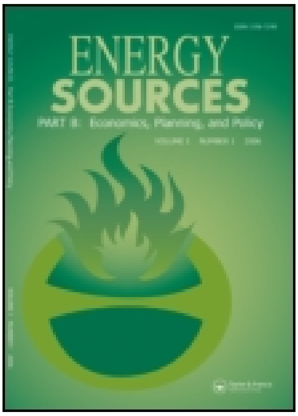

Energy Sources, Part B: Economics, Planning, and Policy
Publication details, including instructions for authors and subscription information:
http://www.tandfonline.com/loi/uesb20

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Published online: 05 Nov 2012.

To cite this article: J.-L. Hu & C.-H. Lin (2013) Examining Asymmetric Behavior Between Energy Consumption and Economic Growth in Taiwan, Energy Sources, Part B: Economics, Planning, and Policy, 8:1, 76-85, DOI: 10.1080/15567240903160922

To link to this article: http://dx.doi.org/10.1080/15567240903160922

Energy Sources, Part B, 8:76–85, 2013

ISSN: 1556-7249 print/1556-7257 online
DOI: 10.1080/15567240903160922

Examining Asymmetric Behavior Between Energy Consumption and Economic Growth in Taiwan

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Energy consumption growth has recently been higher than economic growth in Taiwan, worsening energy efficiency. This article utilizes the momentum-threshold autoregressive cointegration method to examine the long-run equilibrium relationship between economic growth and energy consumption growth, particularly focusing on the most energy-consuming sectors in Taiwan. Allowing for a threshold effect sheds new light on the explanation of the non-linear characteristics of the energy-growth link. The results indicate that economic growth is non-linearly cointegrated with energy consumption when the asymmetric adjusting behavior is confirmed. Specifically, we find that the deviations adjust persistently toward equilibrium in a relative energy efficiency regime for the aggregate-level, while for the sector level in a relative energy inefficiency regime. The government authority should conduct effective energy demand-side management to improve energy efficiency.

Keywords: asymmetric adjustments, energy efficiency, momentum-threshold autoregressive (M-TAR), non-linearity, threshold cointegration

1. INTRODUCTION

Energy is one of the critical determinants for economic growth. To maintain higher economic growth, rapidly growing developing economies are confronted with substantial demand for various energy sources. Since the early 1980s, energy demand on a national and international basis has been extensively analyzed, initially motivated by concerns about security due to energy supply in view of the twin oil price shocks in the 1970s and later due to concerns about climate change. On account of growing pressure exerted on governments to reduce carbon dioxide emissions in order to ease up the rate of climate change, many countries worry about the negative impact on economic growth caused by the restricted use of fossil fuels. Hence, various economic policies and options have been studied to practice energy conservation without harming economic growth.

Because of the critical role played by energy in economic growth, an energy conservation policy (whether or not it can successfully be propagated within an individual country) has been a striking topic widely explored since the late 1970s. The directions of the causal relationship between energy consumption and economic growth can be categorized into four types and evidence in either direction has important implications for an energy policy. First, if there is unidirectional

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Granger causality running from economic growth to energy consumption, then it implies that the policies for reducing energy consumption may be implemented with little or no adverse effects on economic growth, such as in a less energy-dependent economy (Chontanawat et al., 2008). Additionally, a permanent increase in economic growth may lead to a permanent increase in energy consumption. Second, if the unidirectional causality runs from energy consumption to economic growth, then it suggests that restrictions on the use of energy consumption may adversely affect economic growth (Narayan and Singh, 2007). Third, if there is a bidirectional causal relationship, then it implies that energy consumption and economic growth are jointly determined and affected at the same time (Yang, 2000). Finally, the finding of no causality in either direction, the so-called "neutrality hypothesis," means that energy conservation policies do not affect economic growth (Wolde-Rufael, 2005).

Although existing studies have broadly shown whether energy consumption is a factor of economic growth and/or vice versa, it is necessary to detect their cointegrated or long-run equilibrium relationships either for a bivariate or multivariate framework. The cointegration test is preferred over conventional methodology, because the relationship found by using ordinary regression analysis with a time series variable could be spurious. Yu and Jin (1992) provide a pioneering work to detect whether energy and output are cointegrated. They find that no such relationship exists between energy use and either employment or an index of industrial production. Stern (2000) argues that bivariate tests may fail to detect a causal relationship, because of the substitution effects that may occur between energy and other inputs. He also notes that the multivariate methodology is important, because changes in energy consumption are frequently countered by the substitution of other factors of production, resulting in an insignificant overall effect of energy use on economic growth.

The aim of this study is to analyze the long-run relationship between energy consumption and economic growth for Taiwan using a newly developed momentum threshold-autoregressive (M-TAR) cointegration method by Enders and Siklos (2001), who allow for asymmetric behavior. We examine energy consumption not only in the aggregate level, but also classified into industrial, transportation and residential sectors, separately, since they are by far the most important energy-consuming sectors. Why do we need to take account of any asymmetric adjustment behavior between energy consumption and economic growth in Taiwan? The plausible answer is that the energy market has compact allied relationships with an economic system and is extremely correlated with the economic system. We should be conscious of the external factors that could impact the world's energy consumption and economic activities. Moreover, the M-TAR adjustment can be particularly useful when policy makers are viewed as attempting to smooth out any large changes in a series.

2. METHOD

2.1. Data Source

The data used in this study consist of total energy consumption (EC), the most energy consuming sectors of industrial (IND), transportation (TRA) and residential (RES), and gross domestic production (GDP). The original unit of energy consumption is measured as kiloliters of oil equivalent (KLOE). The nominal GDP series in the national currency is transformed into real GDP in 2001 prices, using GDP deflators. The data are in quarterly frequency covering from 1982:1 to 2006:4, compiled from the Advanced Retrieval Econometric Modeling System (AREMOS) economic-statistic database. To obtain the reliable outcomes, all variables are seasonally adjusted and converted into natural logarithms so that they can be interpreted in growth terms after taking first difference.

2.2. The Threshold Cointegration Model

Threshold cointegration was developed by Enders and Siklos (2001), who make an extension to the residual-based two-stage symmetric Engle and Granger (1987) estimation strategy to test the long-run relationship between two time series of variables, taking the property of asymmetry into account. Enders and Siklos (2001) generalize the Enders and Granger (1998) M-TAR test for unit roots to a multivariate context. The resulting M-TAR testing procedure has shown good power and size properties relative to the alternative assumption of symmetric adjustment.

When examining the behavior of GDP and energy consumption, we need to consider the long-run relationship as follows:

$$Y_t = \alpha_0 + \alpha_1 EC_t + u_t, \tag{1}$$

where Y_t represents real GDP, EC_t is energy consumption with respect to aggregate, IND, TRA, and RES, and u_t is the disturbance term that may be serially correlated. The existence of a long-run equilibrium relationship involves the stationarity of u_t. In order to investigate the stationarity of u_t, whether or not $-2 < \rho < 0$ has to be tested in the second step procedure is given by:

$$\Delta u_t = \rho u_{t-1} + \varepsilon_t, \tag{2}$$

where ε_t is the white-noise disturbance and the residuals from Eq. (1) are used to estimate Eq. (2). Rejecting the null hypothesis of no cointegration (i.e., accepting the alternative hypothesis $-2 < \rho < 0$) implies that the residuals in Eq. (1) are stationary with 0 mean.

The standard cointegration framework assuming symmetric adjustment toward equilibrium in Eq. (2) is misspecified if the adjustment process is asymmetric. Therefore, the residuals, $\hat{u}_t$, from Eq. (1) are then used to estimate the following M-TAR model:

$$\Delta\hat{u}_t = M_t\rho_1\hat{u}_{t-1} + (1 - M_t)\rho_2\hat{u}_{t-1} + \sum_{i=1}^{p}\gamma_i\Delta\hat{u}_{t-1} + \varepsilon_t, \tag{3}$$

where ρ_1 and ρ_2 are the respective speed of adjustment coefficients of $\Delta\hat{u}_t$, and $\varepsilon_t \sim I.I.D$ (0, σ^2). The lagged values of $\Delta\hat{u}_t$ are meant to yield uncorrelated residuals and can be determined by Akaike Information Criterion (AIC) model-selection criteria. Instead of estimating Eq. (3) with the Heaviside indicator depending on the level of u_{t-1}, the decay could be allowed depending on the previous period changes in u_{t-1}. The Heaviside indicator function is denoted as follows:

$$M_t = \begin{cases} 1 & \text{if } \Delta\hat{u}_{t-1} \geq \tau \\ 0 & \text{if } \Delta\hat{u}_{t-1} < \tau \end{cases} \tag{4}$$

and τ is the value of threshold.

The threshold parameter, τ, is endogenously determined utilizing Chan's (1993) method to search for the consistent estimate of the threshold. This method sorts the estimated residual series in ascending order and is called $u_1^\tau < u_2^\tau < \cdots < u_T^\tau$, where T denotes the number of usable observations. The largest and smallest 15% of the $\{u_i^\tau\}$ values are excluded. The estimated threshold yielding the lowest residual sum of squares is deemed to be the appropriate estimate of the threshold over the remaining 70%.

The M-TAR model allows the $\{u_t\}$ series to exhibit more "momentum" in one direction than the other and allows the variable of interest to display various amounts of autoregressive decay depending on whether the series is increasing or decreasing—that is, the adjustment is modeled by $\rho_1\hat{u}_{t-1}$ if $\Delta\hat{u}_{t-1}$ exceeds the endogenous threshold and by $\rho_2\hat{u}_{t-1}$ if $\Delta\hat{u}_{t-1}$ is below the respective threshold. To test whether there is threshold cointegration, we test for symmetric adjustment when the null hypothesis of non-cointegration ($H_o : \rho_1 = \rho_2 = 0$) is rejected. Moreover, we also proceed with another test for symmetric adjustment such as *Ho*: $H_o : \rho_1 = \rho_2$, using a standard F-test. If the null hypothesis of symmetric adjustment is rejected and $|\rho_1| < |\rho_2|$, then the M-TAR model exhibits little decay for positive Δu_{t-1}, but substantial decay for negative Δu_{t-1}. In a sense, in the M-TAR model increases tend to persist, but decreases tend to revert quickly back toward the threshold. The critical values for the null hypothesis $\rho_1 = \rho_2 = 0$ depend on the number of variables in the cointegrating vector and the number of lags in the form of Eq. (3). Enders and Siklos (2001) report critical values for the M-TAR models, called Φ_μ^* test statistic.

3. EMPIRICAL RESULTS

3.1. Unit Root Tests

Before performing cointegration analysis, we use the Augmented Dickey-Fuller (ADF; Dickey and Fuller, 1979) and the Kwiatkowski-Phillips-Schmidt-Shinn (KPSS; Kwiatkowski et al., 1992) methods to identify the order of integration for each variable. We include a constant term and a time trend in these tests. In Table 1, the ADF tests show that the unit root hypothesis cannot be rejected at any significant level for each variable in levels. Further investigations of the unit root hypothesis in the first differenced variables are stationary. We also apply another KPSS unit root test based on the null hypothesis of no unit root. The results indicate that the null hypothesis of stationarity is rejected at least at the 5% significance level. Thus, all series are found to be integrated of order $I(1)$.

3.2. Cointegration Analysis With Asymmetric Adjustment

Table 2 contains cointegration test results of long-run equilibrium relationships at different energy consumption sectors in the form of Eq. (1), when considering threshold and momentum adjustment. The table reports values of the adjustment coefficients ρ_1 and ρ_2, the Φ_μ^*-statistics for the null hypothesis of a unit root in u_t (i.e., no cointegration) against the alternative of cointegration with asymmetric adjustment. The lag length is selected such that the AIC is minimized. We also use the

TABLE 1
Unit Root Tests

	Levels		*First Differences*	
	ADF	*KPSS*	*ADF*	*KPSS*
GDP	−0.915[8]	0.302[8]***	−3.280[7]*	0.095[6]
EC	−0.494[4]	0.162[6]**	−7.955[3]***	0.054[5]
IND	−2.361[1]	0.317[8]***	−14.859[0]***	0.083[12]
TRA	0.128[1]	0.306[8]***	−12.869[0]***	0.117[3]
RES	−0.399[4]	0.259[9]***	−8.985[3]***	0.041[6]

Note: (***), (**), and (*) in ADF tests respectively indicate the rejection of the null hypothesis of series has a unit root at 1%, 5%, and 10% levels of significance, while in KPSS tests indicate the rejection of the null hypothesis of series is stationary. The numbers inside the brackets are the optimum lag lengths determined using AIC in ADF tests and the bandwidth is used using the Newey-West method in KPSS tests.

TABLE 2
Results of the Asymmetric Cointegration Tests for GDP and Energy Consumption

	Total	*Industrial*	*Transportation*	*Residential*
τ	−0.00298	−0.02281	−0.00122	−0.00622
ρ_1	−0.201 (−2.87)	−0.036 (−0.94)	−0.084 (−1.71)	−0.516 (−2.79)
ρ_2	−0.071 (−0.89)	−0.491 (−3.38)	−0.162 (−2.51)	−0.297 (−1.41)
Φ_{μ}^{*}	7.002**	9.397***	9.842***	1.117
$\rho_1 = \rho_2$ (F-test)	4.703**	6.060***	4.926***	2.074
Lags	2	1	3	4
$Q(4)$	1.698 (0.79)	6.166 (0.19)	4.286 (0.37)	0.631 (0.96)
AIC	−6.497	−5.927	−6.570	−4.999

Note: Values for ρ_1 and ρ_2 in the parentheses are t-statistic. τ is the estimated threshold. For the Φ_{μ}^{*} test statistic (i.e., test for the null hypothesis of $\rho_1 = \rho_2 = 0$), critical values are approximately 5.20 for the 10% significance level, 6.28 for the 5% significance level, and 8.82 for the 1% significance level. Lags denote the lag order of the differenced residuals calculated from the base model. $Q(4)$ is the Ljung-Box statistic that the first 4 of the residual autocorrelations are jointly equal to 0 with the significance level in parentheses. Levels of significance are indicated by (***) and (**) for 1% and 5%, respectively.

F-test to test whether the adjustment back to equilibrium is symmetric $\rho_1 = \rho_2$. The consistent estimates of the threshold as well as the values of AIC are also presented in the table.

The estimated Φ_{μ}^{*}-statistics for the relationship between total energy consumption and GDP is 7.002, where the critical values reported in Enders and Siklos (2001) at the 5% significance level with 100 observations and one lagged change is 6.28. As such, we reject the null hypothesis of a unit root in favor of cointegration with asymmetric adjustment. Except for the residential sector, we also find similar evidence in the cases of the relationships between GDP and sector-based energy consumption such as IND and TRA. The unreported results show that TAR adjustment has a less statistical significance for all energy-GDP specifications, favoring the M-TAR adjustment. With regard to the null hypothesis of symmetric adjustment $\rho_1 = \rho_2$, the F-statistics strongly reject symmetric adjustment for M-TAR specifications at conventional significance levels in all models (except for the residential sector). This evidence favors that the adjustment back to equilibrium between GDP and energy is non-linear.

The point estimates for ρ_1 and ρ_2 suggest substantially faster convergence for positive deviations (i.e., above the threshold and defined as the energy-efficiency regime) from long-run equilibrium than negative deviations (i.e., below the threshold and defined as the energy-inefficiency regime) with the exception of the total-GDP nexus model. For example, in the total energy consumption-GDP model, the point estimates of ρ_1 and ρ_2 suggest that negative deviations from the long-run equilibrium resulting from decreases in industrial sector energy or increases in GDP ($\Delta u_{t-1} < -0.00298$) are eliminated lower than positive deviations are eliminated. When comparing these models, the largest discrepancy between the elimination of below and above threshold deviations occurs at the transmission from industrial sector energy to GDP where negative deviations are eliminated at 49.1% per quarter, while positive deviations are eliminated only at a rate of 3.6%.

3.3. Estimates of the Momentum Threshold Error-correction Model

Given the threshold cointegration results confirmed previously, Enders and Siklos (2001) further elaborate explicitly for cointegration with asymmetric error-correction. As noted by Enders and Siklos (2001), a more general specification may incorporate threshold effects of lagged ΔY_t and

ΔEC_t depending on whether the error-correction term is positive or negative. Utilizing the long-run equilibrium between energy consumption and GDP, we estimate the following momentum threshold error-correction model:

$$\Delta Y_t = \mu + \lambda_1 Z_{t-1}^{+} + \lambda_2 Z_{t-1}^{-} + \sum_{k=1}^{p} \alpha_k \Delta Y_{t-k} + \sum_{k=1}^{p} \beta_k \Delta EC_{t-k} + e_{1t} \tag{5}$$

$$\Delta EC_t = \mu + \lambda_1 Z_{t-1}^{+} + \lambda_2 Z_{t-1}^{-} + \sum_{k=1}^{p} \alpha_k \Delta Y_{t-k} + \sum_{k=1}^{p} \beta_k \Delta EC_{t-k} + e_{2t}, \tag{6}$$

where $Z_{t-1}^{+} = I_t \hat{u}_{t-1}$ and $Z_{t-1}^{-} = (1 - I_t)\hat{u}_{t-1}$, and $\hat{u}_{t-1}$ is obtained from the estimated long-run equilibrium. The Heaviside indicator is set in accordance with Eq. (4), given $I_t = 1$ if Δu_{t-1} exceeds a certain threshold value and 0 otherwise. Here, e_{1t} and e_{2t} are white-noise disturbances. The previously-mentioned specification also distinguishes between long- and short-run adjustments. The long-run adjustment is determined by the parameters λ_1 and λ_2. The short-run adjustment is governed by the parameters α_k and β_k (for $k = 1, 2, \ldots, p$) and may come either from its own history of lagged dynamics or from the lagged effects of changes in energy consumption.

Table 3 presents estimates of the error-correction parameters along with test statistics regarding weak exogeneity and Granger causality. It is clear that the point estimates of λ_1 and λ_2, the error-correction coefficients, are noticeably different in all models. While the adjustment speed on the exceeding or underlying threshold level in all GDP models is the “right” direction by acting to eliminate deviations from the long-run equilibrium, the energy consumption model adjusts to the “wrong” direction (have a positive sign) for either or both regimes. In all GDP equations, the adjustment speed responds faster in an energy-inefficiency regime from the long-run equilibrium than in an energy-efficiency regime. For example, in the GDP-industrial model, economic growth adjusts by about 1.5% of an above threshold deviation from the long-run equilibrium (such that $\Delta u_{t-1} \geq -0.02281$), but by 14.7% of a below threshold deviation. However, in the GDP-transportation model, these differences are less pronounced. The t-statistics for the error-correction terms indicate that with the exception of the GDP equation in the GDP-transportation representation model, all adjusting coefficients of the GDP-energy nexus are weakly endogenous with respect to the long-run equilibrium for at least one regime.

Tests of symmetry based on the error-correction model in Eqs. (5) and (6) and short-run and long-run causal relationships are also conducted. The F-statistics for Granger causality indicate that there is a unidirectional Granger causality from total energy consumption to economic growth at conventional significance levels and the GDP-industrial nexus as well. The relationship between GDP and transportation sector energy consumption fails to cause movements in economic growth, and vice versa. From the present time and into the future, energy acts as an engine of economic growth for Taiwan. Furthermore, the jointly coefficients of lagged terms of transportation sector energy are insignificant in the economic growth equation (i.e., F-statistics $= 1.462$, p-value $=$ 0.237), and the jointly coefficients of lagged terms of economic growth are insignificant in the transportation sector energy equation (i.e., F-statistics $= 0.520$, p-value $= 0.596$). This evidence suggests that the neutrality hypothesis can be supported.

3.4. Detecting the Constancy of the Cointegration Space

One problem with time series regression models is that the estimated parameters may change over time. The estimated periods in our study cover a somewhat volatile time of accidental economic events in Taiwan. Unstable parameters can result in model specification and, if left undetected,

TABLE 3
Estimates of the Momentum Threshold Error-correction Model

	Total		*Industrial*		*Transportation*	
	ΔGDP	ΔEC	ΔGDP	ΔIND	ΔGDP	ΔTRA
Intercept	0.006 (3.796)	0.007 (4.352)	0.007 (5.117)	0.007 (3.786)	0.008 (5.841)	0.007 (3.952)
λ_1	−0.009 (−1.857)	0.161 (2.509)	−0.015 (−0.629)	0.005 (0.157)	−0.006 (−0.150)	−0.087 (−1.475)
λ_2	−0.059 (−0.891)	−0.101 (−1.378)	−0.147 (−1.974)	0.249 (1.933)	−0.025 (−0.424)	0.152 (1.889)
$\lambda_1 = \lambda_2$	0.313 [0.577]	7.009 [0.009]	1.978 [0.163]	3.469 [0.066]	0.066 [0.798]	5.862 [0.017]
$\lambda_1 = \lambda_2 = 0$	0.414 [0.662]	3.988 [0.022]	1.406 [0.251]	1.870 [0.160]	0.100 [0.905]	2.936 [0.058]
$\alpha_1 = \alpha_2 = 0$	2.365 [0.099]	0.728 [0.486]	1.146 [0.322]	1.072 [0.347]	1.998 [0.142]	0.520 [0.596]
$\beta_1 = \beta_2 = 0$	7.414 [0.001]	5.762 [0.004]	5.550 [0.005]	4.032 [0.021]	1.462 [0.237]	1.053 [0.353]
$\lambda_1 = \alpha_1 = \alpha_2 = 0$	1.601 [0.195]	2.700 [0.050]	0.932 [0.429]	0.736 [0.533]	1.421 [0.242]	0.900 [0.445]
$\lambda_2 = \alpha_1 = \alpha_2 = 0$	2.541 [0.061]	0.870 [0.460]	1.839 [0.146]	2.489 [0.065]	1.453 [0.233]	1.641 [0.186]
$\lambda_1 = \beta_1 = \beta_2 = 0$	5.055 [0.003]	7.704 [0.000]	3.863 [0.012]	2.763 [0.047]	1.091 [0.357]	1.237 [0.301]
$\lambda_2 = \beta_1 = \beta_2 = 0$	5.721 [0.001]	4.086 [0.009]	4.102 [0.009]	6.852 [0.000]	1.171 [0.325]	2.555 [0.060]
Number of lags	2	2	2	2	2	2
Q(4)	0.109 (0.99)	2.190 (0.70)	0.158 (0.98)	0.607 (0.96)	0.142 (0.98)	3.908 (0.42)

Note: Values for λ_1 and λ_2 in the parentheses are t-statistic. Values in the brackets are the p-values of F-statistic to test the equality of the coefficients of two-regime error-correction terms and dynamic lagged terms. Q(4) is the Ljung-Box statistic that the first 4 of the residual autocorrelations are jointly equal to 0 with the significance level in parentheses.

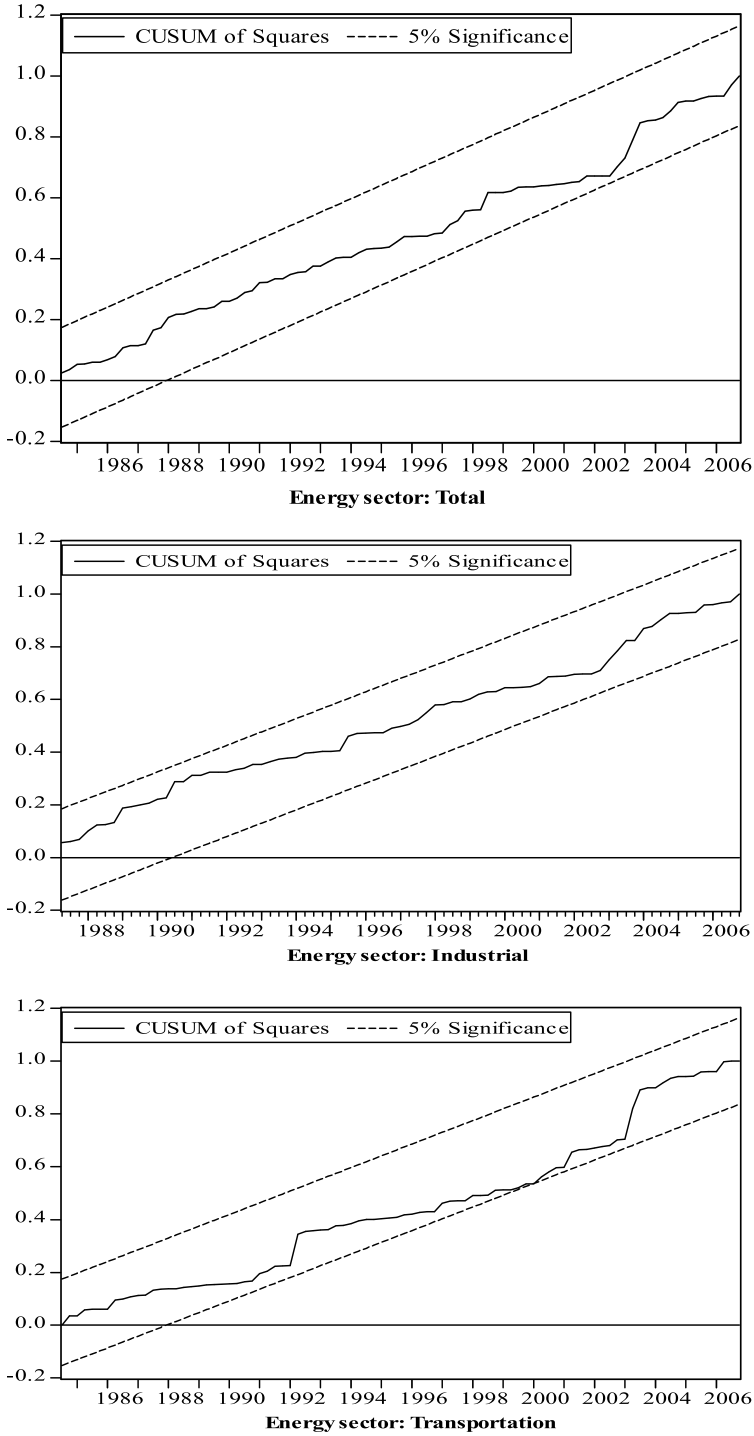

FIGURE 1 Plots of the CUSUM square test.

have the potential to bias the results. To account for this, it is important to check whether the estimated elasticities are stable over time, using the Pesaran and Pesaran (1997) test for parameter instability.

The Pesaran and Pesaran (1997) test amounts to estimating the error-correction models by taking each differenced variable as a dependent variable together with the lagged error-correction term. Testing for the stability of the long-run coefficients in estimating Eq. (5) is carried out using the cumulative sum of recursive residuals (CUSUM) and the CUSUM of square (CUMSUMSQ) tests proposed by Brown et al. (1975). This test uses CUSUMSQ which is updated recursively and are depicted against the break points in the broken sample points to test the null hypothesis that all the coefficients in the growth non-linear model are stable. The graphical representations of the tests are plotted in Figure 1. The CUSUMSQ plots are confined within the 5% critical bounds, suggesting that the residual variance is somewhat stable over time.

4. CONCLUSIONS AND POLICY IMPLICATIONS

This article has considered the possibility of an asymmetric adjusting effect for aggregate/disaggregate sector energy consumption and GDP due to the insight that energy consumption is higher than economic growth in Taiwan, worsening energy efficiency. To address such an issue, we apply the threshold cointegration analysis which allows for an asymmetric adjusting behavior of energy consumption and GDP, and the major findings are as follows. First, we find that there is a non-linear cointegration relationship between energy consumption and GDP with the exception of the residential sector. Second, the estimated vector error-correction models provide strong evidence that the deviations persistently adjust toward long-run equilibrium in a relative energy-efficient regime for the aggregate-level and in a relative energy-inefficient regime for the sector-level. Third, the short-run weak exogeneity Granger causality tests support that there is Granger causality from energy to economic growth in the cases of the aggregate and industrial sector. This evidence suggests that energy plays a crucial component in driving economic growth. Policymakers should take into account the asymmetric adjustment behavior of the energy-growth nexus when building estimation and prediction models of economic growth for Taiwan in the future.

To promote greater energy efficiency and pursue sustainable energy development, Taiwan's government has to implement some courses in the energy market. The developed and developing countries are facing the challenge of sustainable energy development. Energy demand-side management is possibly a useful tool to entail action that influences the quantity or patterns of energy demand consumed by end-users such as actions targeting a reduction of peak demand when energy supply systems are constrained. On the other hand, to pursue economic sustainable development, government authorities should amend renewable energy development targets to enhance the determinants towards the utilization of clean energy such as wind power, solar energy, and biomass fuels. Improving energy efficiency may alleviate the problems of greenhouse gas emissions as well as the "decoupling relationship" between energy consumption and economic growth.

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