsilon}_{\mathbf{t+n-i}}$$

Forecasting solely on the $\{x_{1t}\}$ sequence, the *n*-step ahead forecast error is:

$$\begin{aligned}\mathbf{y_{1t+n}} - E_t \mathbf{y_{1t+n}} = {} & \phi_{11}(0)\boldsymbol{\varepsilon}_{\mathbf{y_1 t+n}} + \phi_{11}(1)\boldsymbol{\varepsilon}_{\mathbf{y_1 t+n-1}} + \cdots + \phi_{11}(n-1)\boldsymbol{\varepsilon}_{\mathbf{y_1 t+1}} \\ & + \phi_{12}(0)\boldsymbol{\varepsilon}_{\mathbf{y_2 t+n}} + \phi_{12}(1)\boldsymbol{\varepsilon}_{\mathbf{y_2 t+n-1}} + \cdots + \phi_{12}(n-1)\boldsymbol{\varepsilon}_{\mathbf{y_2 t+1}} \\ & + \phi_{13}(0)\boldsymbol{\varepsilon}_{\mathbf{y_3 t+n}} + \phi_{13}(1)\boldsymbol{\varepsilon}_{\mathbf{y_3 t+n-1}} + \cdots + \phi_{13}(n-1)\boldsymbol{\varepsilon}_{\mathbf{y_3 t+1}} \\ & + \cdots + \phi_{1n}(0)\boldsymbol{\varepsilon}_{\mathbf{y_n t+n}} + \phi_{1n}(1)\boldsymbol{\varepsilon}_{\mathbf{y_n t+n-1}} + \cdots + \phi_{1n}(n-1)\boldsymbol{\varepsilon}_{\mathbf{y_n t+1}}\end{aligned}$$

Denote the variance of the *n*-step ahead forecast error variance of y_{1t+n} as $\sigma_{y_1}(n)^2$:

$$\begin{aligned}\sigma_{y_1}(n)^2 = {} & \sigma_{y_1}^2\left[\phi_{11}(0)^2 + \phi_{11}(1)^2 \cdots + \phi_{11}(n-1)^2\right] + \sigma_{y_2}^2\left[\phi_{12}(0)^2 + \phi_{12}(1)^2 \cdots + \phi_{12}(n-1)^2\right] \\ & + \sigma_{y_3}^2\left[\phi_{13}(0)^2 + \phi_{13}(1)^2 \cdots + \phi_{13}(n-1)^2\right] + \cdots + \sigma_{y_n}^2\left[\phi_{1n}(0)^2 + \phi_{1n}(1)^2 \cdots + \phi_{1n}(n-1)^2\right]\end{aligned}$$

Since all values of $\phi_{jk}(i)^2$ are necessarily nonnegative, the variance of the forecast error increases as the forecast horizon *n* increases. Note that it is possible to decompose the *n*-step ahead forecast error variance due to each one of the shocks. The proportions of $\sigma_{y_1}(n)^2$ due to shocks in the $\{\varepsilon_{y_1 t}\}, \{\varepsilon_{y_2 t}\}, \ldots, \{\varepsilon_{y_7 t}\}$ sequences are:

$$\frac{\sigma_{y_1}^2\left[\phi_{11}(0)^2 + \phi_{11}(1)^2 + \cdots + \phi_{11}(n-1)^2\right]}{\sigma_{y_1}(n)^2} \tag{28}$$

$$\frac{\sigma_{y_2}^2\left[\phi_{12}(0)^2 + \phi_{12}(1)^2 + \cdots + \phi_{12}(n-1)^2\right]}{\sigma_{y_1}(n)^2} \tag{29}$$

$\vdots$

$$\frac{\sigma_{y_7}^2\left[\phi_{1n}(0)^2 + \phi_{1n}(1)^2 + \cdots + \phi_{1n}(n-1)^2\right]}{\sigma_{y_1}(n)^2} \tag{30}$$

Equations (28), (29) and (30) are the forecast error variance decomposition (VDC), showing the proportion of the movements in a sequence due to its own shocks versus shocks to the other variable. If $\{\varepsilon_{y_2 t}\}, \{\varepsilon_{y_3 t}\}, \ldots, \{\varepsilon_{y_7 t}\}$ shocks explain none of the forecast error variance of $\{y_{1t}\}$ at all forecast horizons, we can say that the $\{y_{1t}\}$ sequence is exogenous. In applied research, it is typical for a variable to explain almost all its forecast error variance at short horizons and smaller proportions at longer horizons.

However, impulse response analysis and variance decompositions can be useful tools to examine the relationships among economic variables. If the correlations among the various innovations are small, the identification problem is not likely to be particularly important. The alternative orderings should yield similar impulse response and variance decompositions.

4 Empirical Results

4.1 Data Description

A total of nine time series datasets, including three energy prices and six macroeconomic variables, are applied in this study. The oil price (oil) data are collected from the West Texas Intermediate (WTI) crude oil spot price index in the commodity prices section. The gas price (gas) data are collected from the Russian Federation natural gas spot price index. The coal price (coal) data are collected from the Australia coal spot price index. Following Sadorsky (1999), we employ the six macroeconomic variables: industrial production index (ip), stock prices (sp), interest rate (r), unemployment rate (un), exports (ex) and imports (im). The industrial production index represents the level of output produced within an economy in a given year. In order to test for the impact in the labor market, the unemployment rate is chosen as a desirable proxy.

All data used in this study are monthly frequencies. Since the VAR or VECM model is used to estimate the non-linear relation, at least 200 data points are needed for a delay of 12 periods as suggested by Hamilton and Herrera (2004). The length of the available data is different and covers the period from 1975:M7-2008:M5 (oil price), 1979:M2-2008:M5 (coal price), and 1985:M1-2008:M5 (natural gas price). The energy price data are obtained from International Financial Statistics (IFS) CD-ROM. The macroeconomic variables are obtained from Taiwan Economic Journal (TEJ) and Advanced Retrieval Econometric Modeling System (AREMOS). All variables are deflated by the base year 2006 consumer price index (CPI) and a natural logarithm (except for interest rate and unemployment rate) is taken before conducting the analysis. Table 4.1 summarizes a description of all variables.

Table 4.1 Definitions of Variables

Variables	Definitions of variables	Source
oil	Logarithmic transformation of monthly real West Texas Intermediate crude oil spot price index in US dollar (in 2006 prices)	IFS (2008)
gas	Logarithmic transformation of monthly real Russian Federation natural gas spot price index in US dollar (in 2006 prices)	IFS (2008)
coal	Logarithmic transformation of monthly real Australia coal spot price index in US dollar (in 2006 prices)	IFS (2008)
ip	Logarithmic transformation of monthly real industrial production index in NT dollar (in 2006 prices)	TEJ
sp	Logarithmic transformation of monthly real stock prices in NT dollar (in 2006 prices)	TEJ
r	Monthly real interest rate	TEJ
un	Monthly unemployment rate	TEJ
ex	Logarithmic transformation of monthly real exports in NT dollars (in 2006 prices)	AREMOS
im	Logarithmic transformation of monthly real imports in millions NT dollar (in 2006 prices)	AREMOS

4.2 One-regime VAR analysis

4.2.1 Results of Tests for Unit Roots and Cointegration

An examination of Table 4.2 indicates our results are consistent, irrespective of using either the ADF unit root or KPSS unit root test. The statistic indicates that all of the individual series in first differences are stationary at the 1% significance level. This outcome suggests that all variables are integrated of order one or *I*(1). Thus, we use the differenced variables in the following analysis.

Table 4.2 Results of Unit Root Tests

Panel A. Oil price (1975:7-2008:5)				
	ADF		KPSS	
	Level	First differences	Level	First differences
oil	-0.989	-15.422***	0.402***	0.106
y	-0.357	-4.790***	2.408***	0.010
sp	-1.286	-18.248***	1.755***	0.080
r	-1.236	-16.639***	1.567***	0.048
un	-1.790	-4.334***	1.484***	0.126
ex	-2.157	-4.773***	0.292***	0.102
im	-0.524	-6.080***	2.359***	0.148
Panel B. Coal price (1979:2-2008:5)				
	ADF		KPSS	
	Level	First differences	Level	First differences
coal	-0.331	-14.577***	0.768***	0.455
y	-0.329	-4.507***	2.254***	0.014
sp	-1.375	-17.211***	1.431***	0.096
r	-0.990	-15.531***	1.548***	0.095
un	-2.209	-4.486***	1.417***	0.070
ex	-0.152	-4.709***	2.224***	0.055
im	0.048	-14.774***	2.261**	0.028
Panel C. Natural gas price (1985:1-2008:5)				
	ADF		KPSS	
	Level	First differences	Level	First differences
gas	-0.918	-6.374***	0.594**	0.446
y	-0.325	-4.273***	1.957***	0.020
sp	-2.798	-15.456***	0.489**	0.187
r	-1.291	-12.323***	1.255***	0.082
un	-1.606	-3.635***	1.358***	0.088

ex	0.048	-4.462***	1.929***	0.146
im	-1.013	-23.082***	1.920***	0.109

Note: '***' and '**' denote significance at 1% and 5%, respectively. Values in the parenthesis in ADF and KPSS unit root tests are *p*-values provided by Mackinoon (1996) and Kwiatkowski et al. (1992), respectively.

We apply the maximum eigenvalue and trace statistic proposed by Johansen (1988) to test the existence of a cointegration relation for these $I(1)$ variables. To determine the optimal lag length of the VAR model three versions of system are initially estimated: 2, 5, and 6-lag versions. A BIC is then employed to test that all three specifications are statistically equivalent. As shown in Table 4.3, there exist cointegration relations among variables. On the basis of the results the existence of a long-run relationship for all specifications finds statistical support in Taiwan over the period under examination.

Table 4.3 Results of the Johansen Cointegration Tests

Energy price: Oil price						
H_0	Eigenvalue	Trace	*p*-value	Eigenvalue	Max-Eigen	*p*-value
$r = 0$	0.15	153.33***	0.00	0.15	61.72***	0.00
$r \leq 1$	0.08	91.61	0.24	0.08	31.75	0.37
$r \leq 2$	0.05	59.86	0.48	0.05	20.41	0.79
Energy price: Coal price						
H_0	Eigenvalue	Trace	*p*-value	Eigenvalue	Max-Eigen	*p*-value
$r = 0$	0.17	133.06***	0.00	0.17	65.18***	0.00
$r \leq 1$	0.07	67.87	0.41	0.07	25.35	0.54
$r \leq 2$	0.05	42.52	0.59	0.05	18.12	0.69
Energy price: Natural gas price						
H_0	Eigenvalue	Trace	*p*-value	Eigenvalue	Max-Eigen	*p*-value
$r = 0$	0.15	124.22***	0.01	0.15	43.98**	0.04
$r \leq 1$	0.11	80.24	0.09	0.11	33.17	0.12
$r \leq 2$	0.10	47.07	0.38	0.10	28.30	0.09

Note: '***' and '**' denote significance at 1% and 5%, respectively.

4.2.2 Results of the Variance Decomposition

Table 4.4 presents the variance decomposition results based on the VECM model for energy price. Each percentage shows how much of the unanticipated changes in macroeconomic variables are explained by the energy price variable over a 12-month horizon. The industrial production variable's own shocks account for 77.59% to 93.95% of the forecast variance. After a year (12 months), oil prices, stock prices, interest rate, unemployment rate, exports and imports account for 1.94%, 2.72%, 1.95%, 7.55%, 5.33% and 2.91% of the industrial production forecast error variance, respectively. Compared to the other energy prices (i.e., coal price and natural gas price), oil price changes in Taiwan have the largest explanatory effect for industrial production.

After one year (12 months), 89.87% of the stock price variability is attributed to changes in itself, 1.54% to oil price changes, 2.22% to interest rate changes, and 2.75% to unemployment rate changes. Moreover, coal price changes explain a 0.22% change in stock prices, slightly lower than 0.91% explained by the interest rate. Natural gas price shocks are important driving forces behind stock price variability, explaining almost 3.53% of the variation in stock prices in the short term (about a year). The interest rate variable's own shocks account for most of the forecast error variance. The oil price change explains about 3.88% of the interest rate change (greater than 1.64% explained by stock price change). Compared to other energy prices, a natural gas price change has stronger explanatory power on the interest rate.

For the unemployment variable, the unemployment variations are still mainly due to its own changes of about 62.77% - 81.18%, while approximately 2.52% is attributed to oil price changes, 0.40% to coal price changes, and 2.66% to natural gas price changes. In other words, a natural gas price change has stronger explanatory power on unemployment.

Table 4.4 Variance Decompositions of Forecast Error Variance in One-regime VAR Model (12 Periods Forward)

	Shock sources						
	ε^{y}	ε^{ep}	ε^{sp}	ε^{r}	ε^{un}	ε^{ex}	ε^{im}
Panel A. Energy price: Oil price							
y	77.59	1.94	2.72	1.95	7.55	5.33	2.91
ep	1.99	90.32	2.22	0.96	1.00	1.97	1.54
sp	1.09	1.54	89.87	2.22	2.75	0.90	1.62
r	1.50	3.88	1.64	84.49	4.02	2.48	1.99
un	20.27	2.52	1.97	2.04	66.60	3.29	3.31
ex	28.61	2.63	1.90	0.51	2.26	60.10	4.00
im	26.95	2.09	2.96	0.86	1.51	23.16	42.46
Panel B. Energy price: Coal price							
y	93.95	0.52	0.14	0.08	0.27	4.97	0.09
ep	0.57	93.97	1.60	0.04	1.67	0.27	1.88
sp	0.26	0.22	95.98	0.91	1.61	0.75	0.27
r	2.33	0.96	0.86	92.80	1.48	0.35	1.22
un	14.45	0.40	0.96	0.53	81.18	0.76	1.73
ex	47.60	1.57	1.05	0.28	0.64	47.72	1.15
im	40.29	1.45	0.45	0.59	1.27	13.05	42.90
Panel C. Energy price: Natural gas price							
y	81.02	0.86	2.43	2.22	2.99	8.07	2.41
ep	1.43	89.22	1.41	2.35	0.83	0.96	3.80
sp	0.80	3.53	86.41	1.58	2.79	2.56	2.38
r	1.66	1.18	1.12	86.88	4.86	2.48	1.83
un	19.98	2.66	3.58	1.22	62.77	4.48	5.31
ex	46.46	2.54	1.89	1.35	0.85	39.33	7.59
im	42.29	3.02	2.23	1.31	0.52	14.99	35.63

Note: Values in the parenthesis are standard errors estimated through 500 Monte Carlo replications. Variance decomposition explaining the variation in variables is due to industrial production shocks (ε^{y}), energy prices shocks (ε^{ep}), stock price shocks (ε^{sp}), interest rate shocks (ε^{r}), unemployment rate shocks (ε^{un}), export shocks (ε^{ex}), and import shocks (ε^{im}).

4.2.3 Results of the Impulse Response Analysis

The impulse response functions illustrate the qualitative response of the variables in the system to shocks in energy prices. Figure 4.1 presents the impulse response functions of each oil price change (DWTI), coal price change (DCOAL), and natural gas price change (DNG) from one-standard deviation shocks to industrial production (DIP), stock price (DSP), real interest rates (DR), unemployment rates (DUN), exports (DEX) and imports (DIM) in the one-regime VAR model.

An oil price shock has a negative impact on industrial production. It responds negatively in period 2, and its responses exhibit more volatility. An oil price shock has a delayed negative impact on industrial production.

Following an oil price shock, stock prices decrease immediately by 0.3%, showing that an oil price shock has a negative impact on stock prices. An oil price shock has a positive impact on interest rates. The phenomenon sustains itself for approximately 12 periods. The maximum effect is reached after 4-5 periods, when the interest rate has increased by 4%. An oil price shock has a positive impact on the unemployment rate which increases by 2% in period 2.

The industrial production reacts a negatively and significantly to a coal price change in the first period. A negative response from stock prices is observed in period 2, but the effect is small and not significantly different from zero. A coal price shock always keeps a positive impact on the unemployment rate that lasts for approximately 6 periods. Following a coal price shock, both exports and imports decrease immediately by 0.5% in period 1 and then increase by 0.5% in period 2.

The industrial production decreased in periods 2 and 7 from a natural price shock. A natural gas price shock has a delayed negative impact on stock prices. With a delayed stock price response in Taiwan, stock prices at first rise and then decrease, this lasts for approximately 5-6 periods. The unemployment rate only starts to decrease in period 1 and it exhibits an upward inclination pattern in period 3.

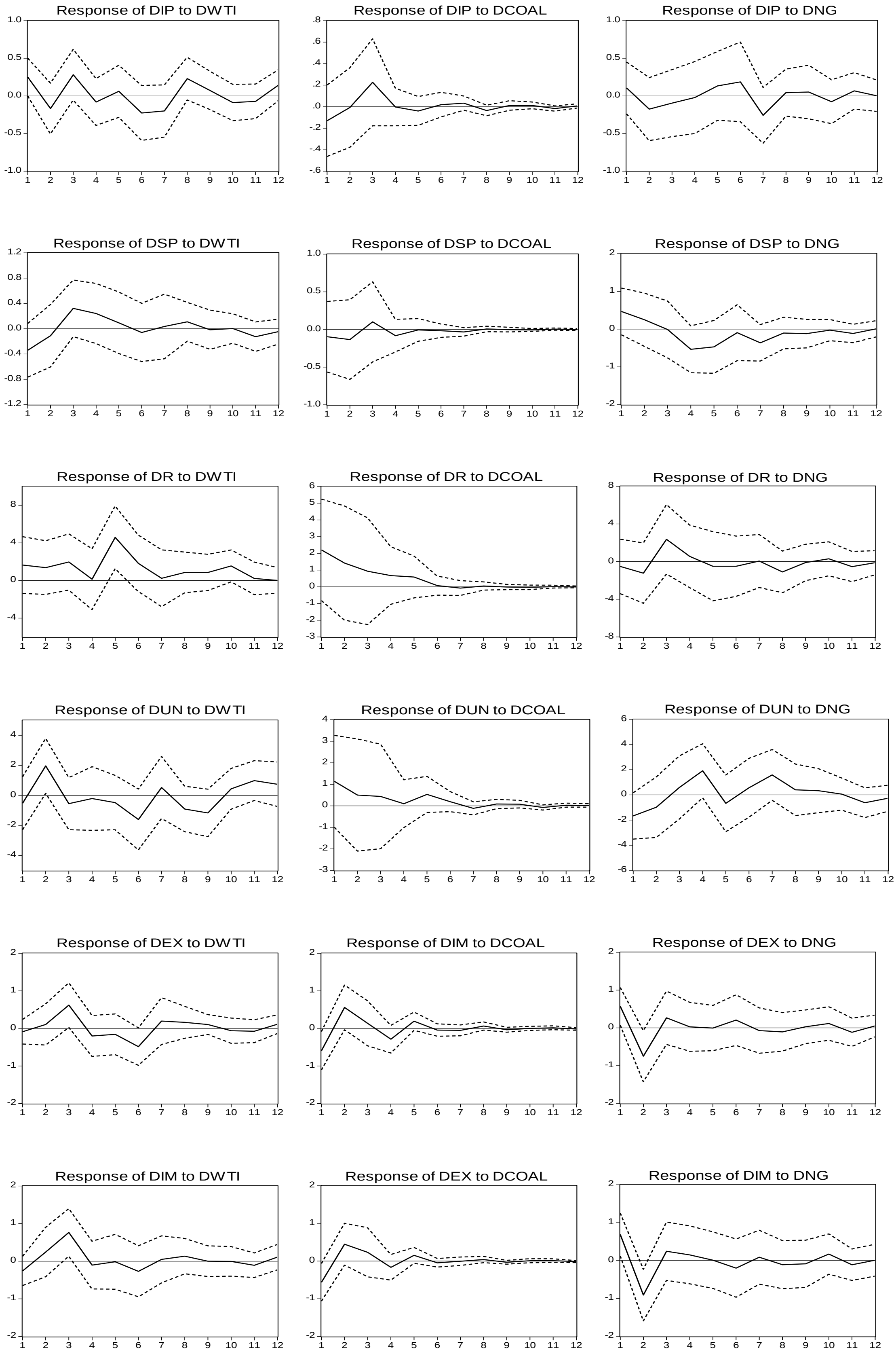

Figure 4.1 Impulse Responses from One Standard Deviation Shock of Energy Price Change in the Linear VAR Model (12 Periods Forward)

Each macroeconomic variable should exhibit significant responses when energy price changes are modest or more. That is to say, we need to offer more detailed responses. To overcome the problem, we use the multivariate threshold error correction (MVTEC) model developed by Tsay (1998).

4.3 Two-regime VAR analysis

4.3.1 Estimating the Threshold Levels and the Delay of Threshold Variables

Before employing the MVTEC model, it is necessary to test the existence of the non-linear relationship in terms of the threshold variables (i.e., oil price, coal price, and natural gas price) during these two periods. The *C(d)* statistic based on the arranged regression (Tsay, 1998) can be used to test the linear relation. Table 4.5 displays the tests results.

Table 4.5 Results of Threshold Effect Tests

Threshold variable	Delay (*d*)	C(*d*) *p*_value	Threshold value (c^*)	Regime one	Regime two
Oil	1	354.36 (0.04)	2.48%	326	68
Coal	1	180.52 (0.00)	0.22%	223	128
Natural Gas	1	397.87 (0.00)	0.87%	239	41

Note: Regime one refers to $Z_{t-d} \leq c$ and Regime two $Z_{t-d} \geq c$. c^* is the optimal threshold value determined by the location of the minimum log det|∑|, and ∑ is the variance–covariance matrix for the corresponding multivariate VECM models.

Based on Table 4.5, we can reject the null hypothesis of the linear model using oil price as a threshold variable. It means that the result favors the MVTEC model. By the same token, we find a similar result when coal price or natural gas price is used as a threshold variable. There is a non-linear relationship between energy price and industrial production. The delay (d) of the threshold variable reflects the speed of response based on the economic impact of a positive energy price change and its shock. Results about the length of delay are similar to Huang's (2008) result:

when a country has a higher energy import ratio, it will have a shorter delay in terms of its economic response from the positive impact of an energy price change. In this study the impact of energy price changes (i.e., oil price, coal price, and natural gas price) on production is rapid (one month).

The threshold value (c) reflects the critical level of the impact. In order to estimate the optimal threshold value c^*, we search procedure targets at the middle 60% to 80% of the arranged dataset. From the estimated results of threshold effect tests in Table 4.5, the optimal threshold levels c are as follows: the highest level is oil price at 2.48%, the next highest is natural gas price at 0.87%, and the lowest level is coal price at 0.22%. When the oil price change (1 month before) exceeds 2.48%, its impact is significantly different from that if the price change is less than 2.48%. Similarly, when the coal price change (1 month before) exceeds 0.22%, its impact is significantly different from that if the price change is less than 0.22%. The impact on industrial production is different from when the natural gas price change (1 month before) exceeds 0.87%.

4.3.2 Results of the Variance Decomposition in the MVTEC Model

In order to depict the response of macroeconomic activities in regime one (energy price changes are less than or equal to c^*) and regime two (energy price changes exceed c^*), we employ the VDC and the IRF analyses for each regime. Table 4.6 reports proportions of impacts emanating from an oil price change in terms of VDC. When an oil price change is below the threshold value c^*, an oil price change explains about 7.12% of industrial production change (greater than 1.98% explained by an interest rate change). Similarly, the proportions of the explanatory power of oil price and interest rate on unemployment change are roughly the same: 2.09% vs. 2.06%. On the other hand, when an oil price change exceeds the threshold value c^*, it explains much more on industrial production than does the interest rate (17.22% vs. 4.39%). However, oil price changes explain less significantly on stock prices (7.45% vs. 14.85%) and unemployment (6.87% vs. 14.11%) than the interest rate.

Table 4.6 Variance Decomposition Results Using Oil Price Changes as Threshold Variable (12 Periods Forward)

	Shock sources						
	ε^y	ε^{ep}	ε^{sp}	ε^r	ε^{un}	ε^{ex}	ε^{im}
Panel A. Regime one							
y	79.60	7.12	3.53	1.98	3.38	1.62	2.77
ep	2.16	89.71	2.62	0.65	2.05	1.23	1.59
sp	0.81	1.01	90.38	1.38	3.57	0.96	1.90
r	2.97	1.09	4.63	79.64	4.78	4.67	2.21
un	13.62	2.09	2.50	2.06	73.74	0.89	5.09
ex	34.81	2.98	2.08	0.39	1.00	57.53	1.22
im	28.36	4.12	3.68	0.68	3.63	22.72	36.79
Panel B. Regime two							
y	51.86	17.22	7.79	4.39	3.61	8.14	6.99
ep	4.85	30.93	4.63	17.89	4.88	18.28	18.53
sp	18.11	7.45	34.99	14.85	10.98	4.74	8.88
r	9.91	17.84	8.01	33.91	6.59	9.55	14.18
un	13.12	6.87	10.17	14.11	37.27	9.83	8.64
ex	34.28	15.60	12.75	6.50	10.09	16.67	4.13
im	36.28	21.16	13.75	6.68	5.51	7.70	8.93

Note: Regime one pertains to $Zt\text{-}d \leq c^*$ while regime two pertains to $Zt\text{-}d > c^*$.

Table 4.7 illustrates the impact of coal price changes on macroeconomic variables in terms of the VDC. When a coal price change is below the threshold value, it explains a significant portion of change in industrial production (4.11%), unemployment (3.19%), exports (4.29%), and imports (5.82%). When a coal price change exceeds the threshold value c^* (regime two), the coal price change explains about 3.88% of the stock price change (greater than 2.23% explained by the interest rate change). Within the linear model, a coal price change exerts less significant impact than an interest rate change (0.22% vs. 0.91%). This is strikingly different from the results of the two -regime model, which displays significant responses from stock markets when the coal price change is modest or more.

Table 4.7 Variance Decomposition Results Using Coal Price Changes as Threshold Variable (12 Periods Forward)

	Shock sources						
	ε^{y}	ε^{ep}	ε^{sp}	ε^{r}	ε^{un}	ε^{ex}	ε^{im}
Panel A. Regime one							
y	87.78	4.11	3.53	0.26	0.69	2.30	1.32
ep	0.98	91.53	0.14	0.87	1.23	4.87	0.38
sp	0.37	0.53	93.26	0.96	2.22	1.61	1.05
r	2.85	0.93	0.65	91.45	3.18	0.42	0.52
un	11.45	3.19	0.53	0.40	81.65	2.30	0.50
ex	37.91	4.29	0.83	0.37	0.69	54.55	1.36
im	34.42	5.82	0.79	1.04	1.93	17.74	38.26
Panel B. Regime two							
y	82.71	4.20	1.46	3.44	1.03	3.85	3.32
ep	2.66	83.08	2.30	1.11	7.34	1.66	1.86
sp	4.28	3.88	77.08	2.23	4.49	2.90	5.15
r	3.55	4.60	5.78	75.03	3.27	3.93	3.84
un	6.50	2.75	4.44	4.01	76.10	4.64	1.56
ex	31.57	8.02	6.20	7.42	3.64	36.98	6.18
im	23.01	10.22	8.80	3.60	2.81	21.76	29.80

Note: Regime one pertains to $Zt\text{-}d \leq c^{*}$ while regime two pertains to $Zt\text{-}d > c^{*}$.

Table 4.8 presents the VDC results from natural gas price changes. As can be seen from the table, a natural gas price change has significant explanatory power. It indicates that a natural gas price change (1) explains more on industrial production change than does an interest rate change (6.49% vs. 1.42%); (2) accounts more on stock price change than does an interest rate change (3.89% vs. 1.91%); (3) accounts more on unemployment in comparison to the interest rate (6.06% vs. 3.17%) and (4) explains more on exports (4.26%) and imports (6.20%) than other macroeconomic variables. In regime two, it indicates that a natural gas price change can explain more on unemployment than does an interest rate change (25.15% vs. 4.68%). Furthermore, a natural gas price change has significant explanatory power on exports (11.42% vs. 4.26%) and imports (18.16% vs. 6.20%) in comparison to regime one. In particular, the explanatory power of a natural gas price change is greater than the interest rate under two regimes. This result is consistent with findings by Park and Ratti (2008) in that the contributions from energy price shocks are greater than that of interest rates on the stock market.

Table 4.8 Variance Decomposition Results Using Natural Gas Price Change as Threshold Variable (12 Periods Forward)

	Shock sources						
	ε^{y}	ε^{ep}	ε^{sp}	ε^{r}	ε^{un}	ε^{ex}	ε^{im}
Panel A. Regime one							
y	76.64	6.49	4.71	1.42	3.54	4.95	2.26
ep	2.58	84.70	0.92	1.76	2.15	2.16	5.73
sp	0.47	3.89	80.22	1.91	3.37	8.11	2.03
r	1.27	1.12	2.79	85.42	5.21	2.66	1.52
un	11.82	6.06	2.34	3.17	63.82	7.36	5.43
ex	42.72	4.26	2.74	1.67	1.89	43.08	3.64
im	35.74	6.20	1.90	5.15	1.78	21.40	27.83
Panel B. Regime two							
y	73.87	4.82	3.56	13.18	1.65	0.65	2.27
ep	10.96	52.03	9.11	8.82	3.14	9.50	6.44
sp	4.37	5.37	77.60	7.09	1.86	3.01	0.70
r	14.19	8.84	18.19	47.99	7.60	2.17	1.02
un	6.51	25.15	16.96	4.68	37.65	2.48	6.57
ex	23.81	11.42	25.43	3.49	2.35	31.38	2.12
im	30.26	18.16	24.66	5.25	1.07	10.60	9.99

Note: Regime one pertains to $Zt\text{-}d \leq c^{*}$ while regime two pertains to $Zt\text{-}d > c^{*}$.

4.3.3 Results of the Impulse Response Analysis in the MVTEC Model

The left of Figure 4.2 presents the impulse responses of macroeconomic variables to an oil price shock in regime one. When an oil price change is below the threshold value c^* (regime one), it can be seen that an oil price shock has a positive impact on industrial production. The response of industrial production to oil price shocks is rising in periods 1 and 3. An oil price shock has a persistently negative impact on stock prices over 11 periods. The oil price shock has an immediate positive response in the interest rate, and then falls. This result can be expected as increases in oil price create inflationary effects in the economy which consequently bring an upward pressure on interest rates. The results for the unemployment rate are somewhat stronger. Except for the first one minor negative response, the graph shows persistent positive responses of unemployment to a shock in oil price. The maximum effect is reached in the second period when the unemployment rate increases by 2%.

The left of Figure 4.3 presents the impulse responses of macroeconomic variables to an oil price shock in regime two. When an oil price change exceeds the threshold value c^* (regime two), an oil price shock has an immediate positive impact on industrial production, and after a minor negative shock it tends to remain for a significant period of time. The IRF analysis shows that oil price shocks exhibit more volatility in the one-regime model than in the two-regime model. An oil price shock has an initial minor negative impact on interest rates and then the increase lasts for approximately 12 periods. The responses of the unemployment rate are only after an initial slight positive impact and then fall.

When a coal price change is below the threshold value c^* (regime one), it can be observed that a positive coal price shock increases industrial production in period 3. A coal price shock has a slight negative impact on stock prices that lasts for approximately 4 periods. In particular, the unemployment rate initial rise lasts for

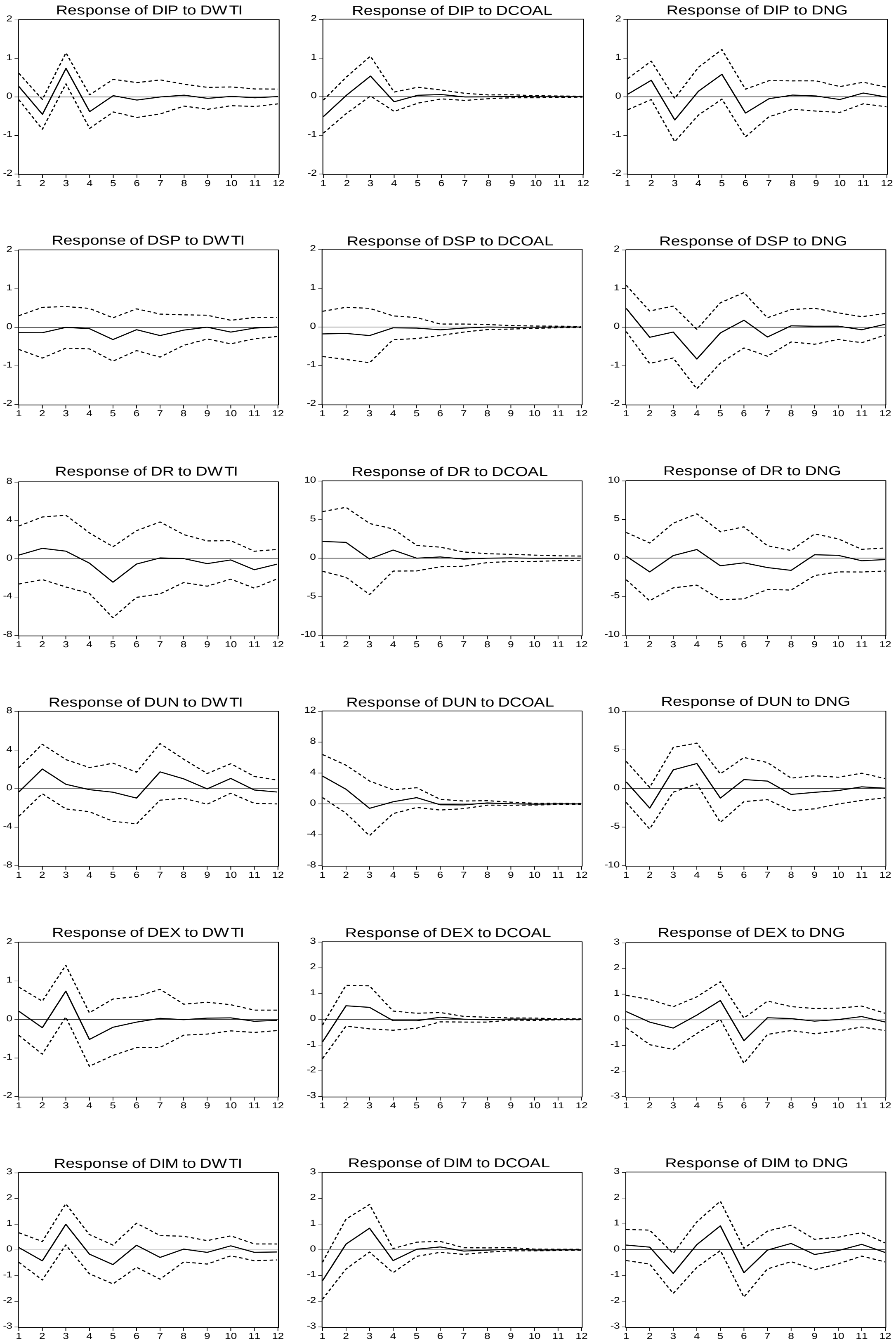

Figure 4.2 Impulse Responses from One Standard Deviation Shock of Energy Price Change in the Regime One VAR Model (12 Periods Forward)

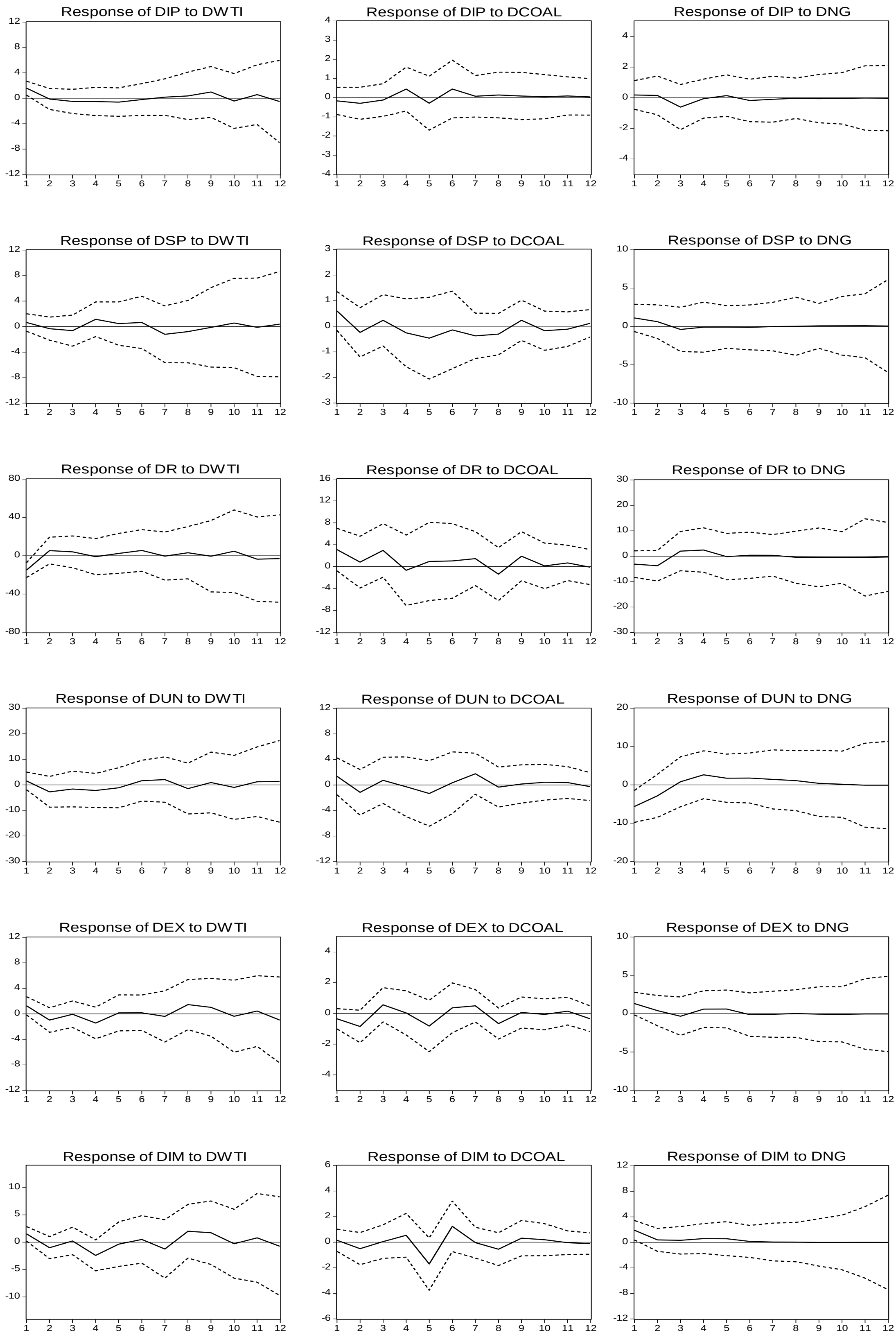

Figure 4.3 Impulse Responses from One Standard Deviation Shock of Energy Price Change in the Regime Two VAR Model (12 Periods Forward)

approximately 2 periods and then decreases in period 3. The results show that following a 10% increase in coal price, the unemployment rate increases immediately by 3–4%. The maximum effect is reached in the first period, after which the effect gradually dies out.

When a coal price change exceeds the threshold value c^* (regime two), a coal price shock has a slight negative impact on industrial production that lasts for approximately 3 periods. In periods 4 and 6, the response of industrial production to shocks in coal price is increasing and then the effect dies out. As expected, a coal price shock has a negative impact on stock prices. The graph presents that the response of stock prices to shocks in coal price is positive up to the first period and it eventually declines. A coal price shock has a negative impact on the interest rate expected for periods 4 and 8.

When a natural gas price change is below the threshold value c^* (regime one), a natural gas price shock has a significant positive impact on industrial production. The industrial production initial rise lasts for approximately 2 periods and increases in periods 4-5. The figure also shows that following a natural gas price shock, stock prices increase immediately by 0.5%. After that, stock prices decrease persistently for 2-4 periods. The maximum effect is reached in period 4, when stock prices have decreased by 1%. The responses exhibit more volatility and last for a long term.

When a natural gas price change exceeds the threshold value c^* (regime two), a natural gas price shock has a small negative impact on the interest rate. After 3-4 periods, the effect gradually dies out. As expected, stock prices initially rise, probably affected by the rise in economic activity, and then die out in period 3.

4.3.4 Results of the Parameter Stability Tests

We use the Pesaran and Pesaran (1997) tests for general parameter stability.

They suggest applying the cumulative sum of recursive residuals (CUSUM) and the CUSUM of square (CUSUMSQ) tests proposed by Brown et al. (1975) to assess the parameter constancy. The CUSUM and CUSUMSQ tests both plot the cumulative sum together with the 5% critical lines to find parameter instability if the cumulative sum goes outside the area between the two critical lines. Assuming there are k parameters in the model, the CUSUM test is based on the statistic:

$$W_t = \sum_{r=k+1}^{t} \omega_r / s \quad \text{for} \quad t = k+1, \ldots, T, \tag{31}$$

where ω_r is the recursive residual and s is the standard error of the regression fitted to all T sample points. The significance of any departure from the zero line is assessed by reference to a pair of 5% significance lines, and the distance between which increases with t. The CUSUM of squares test is:

$$W_t = \left(\sum_{r=k+1}^{t} \omega_r^2 \right) \Bigg/ \left(\sum_{r=k+1}^{T} \omega_r^2 \right) \tag{32}$$

The expected value of s_t under the hypothesis of parameter constancy is

$$E(s_t) = (t-k)/(T-k),$$

which goes from zero at $t=k$ to unity at $t=T$. The significance of the departure from its expected value is assessed by reference to a pair of parallel straight lines around the expected value.

Figure 4.4 plots the CUSUM and CUSUMSQ statistics when energy price is the dependent variable and energy price changes are less than or equal to c^* (regime one). The results indicate no instability in the coefficients as the plots of the CUSUM and CUSUMSQ statistics are confined within the 5% critical bounds of parameter stability. On the other hand, when energy price changes exceed the threshold value c^* (regime two), the graphical representations of the tests are plotted in Figure 4.5. Both the CUSUM and the CUSUMSQ plots are confined within the 5% critical bounds, suggesting that the residual variance is somewhat stable over time. In other words, if there is a structural break, then they will tend to drift above the bounding lines at the 5% level of significance. As shown in Figure 4.4, both tests suggest that the null hypothesis of the absence of a structural break cannot be rejected at the 5% level of significance. Thus, the models are stable over time. It appears that applying two-regime error correction models does not suffer from any problem caused by a structural break. Similar conclusions can be found from Figure 4.5.

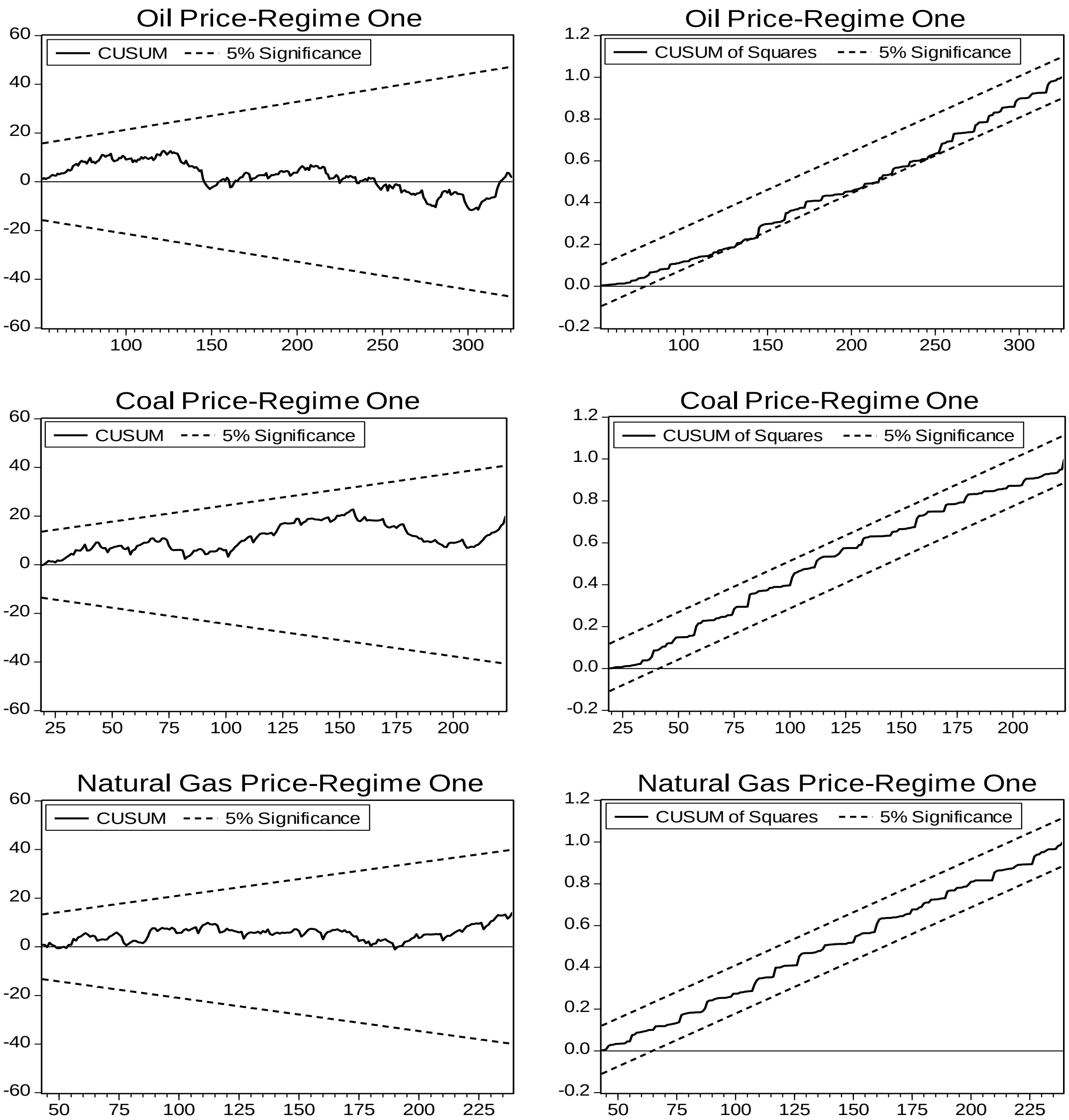

Figure 4.4 Plots of the CUSUM and CUSUM of Square Tests in Regime One

Note: Values in the vertical axis refer to the CUSUM statistic and in horizontal axis represent the time point in *t* of regime one.

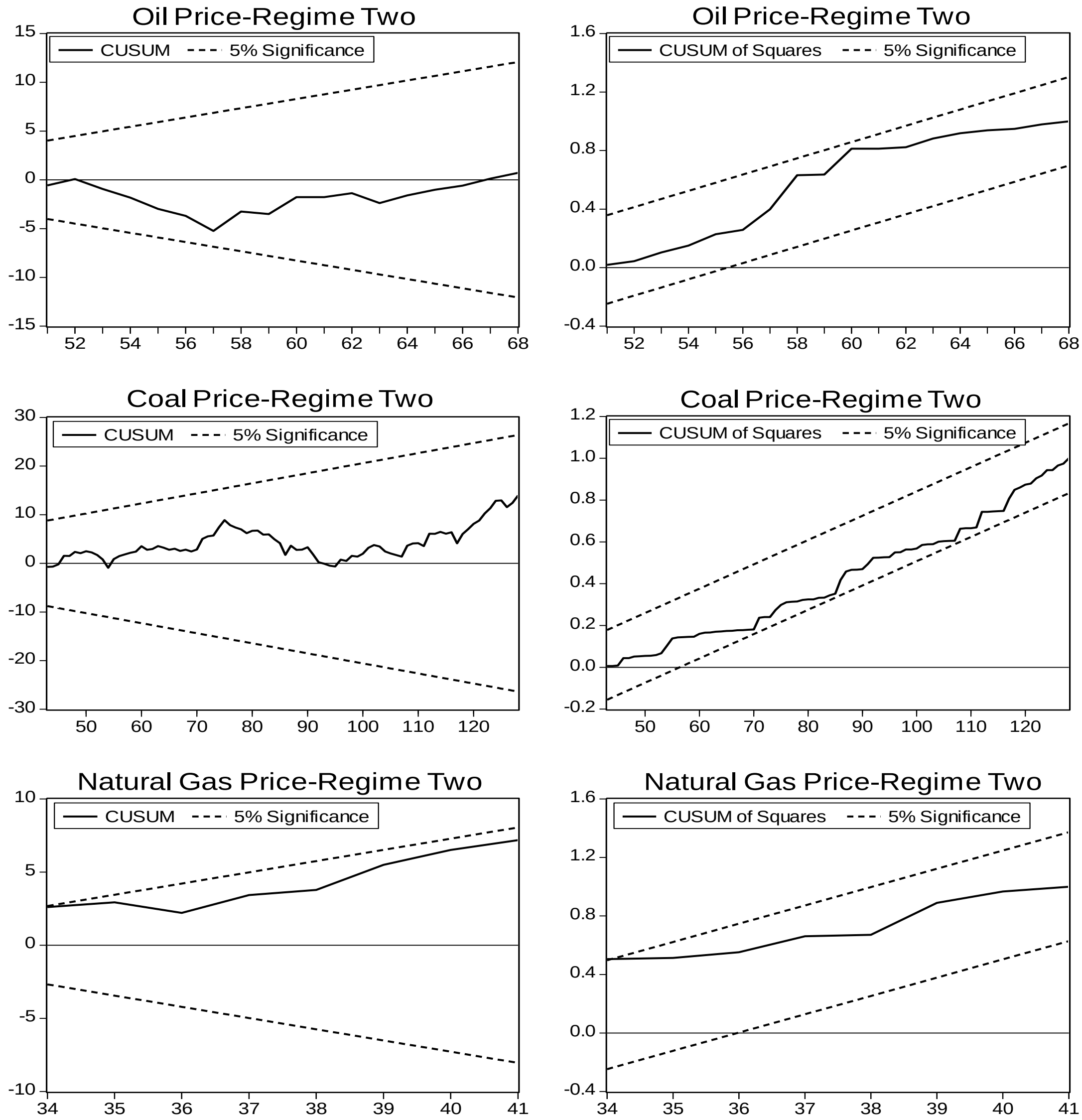

Figure 4.5 Plots of the CUSUM and CUSUM of Square Tests in Regime Two

Note: Values in the vertical axis refer to the CUSUM statistic and in horizontal axis represent the time point in *t* of regime two.

5 Preliminary Conclusions and Policy Implications

The main purpose of this paper explores the effects of international energy price shocks and macroeconomic activity in Taiwan. The preliminary findings are: (1) There is a threshold non-linearity relationship between energy price variables and macroeconomic variables. (2) The optimal threshold levels are 2.48% in terms of oil price change, 0.87% in terms of natural gas price change, and 0.22% in terms of coal price change. Due to Taiwan's higher economic development, the threshold of critical level is greater as evidence by the positive impact of an oil price change and its shock. The optimal threshold value seems to vary according to how an economy depends on imported energy and the attitude towards accepting energy-saving technology. (2) If a country has a higher energy import ratio and acquires a higher ratio of energy use in the industrial sector, then it will have a shorter delay in terms of its economic response from the positive impact of an energy price change. As our results show, the delays of the threshold variable are only one month and their responses are very quick. (3) Compared to the other energy prices (i.e., coal price and natural gas price), an oil price change has the largest explanatory effect on Taiwan's industrial production. Moreover, it better explains industrial production than the real interest rate when an oil price change exceeds the threshold value (regime two). (4) A coal price change significantly explains stock prices in the two-regime model compared to the one-regime model. A natural gas price change has higher explanatory power on stock prices than the interest rate when a natural gas price change is below the threshold value (regime one). In a similar vein, a natural gas price change has stronger explanatory power on the unemployment rate. (5) Energy price shocks have a negative impact on Taiwan's macroeconomic activities especially in industrial production and stock prices in regime two. Both oil price shocks and natural gas shocks have a delayed negative impact on industrial production with one lag when energy price changes exceed the threshold level. By the same token, energy price shocks have delayed negative impacts on the stock market. (6) To Taiwan's labor market, international energy price shocks have a positive effect on the unemployment rate in the short term. It means that an increase in energy prices will increase the cost of production which in turn results in higher levels of unemployment. (7) In summary, the findings speak to the fact that the two-regime model seems to offer more detailed and noticeable responses.

Based on the aforementioned findings, we observe that energy prices have significant impacts on Taiwan's macroeconomic activity. In order to reduce the impact of energy price shocks and promote sustainable development in Taiwan, we further address the trend of Taiwan's energy development and some energy strategies for domestic policy makers. The first one is to actively develop the domestic renewable (or green) industry. The second one is to promote greater scale efficiency and to obtain competitive advantages for the domestic energy industry. The final one is to achieve energy technological breakthroughs.

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國立交通大學
National Chiao Tung University

類別：學術研討會

題目：兩岸博鰲學術會議與企業實務論壇

服務機關：國立交通大學經營管理研究所
姓名職稱：胡均立 教授兼所長
前往國家：中國大陸海南省博鰲
出國期間：自 99 年 4 月 1 日至 99 年 4 月 5 日
報告日期：99 年 4 月 26 日

一、參加經過

赴中國大陸參加學術研討會，作第一手觀察與學習，跟大陸的學者交換意見，對於長期從事中國大陸經濟研究的敝人，幫助頗大。此次係赴海南省博鰲參加兩岸學術會議。敝人是在 4 月 1 日跟隨成功大學管理學院教授及 EMBA 學生團體從高雄機場飛到澳門機場，從澳門過境經陸路進入中國大陸珠海市，再從珠海機場飛海口機場。並於 4 月 5 日上午由海口機場飛珠海機場，從珠海市經陸路過境澳門，再由澳門機場飛返高雄機場。正式會議

期間為 4 月 2 日至 4 月 4 日。4 月 2 日為報到，4 月 3 日舉行開幕式及論壇演講，4 月 4 日舉行分組論文發表。4 月 3 日的主題演講議題為成功大學 EMBA 執行長談台南科學園區的發展經驗、國電大廈垻發電有限公司總經理陳大兵的實務分享等。

二、心得

敝人係於 4 月 4 日上午發表「A Comparative Efficiency Study of Life Insurance Companies in Mainland China and Taiwan」一文，與南開大學支燕副教授合著。主要以海峽兩岸的人身保險公司為分析對象，樣本期間自 2002 年到 2007 年共六年期間。台灣地區共選取 28 家壽險公司，大陸地區共選取 41 家壽險公司為研究對象，共計 69 家壽險業者作為決策單位（DMU）。投入變數包含負債資本、權益資本及員工人數，產出項包含保費收入及投資收入。數據分別來自於《台灣保險年鑑》（2002-2007）及《中國保險年鑑》（2003-2008）。

研究方法採取 Fried et al. (2002) 的三階段方法。第一階段先以資料包絡分析法逐年求解個別壽險公司的效率值及投入差額變數。第二階段再以隨機邊界成本函數估計影響投入差額變數的方程式。對投入項進行排除環境效果及統計噪音的調整。第三階段再以經環境效果及統計噪音調整後的投入項，進行資料包絡分析法計算，得到排除環境效果及統計噪音影響後的效率值。

實證結果發現：第一階段之實證結果發現 2002 年至 2006 年間，兩岸壽險公司之平均效率值差異不大，表現較佳之公司皆為年輕且資產規模較小之壽險公司。第二階段實證結果發現無論總體經濟層面或個體企業層面之環境因素皆對壽險公司之經營效率有顯著之影響。此外，除了權益資本的無效率是顯著來自於管理無效率外，負債資本與員工人數的無效率則多來自運氣的好壞(統計噪音項)。第三階段實證結果發現經由環境及隨機干擾項調整過後的兩岸壽險業整體平均效率皆有明顯的下降。因此在研究效率值時，應將環境因素及隨機干擾的影響加以考慮並扣除，才能得到真正的效率值。而資產規模較大之壽險公司經營效率較佳。

提升兩岸壽險業效率需從總體之經濟環境及個別企業管理兩個方面同時改進：

1. 提升經濟自由度並引導消費，降低儲蓄。使企業依據市場規律自由競爭，在經濟發展的同時擴充社會可支配資金量。
2. 引導壽險公司發展成適度規模。當壽險企業置於同一環境狀態下，規模大的壽險企業效

率值比較高。但當規模成長至一定的程度時，須注意因規模過度擴大而造成資本及人力投入的無效率。

3. 就企業股東性質與地區別而言，外商企業、合資企業、大陸或臺灣企業在資源投入效率上各有所長。

敝人發表論文的評論人為海南大學嚴傳東教授。嚴教授的主要建議為計量模型的實證結果合理，但宜加強對兩岸壽險市場背景的描述，並強調制度差異對兩岸壽險業者經營效率的影響。

敝人同時擔任海南大學嚴傳東教授發表的「碳權性質初探」論文的評論人，敝人以參與台灣環保署碳權市場建制的研究計畫為例，建議可由聯合國清潔發展機制、兩岸碳權議題合合作，進一步發展這篇文章。

三、考察參觀活動

成大管理學院團隊於 4 月 4 日下午參觀了博鰲亞洲論壇的永久會址，今年的大會主題是綠色復甦，於 4 月 9-11 日召開。中國大陸於 2000 年首度召開博鰲亞洲論壇，邀請世界各國領袖與企業家前來與會，以促進區域對話的方式增進對區域的政治及經濟影響力。

四、建議

在兩岸經貿往來日漸深化之際，兩岸企業經營效率的比較研究，有助於讓我們提前掌握兩岸企業在市場中的競爭優勢與可能的均衡結果。兩岸企業經營效率的研究應建立在事實（facts）而非意見（opinion）之上。此外，台灣經濟與管理學門的英文寫作能力及計量分析訓練仍領先於大陸學者，若能配合大陸經濟發展的統計數據及制度演化，在優良類國際期刊上發表的機會大增，並能對兩岸提供經濟發展與企業管理上具有理論高度的建議。

五、攜回資料名稱及內容

議程及會議論文光碟。

六、其他

無。

類別：國際學術會議

題目：International Conference on Business and Information 2010

服務機關：國立交通大學經營管理研究所
姓名職稱：胡均立 教授兼所長
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出國期間：自 99 年 7 月 4 日至 99 年 7 月 9 日
報告日期：99 年 7 月 31 日

一、參加經過

近年敝人將經濟分析應用於管理議題的研究上，包含銀行、旅館、電子電機、醫院等產業的經營績效之研究。2008 年曾經參加過在韓國首爾舉行的 International Conference on Business and Information (BAI)，主辦單位之一為國立台北大學商學院，跟各國與會者互動良好。今年亦報名參加在日本北九州市舉行的 BAI 國際學術會議，除了發表論文以外，並順道訪問在日本九州產業大學任教的共同著作人本間聰（Satoshi Honma）教授，並當面研

擬未來一年的學術合作議題。值得一提的是，Honma 教授與敝人合著的「Analyzing Japanese hotel efficiency: a DEA application」一文榮獲本屆大會的最佳論文獎（Best Paper Award）。敝人是在 7 月 4 日下午跟隨大同大學教授及研究生團體從桃園機場飛到福岡機場，並於 7 月 9 日上午由福岡機場飛回桃園機場。會議期間為 7 月 5 日至 7 月 7 日。

二、心得

敝人係於 7 月 7 日上午發表「R&D Efficiency and National Innovation System: An International Comparison Using the Distance Function Approach」一文，與中央大學楊志海教授及陳疆平博士候選人合著。這篇論文以超越對數距離函數方法，比較 24 個國家的研發效率，樣本期間為 1998 年至 2005 年。計量模型中有兩個投入項，分別為研發資本量及研發人力；有三個產出項，分別為專利數、權利金及授權金收入及科學期刊論文篇數。經過資料處理及移項，這篇論文利用 Battese and Coelli (1995) 的隨機邊界法（stochastic frontier analysis; SFA）一階段來聯立估計多產出、多投入的超越對數距離函數及無效率方程式。實證結果發現：智慧財產權保護、人力資本累積、企業部門間的技術合作、企業部門與高等教育機構間的技術

移轉顯著改善各國的研發效率，但是政府的研發支出對於研發效率並無顯著影響。與會者普遍認為這篇研究方法嚴謹且結論大致合理，但文獻定位仍應予以比較釐清。

本間聰教授則於 7 月 7 日下午發表與敝人合著的「Analyzing Japanese hotel efficiency: a DEA application」一文，並擔任該場次的會議主席。如前所述，這篇文章獲得本次會議的最佳論文獎。這篇文章分析日本主要旅館業者 2004 年至 2008 年的經營效率。我們採用資料包絡分析法（data envelopment analysis; DEA）來計算個別旅館的效率值。模型中有 4 項投入（全職員工人數、臨時員工人數、餐廳及酒吧座位數、客房數目）及 1 項產出（總收益）。實證結果發現：主要的 15 家日本旅館業者的平均整體效率值達到 80.3%。一半以上的日本主要旅館業者處在規模報酬遞增階段。同時在數個城市營運的日本旅館業者顯著享有規模效率上的優勢。敝人於會後也立即告知 Honma 教授一些寫作上的不夠精確的描述，在我們投稿國際期刊前必須予以修正。

敝人近年的效率與生產力研究將研究對象由過去的能源、污染、銀行、保險、證券產業，延伸至旅館。與本間教授的合作也由日本的能源及污染議題，延伸至日本的旅館業者研究。

隨著世界各國人均所得的提升，休閒觀光的需求日益提升。休閒觀光產業有服務的面向，也有自然與人文資源的面向，是結合服務科學及資源經濟研究典範與方法的重要新興領域。

三、考察參觀活動

敝人跟共同著作人 Honma 教授在會議期間密集討論了未來一年的合作主題。合作模式仍由敝人提供研究方法及研究命題，Honma 教授提供日本數據及背景描述。目前初步擬定新議題包含：日本各區域對外貿易與總要素能源效率之關係、旅客國籍來源對日本旅館業者經營效率之影響、OECD 國家各部門總要素能源效率之比較等。預計將有另一波合作國際期刊發表的產生。

亦獲得本屆會議最佳論文獎的後藤美香（Mika Goto）教授，現任職於日本的電力中央研究所擔任上席研究員。她利用 DEA 方法在優良類國際期刊發表多篇論文，並對 DEA 方法之發展有所貢獻。她對於 Honma 教授與敝人的合作模式及成果非常感興趣。她亦於 7 月 21-24 日來台參加亞洲生產力會議。她希望 Honma 教授就能源效率議題進行合作研究。

此外，敝人亦與紐西蘭 Christchurch Polytechnic Institute of Technology 的 Mehdhi Asgarkhani 教授討論了招收外籍學生對各大學經營績效的影響。該校進行的外籍學生學習輔導計畫顯示，經過專責導師輔導後，非英語系外籍學生的學業表現經常超越本地的學生。

任教於美國 Gonza University 的莊道達教授（本所校友）對於台灣管理學的教育現況非常關心，他認為國際化是必須走的路，應該鼓勵學生赴海外留學或短期進修。

擔任多本優良類國際期刊編輯委員的 Ali M. Kutan 教授（Southern Illinois University），則主動進行邀稿。他強調投稿管理類國際期刊時，應該加強英文寫作及管理意涵之論述，不宜單純套用（賣弄）計量經濟技巧，必須對管理期刊的讀者有所啟發。他也表示願意到台灣與教授及研究生進行交流。

四、建議

藉由這次赴北九州市參加國際會議，敝人除了宣讀論文、獲頒研究獎項外，並與海外共

同著作人面對面討論未來一年的合作研究議題。畢竟，平時電子郵件的密集往返討論畢竟不如面對面的討論。未來可以繼續主動建議海外共同著作人參加同一個國際研討會，藉由與會期間當面討論並介紹彼此的學術界友人。

觀光休閒研究具有學術及實務上的雙重重要性，值得敝人進一步投入發展。影響國家創新系統的效率的變數很多，這篇文章考慮的影響無效率之因素常見於過去科技管理文獻中，未來宜考慮其他變數，例如：公部門治理等，對於國家創新系統效率之影響。

五、攜回資料名稱及內容

會議手冊、會議論文光碟、各國學者名片等。

六、其他

無。

國立交通大學
National Chiao Tung University

類別：國際學術會議

題目：International Conference of Pacific Rim Management: 2010 ACME Annual Meetings

服務機關：國立交通大學經營管理研究所

姓名職稱：胡均立 教授兼所長

前往地區：澳門科技大學 (MUST)

出國期間：自 99 年 7 月 15 日至 99 年 7 月 19 日

報告日期：99 年 7 月 31 日

一、參加經過

近年敝人將經濟分析應用於管理議題的研究上，包含銀行、旅館、電子電機、醫院等產業的經營績效之研究。這次報告的議題是台灣各縣市警察局的效率，採用共同邊界（metafrontier）模型。ACME 的全名是 Association for Chinese Management Educators (華人管理教育者學會)，過去兩屆年會在北美舉行，這屆首度移到東亞舉行，會議地點為澳門科技大學。與會者多為在北美任教的華人管理學者。論文投稿會議時並經過匿名審查，審

查人質疑此篇論文產出項的選取，並認為 Tobit 迴歸中的環境變數不足以代表影響各縣市警察局效率的重要外部因素。由於時間限制，敝人於出發前修改了一些內文寫作，但尚未能就模型設定及實證分析予以重作。敝人是在 7 月 15 日上午跟隨大同大學教授及研究生團體從桃園機場飛到澳門機場，並於 7 月 18 日下午由香港機場飛回桃園機場。會議期間為 7 月 15 日至 7 月 17 日。

二、心得

敝人係於 7 月 16 日下午發表「A Metafrontier Efficiency Analysis of Regional Police Agencies in Taiwan」一文，與交通大學碩士董育穎及博士生羅凱文合著。這篇論文依循 Battese et al. (2004) 所提出的共同邊界法程序，但利用 Färe et al. (1985) 所提出的變動規模報酬、產出導向的資料包絡分法 (DEA) 模型來計算多產出下的各縣市警察局效率值。研究對象為 2003-2007 的台灣本島及澎湖的 23 縣市警察局。模型中有六個產出項，分別為暴力犯罪率、暴力犯罪破案率、非暴力犯罪率、非暴力犯罪破案率、總犯罪率、總犯罪破案率。由於犯罪率是非意欲（undesirable）產出，故在模型計算時予以取倒數處理，即其他因素不變下，犯罪率越高、效率值越低。

我們將 23 縣市依照人口密集度分成兩群，分別計算其分年度分群的效率值；並將所有縣市納入，計算其面對各年度共同邊界下的效率值。對第 t 期的第 i 個縣市警察局而言，Battese et al. (2004) 以推導方式證明下列關係式：

$$\text{共同邊界效率值 } (TE_{it}^*) = \text{ 分群效率值 } (TE_{it}) \times \text{ 技術落差比例 } (TGR_{it})$$

主要實證結果如下：共同邊界法分析發現高人口密集度的縣市警察局相對缺乏管理效率，低人口密度警察局的技術落差較大。Tobit 迴歸結果顯示：教育程度的提升有助於改善該縣市警察局的效率，15-64 歲人口比例的提高顯著惡化該縣市的效率，社會（人口）增加率顯著惡化該縣市警察局的效率。

三、考察參觀活動

在會場與來自北美、台灣、中國大陸及港澳的學者廣泛交換意見，並出席了其他場次的論文發表並主動提問。來自加拿大 University of Windsor 的 Jerry Sun 教授建議敝人可再考慮

警察局進行犯罪防制宣導等產出項，並建議可以考慮放入投入項。國立中正大學鎮明常教授認為拙作實證結果大致合理，但按人口密集度分群則可再斟酌。來自加拿大 University of Windsor 的 Diana Kao 副院長則詢問將研究延伸至兩岸三地比較的可能性。來自加拿大 Ryerson University 的 Guoping Liu 教授則詢問資料來源及台灣官方統計數據可否從網路上取得。敝人自己則發現非意欲產出取倒數將扭曲效率值，未來宜以方向性距離函數（directional distance function）直接處理具有非意欲產出的情形。

來自 Eastern Michigan University 的 Huei Lee 教授則詢問未來台灣接辦 ACME 年會的可能性。文化大學校長吳萬益教授當場建議國立成功大學明年在台南接辦此會議。來自 University of Southern Mississippi 的 Kuo Lane Chen 則鼓勵敝人日後應多出席在北美舉辦的管理學類國際會議，並連結兩岸三地的管理類研究議題。Chen 教授同時指出台灣管理學界的國際化與海外互動仍可進一步加強提升。國立高雄師範大學孫培真所長則建議母校交通大學在管理學教育上應更積極扮演主動角色。

四、建議

藉由這次赴澳門參加國際會議，敝人除了宣讀論文外，並與海外華人管理學者對面討論未來可能的合作研究議題。除了分別作台灣及中國大陸的實證研究外，未來兩岸三地的比較研究是值得進一步發展的方向。此外，將來 ACME 年會回到北美舉辦的時候，也可以把握機會報名宣讀論文。本校除了論文發表者外，未來在國際會議的舉辦上應更積極扮演組團參與者、特定場次籌備者或聯合主辦者的角色。

五、攜回資料名稱及內容

會議議程、會議論文光碟、各國學者名片等。

六、其他

無。

國立交通大學
National Chiao Tung University

類別：國際學術會議

題目： The 1st Congress of East Asian Association of Environmental and Resource Economics (EAAERE)

服務機關：國立交通大學經營管理研究所

姓名職稱：胡均立 教授兼所長

前往國家：日本北海道

出國期間：自 99 年 8 月 16 日至 99 年 8 月 21 日

報告日期：99 年 9 月 15 日

一、參加經過

敝人長期研究資源與環境管理議題。東亞環境與資源經濟的學者過去在東亞地區輪流主辦學術會議。敝人 2006 年 11 月曾經參加過在韓國首爾舉行的 The 2nd East Asian Symposium on Environmental and Natural Resources Economics，2008 年 2 月出席在日本東京的 The 3rd East Asian Symposium on Environmental and Natural Resources，2009 年 3 月聯合主辦並出席了 The 4th East Asian Symposium on Environmental and Natural Resources。經

過多年的密切往來與互動，東亞地區的環境與資源經濟學者決定成立 East Asian Association of Environmental and Resource Economics (EAAERE)，並於今年 8 月在北海道大學舉行成立大會及第一次年會。

敝人是在 8 月 16 日上午跟隨中華經濟研究院蕭代基院長及開南大學觀光運輸學院黃宗煌院長領導的台灣教授及專家團體，從桃園機場飛到新千歲機場，並於 8 月 21 日上午由新千歲機場飛回桃園機場。會議地點在北海道大學，會議期間為 8 月 17 日至 8 月 19 日。8 月 17 日下午是日本、英國、德國、台灣、中國大陸、韓國的貴賓演講，主題為各地區的綠色租稅改革。分組論文發表則安排於 8 月 18、19 兩日。由於蕭代基院長於出發前在桃園機場就特別提醒，敝人於 8 月 17 日至 8 月 19 日三天內全程參與，並踴躍發言。

二、心得

敝人係於 8 月 19 日上午發表「Gresham's Law in Environmental Protection」一文，與高雄大學經營管理研究所楊雅博教授合著。這篇論文係利用賽局理論模型推導命題。其研究動機是見到中國大陸因為環境執法寬鬆，使得無效率的高污染廠商有機會藉由逃避環境保護成本來壓低產品售價，不利於合法負擔環境保護及污染防治成本的廠商，因而造成「劣幣驅逐良幣」的現象。我們利用賽局理論推導證明，在不完全環境執法下，高成本廠商得以藉由策略性地違反環保法規，來阻卻守法的低成本廠商進入市場。僅是訂定更嚴格的環境保護標準，但未同時提高執法機率及提高罰金，有助於違法廠商將守法廠商趕出市場。本理論模型得以用來解釋何以高污染與低生產效率經常並存於開發中經濟體。

本篇文章的評論人為東京大學國際協力學 Tetsuya Tsurumi 教授，但是他表示主要進行實證研究，鮮少做經濟理論的推導。與會者普遍認為這篇研究的方法學嚴謹且結論大致合理，但是理論文章所能處理的變數有限，且並無出人意表之外的新穎發現，僅是從經濟理論來解釋已經可觀察到的現象何以長期存在。

敝人同時擔任政治大學經濟系李慧琳教授的評論人。她發表的題目為「Economic Impact of the Public Expenditure on Building On-Site Wastewater Treatment Facilities in the Taipei Metropolitan Area: An Application of the Regional Input-Output Model」，模擬計算台北都會區地方政府廢水處理工程支出的投入產出分析。其中包含 55 個部門、19 個台北縣市的鄉鎮市。結果發現廢水處理工程支出對帶動地方產業發展及就業具有顯著正向影響，其中又以台北縣

板橋市的支出所帶的乘數效果最大。由於從事中國大陸研究，敝人特別積極參加了中國大陸研究相關的場次。例如：8 月 19 日下午的非再生能源場次中，京都大學 Vivian Leung 博士生報告了「Limited Diffusion of LPG-powered Vehicles in China」，一橋大學博士生 Fu Zhe 報告了「Motorization and Air Pollution in Shanghai」，日本國立環境研究所 Azusa Okagawa 博士報告了「Promotion of Energy Supply from West to East in China」。敝人都舉手發問，並提供論文修改建議。

三、考察參觀活動

日本山口大學陳禮俊教授，是知名的台灣旅日學者。他目前在中國大陸雲南等地區進行田野調查，並且在台灣推廣山坡地生態工法。敝人多次與他交談，也了解了一些日本學界與東亞學術界在永續發展上的合作項目。

韓國釜山國立大學 Sang-Mok Kang,曾多次在不同的國際會議中與敝人相遇,他曾於 2009 年來台參加亞洲生產力會議。他經常使用資料包絡分析法進行產業綠色效率計算，進行南韓及中國大陸廠商綠色效率及生產力的比較研究。他最近從事共同邊界法的應用。敝人建議進行第二階段的迴歸分析，以估計環境變數對綠色效率及生產力的影響，而非僅就效率值的計算結果進行主觀猜測。

敝人與去年曾蒞臨本所演講的日本慶應大學 Eiji Hosoda 教授，就環境管制理論可能的發展進行討論。此次的大會的主題之一是綠色租稅，不同租稅工具的組合或選擇將是有趣的環境管制議題之一。Hosoda 教授是國際期刊 *Environmental Economics and Policy Studies* (EEPS) 的主編，敝人曾經在該期刊發表過學術論文，東亞環境與資源經濟學會計畫將 EEPS 變成該學會的主要出版期刊，並於成立大會上正式宣布。

京都大學 Akihisa Mori 教授，一向對於促進東亞環境與資源學術交流不遺餘力。他提及除了學術研究以外，藉由環境學術交流活動，落實綠色租稅改革理論於東南亞等開發中國家的可能性。

此外，許多同行的台灣學者表示未來合作發表論文的意願。畢竟在台灣進行跨校研究合作，具有可以經常見面討論的優點，也對增加台灣在國際發表上的量能，有所助益。

四、建議

藉由這次赴北海道大學參加國際會議，敝人除了宣讀論文外，並與海外學者面對面討論未來的研究議題趨勢。畢竟，平時電子郵件的密集往返討論不如面對面的討論。未來可以主動建議海外友人參加同一個國際研討會，藉由與會期間當面討論並介紹彼此的學術界朋友。

台灣是東亞環境與資源經濟學會的正式發起團體，我方宜積極參與，鞏固會籍，並藉此與東亞環境與資源經濟學者密切交流，落實國際合作項目。

五、攜回資料名稱及內容

會議手冊、會議論文光碟、各國學者名片等。

六、其他

無。

無研發成果推廣資料

98 年度專題研究計畫研究成果彙整表

<table>
<tr><td colspan="3">計畫主持人：胡均立</td><td colspan="5">計畫編號：98-2410-H-009-055-</td></tr>
<tr><td colspan="8">計畫名稱：國際能源價格與台灣總體經濟變數之關聯性分析</td></tr>
<tr><td colspan="3" rowspan="2">成果項目</td><td colspan="3">量化</td><td rowspan="2">單位</td><td rowspan="2">備註（質化說明：如數個計畫共同成果、成果列為該期刊之封面故事...等）</td></tr>
<tr><td>實際已達成數（被接受或已發表）</td><td>預期總達成數(含實際已達成數)</td><td>本計畫實際貢獻百分比</td></tr>
<tr><td rowspan="5">國內</td><td rowspan="4">論文著作</td><td>期刊論文</td><td>2</td><td>2</td><td>100%</td><td rowspan="4">篇</td><td>Chang, Tzu-Pu and Jin-Li Hu (2010), ’ ’ ’ ’ ’ ’ ’ ’ Total-factor Energy Productivity Growth, Technical Progress, and Efficiency Change: An Empirical Study of China,’ ’ ’ ’ ’ ’ ’ ’ Applied Energy (SCI, IF = 2.209), 87(10), 3262-3270.

Hu, Jin-Li, Jui-Chu Lin, Tzu-Cheng Yang and Meng-Chi Chi (2010), ’ ’ ’ ’ ’ ’ ’ ’ The Subsidy-based Solar Energy Policy in Taiwan,’ ’ ’ ’ ’ ’ ’ ’ Open Environmental Sciences, 4, 8-13.</td></tr>
<tr><td>研究報告/技術報告</td><td>0</td><td>0</td><td>100%</td><td></td></tr>
<tr><td>研討會論文</td><td>5</td><td>3</td><td>100%</td><td>2010 年 4 月 1-5 日，兩岸經濟與經營管理博鰲高峰論壇，中國大陸海南島博鰲。

2010 年 7 月 4-9 日，International Conference on Business and Information, Kitakyushu, Japan. （獲得 Best Paper Award）

2010 年 7 月 15-18 日，ACME Annual Meeting, Macau, China, 2010.

2010 年 7 月 22-23 日，Asia-Pacific Productivity Conference, Taipei, Taiwan.

2010 年 8 月 16-21 日，1st Congress of East Asian Association of Environmental and Natural Resource Economics, Hokkaido, Japan.</td></tr>
<tr><td>專書</td><td>0</td><td>0</td><td>100%</td><td></td></tr>
<tr><td>專利</td><td>申請中件數</td><td>0</td><td>0</td><td>100%</td><td>件</td><td></td></tr>
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		已獲得件數	0	0	100%		
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		研究報告/技術報告	0	0	100%		
		研討會論文	0	0	100%		
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	專利	申請中件數	0	0	100%	件	
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其他成果(無法以量化表達之成果如辦理學術活動、獲得獎項、重要國際合作、研究成果國際影響力及其他協助產業技術發展之具體效益事項等，請以文字敘述填列。)	以資源經濟及休閒管理之結合研究，獲得 International Conference on Business and Information, Kitakyushu, Japan 之 Best Paper Award。 首次建構並提出 Total-factor Energy Productivity Index (TFEPI) 指標，並發表於優良類國際期刊 Applied Energy (SCI, IF = 2.209)。 能源經濟研究成果獲得海峽兩岸學者之高頻率引述及來函索取已發表期刊論文成果。

	成果項目	量化	名稱或內容性質簡述
科教處計畫加填項目	測驗工具(含質性與量性)	0	
	課程/模組	0	
	電腦及網路系統或工具	0	
	教材	0	
	舉辦之活動/競賽	0	
	研討會/工作坊	0	
	電子報、網站	0	

目	計畫成果推廣之參與（閱聽）人數	0	

國科會補助專題研究計畫成果報告自評表

請就研究內容與原計畫相符程度、達成預期目標情況、研究成果之學術或應用價值（簡要敘述成果所代表之意義、價值、影響或進一步發展之可能性）、是否適合在學術期刊發表或申請專利、主要發現或其他有關價值等，作一綜合評估。

1. 請就研究內容與原計畫相符程度、達成預期目標情況作一綜合評估

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論文：■已發表 □未發表之文稿 □撰寫中 □無

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本研究計畫深入分析國際能源價格與台灣總體經濟變數之關聯，正式將能源變數與台灣總體經濟情勢進行連結，應用嚴謹的時間序列模型進行估計及檢定。研究成果除已陸續投稿並發表於優良類國際期刊外，對於台灣未來的能源經濟政策，亦具有諮詢參考價值。