Author: S. Buranakunaporn and E. Oczkowski
Author Address: eoczkowski@csu.edu.au
Title: Structural change and Thailand energy demand
Year: 2007
Journal: International Journal of Energy Research
Volume: 31
Pages: 300-314
ISSN: 0363-907X
URL: http://www3.interscience.wiley.com/cgi-bin/abstract/112728530/ABSTRACT
Keywords: Structural Change, Thailand Energy Demand
Abstract: This paper examines the impact of changes in the structure of the economy, radical changes in economic policy and oil price shocks on the relation between Thailand energy demand and its macroeconomic determinants. The impact of these structural changes on the relationship between energy consumption, income, energy prices and structural variation is examined through unit root and cointegration tests, the cointegration relationship and the error correction model. Methods which endogenize the location of an a priori unknown break point are employed to assess the impact of structural change. In general, the recognition of structural change has lead to some unique insights. In particular, the results of some of the conventional unit root and cointegration tests are reversed once structural changes are recognized. Estimates from the cointegrating regression imply long-run income, price, and structural variation elasticities of 0.568, -0.600 and 1.046 respectively. In comparison, estimates from the error correction model suggest a higher short-run income elasticity (0.788) but lower short-run price and structural variation elasticities (-0.522 and 0.491 respectively). One of the important implications of the estimates pertains to the low price elasticity for aggregate energy demand which implies that the over-pricing of energy as a policy instrument is not likely to be very influential for restraining future energy demand. Additionally, taxes on energy prices are unlikely to achieve government goals for energy conservation and environmental improvement, although they may well be efficient for raising revenue.

Structural Change and Thailand Energy Demand

by

Suthep Buranakunaporn
Faculty of Economics,
Ramkhamhaeng University, Thailand.

and

Edward Oczkowski*
School of Commerce, Charles Sturt University
PO Box 588, Wagga Wagga, NSW, 2678, Australia.
Phone: +61 2 69332377 Fax: +61 2 69332930 E-mail: eoczkowski@csu.edu.au

Keywords: Structural change; Energy demand; Thailand.

* Corresponding author. We gratefully acknowledge the comments from two anonymous referees.

Structural Change and Thailand Energy Demand

Abstract

This paper examines the impact of changes in the structure of the economy, radical changes in economic policy and oil price shocks on the relation between Thailand energy demand and its macroeconomic determinants. The impact of these structural changes on the relationship between energy consumption, income, energy prices and structural variation is examined through unit root and cointegration tests, the cointegration relationship and the error correction model. Methods which endogenize the location of an *a priori* unknown break point are employed to assess the impact of structural change. In general, the recognition of structural change has lead to some unique insights. In particular, the results of some of the conventional unit root and cointegration tests are reversed once structural changes are recognized. Estimates from the cointegrating regression imply long-run income, price, and structural variation elasticities of 0.568, -0.600 and 1.046 respectively. In comparison, estimates from the error correction model suggest a higher short-run income elasticity (0.788) but lower short-run price and structural variation elasticities (-0.522 and 0.491 respectively). One of the important implications of the estimates pertains to the low price elasticity for aggregate energy demand which implies that the over-pricing of energy as a policy instrument is not likely to be very influential for restraining future energy demand. Additionally, taxes on energy prices are unlikely to achieve government goals for energy conservation and environmental improvement, although they may well be efficient for raising revenue.

1. Introduction

The time series relationship between energy demand and macroeconomic variables has been extensively studied, representative examples include, Masih and Masih (1996), Asafu-Adjaye (2000) and Oh and Lee (2004). Typically, unit root and cointegration tests are preformed and the cointegration relationship and error correction models (ECM) developed to examine the relationship between national energy consumption and a series of regressors such as Gross Domestic Product (GDP) and energy prices. One aspect which has received relatively scant attention in this literature, however, is the role of structural change and its impact on the stability of these energy demand relations. The purpose of this paper is to examine structural changes in the time series relationship for Thailand energy demand.

Beyond the contribution from examining structural change, this paper adds to the literature on Thai energy demand. Hoa (1993) and Asafu-Adjaye (2000) are the only studies which have provided a macro-econometric analysis of energy demand for Thailand. Their studies concentrated on the examination of the causation between energy demand and economic activities by using the bivariate causality test. Hoa (1993) made use of cointegration analysis to investigate the long-term relationships between crude oil consumption, output growth, and inflation in Thailand. However, his study is simple in its bivariate form and in its methodology. Asafu-Adjaye (2000) used the ECM framework using energy prices (proxied by the CPI) and income as determinants of energy demand.

Our paper differs from previous Thai studies in terms of model specification, time period and data set. The determinants of energy demand comprise of three variables: energy prices, income and an indicator of structural variation (represented by the share of manufacturing output in income). It utilizes the ECM to specify the short-run adjustment model of energy demand. Additionally, the unit root testing, cointegration analysis and the ECM will cover previously unused energy price data.

At the outset it is important to outline the context to energy consumption in Thailand. The industrialization of the Thai economy and the urbanization of Thai society have had a profound impact not only on the growth of energy consumption but also the types of energy consumed. ‘Modern’ or ‘commercial’ energy consumption has been growing at an average rate of 7.6 % since 1980 compared to the 5.9 % rate of growth of total energy consumption, see the data appendix for sources of this and subsequent information. Modern energy consumption rapidly grew at 10.8 and 11.3 % per year during the periods of 1985–1989 and 1990–1994 respectively, but due to the economic crisis its growth was reduced to 4.0 % per year for the period of 1995–1999. While the demand for modern energy has increased, the consumption of ‘traditional’ energy has been stagnant with an average rate growth of 2.0 % during the period of 1980–1999.

The rapid rate of economic growth experienced in Thailand (real GDP growth of 6.2% during 1980-1999) has been the driving force for higher growth in energy demand. Moreover, as the Thai economy is moving towards a more diversified industrial and service-oriented society, the demand for energy is expected to increase further. During the period 1980-1985, the average level of per capita GDP at 1988 prices was 20,983 baht (USD $830) and it increased by 2.32 times to an average level of 34,105 baht (USD $1349) during the period 1995–1999. For the comparable periods, average per capita energy consumption during the period 1980–1985 was 0.329 tons of crude oil equivalent (TOE) and increased by 2.39 times to the level of

0.787 TOE during the period 1995–1999. It is evident that as the level of economic development is higher, people will use more energy in making their life more comfortable and Thailand appears to be no exception.

Total energy intensity in Thailand, as measured by the proportion of total energy consumption per real GDP at 1988 prices, was at an average level of 15.72 TOE per million baht (USD 0.622 million) during the period 1980-1999. Interestingly, modern energy intensity significantly increased from 9.67 TOE per million baht (USD 0.382 million) during the period 1980-1985, to 13.19 TOE per million baht (USD 0.522 million) during the period 1995-1999.

The remainder of the paper is organized as follows. The next section presents the methodology, econometric methods and data to be employed for analysis. Section three outlines and presents the empirical results. Section four concludes with a discussion of the results and their implications.

2. Methodology

It is well known that the standard unit root tests, such as the augmented Dickey-Fuller (ADF) (Dickey and Fuller 1979 and 1981) and Phillips-Perron (PP) (Phillips and Perron 1988) tests, for the random walk hypothesis have low power against the alternative hypothesis of mean reversion in small samples. The problem is especially serious when there exist structural changes in the underlying series (Perron, 1989). In general, failure to account for breaks can produce misleading tests and result in incorrect inference. To account for structural change possibilities in energy demand variables and relationships, we will employ three strategies. First, the Zivot and

Andrews (1992) unit root tests are employed to test for the presence of structural breaks with unknown timing in the individual series of energy demand and macroeconomic variables. Second, we will employ Gregory and Hansen (1996) tests for cointegration where the structural break is test-determined and the cointegrating vectors are allowed to change at an unknown time period. Finally, we will examine parameter instability through cumulative sum of residuals (CUSUM) tests and recursive parameter estimates to assess structural change occurrences in the cointegration relationship and the ECM for Thailand energy demand.

We now turn to the specification of the econometric model. Most previous energy demand studies (see Atkinson and Manning (1995) for a survey) typically assume energy demand depends upon lagged demand, income and prices. In addition to these determinants this paper employs a measure of structural variation to capture the increasing energy demand requirements of the Thai manufacturing sector. The share of manufacturing in GDP has increased from about 15% in 1960 to 35% in 1999.

The definitions and summary statistics of the key variables are: energy demand (E_t : mean = 19194.7, std. dev = 17841.3) is measured by total final energy consumption deflated by the energy price index (KTOE); income (Y_t: mean = 21637.6, std. dev = 24649.5) is measured as GDP per capita (Millions of Baht); energy prices (P_t: mean = 91.545, std. dev = 10.80) are defined by the price index of the component 'fuel and electricity' from the consumer price index, deflated by the consumer price index; the indicator of structural variation (M_t : mean = 22.34, std. dev = 5.75) is measured as the share of manufacturing output in GDP at constant prices (%). All nominal variables are deflated using 1988 prices. The sources for the data are provided in the data appendix.

Various historical events motivate the examination of structural change in Thai energy demand. Over the time period under consideration significant economic events include: the two oil shocks of 1973/4 and 1979/80; radical changes in economic policy such as the first five-year National Economic and Social Development plan in 1961; the 14.8% devaluation of the baht against the US dollar in 1984; financial liberalization in 1987; and the Asian economic crisis of 1997. Even though this motivation is clearly apparent there are significant dangers in using these events to *a priori* identify specific structural breaks in econometric tests and relationships. The assumption that the location of break is known *a priori* has been criticized by a number of studies, most notably by Christiano (1992). Christiano argued that the choice of the breakpoint is in most cases correlated with the data, and this leads to inaccurate inferences and accusations of data mining. As a consequence the methods used in this paper assume the break point is unknown and hence endogenize the location of the break point.

For unit root testing the conventional ADF and PP tests are contrasted with the Zivot and Andrews (1992) sequential break point selection tests. Here the null hypothesis is that the series is integrated without an exogenous structural break against the alternative that the series can be represented by a trend-stationary process with a once only breakpoint occurring at some unknown time. Three different characterizations of the trend-break alternative are considered:

- (A) The Crash model that allows a break in the intercept;
- (B) The Changing Growth model that allows for a break in the slope with the two segments joined at the breakpoint;

- (C) The Mixed model which allows for a simultaneous break in the intercept and the slope.

The aim of the procedure is to sequentially test breakpoint candidates and select that which gives the most weight to the trend-stationary alternative.

For the three models, Zivot and Andrews estimate the testing equation by allowing the break to take place beginning successively in the second, third, fourth, and so on, observation, up to observation $T - 1$, where T stands for the total sample size used in the estimation. The alternative specifications are estimated by OLS, and the length of the lag (k) for the difference terms is determined by starting at $k = 8$, and working backwards until significant values are identified. Ng and Perron (1995) established that this procedure has better small sample properties than information based lag order selection. This choice of the optimal number of lagged terms is carried out for each possible breakpoint considered in the procedure, (see Zivot and Andrews (1992, p.255) for details). The estimate of the breakpoint is that particular observation corresponding to the minimum t-value for the one period lagged term, for each model A, B, and C. In order to test the unit root hypothesis, this minimum t-value is compared with a set of asymptotic critical values (Zivot and Andrews, 1992, pp.256-257).

For cointegration testing, the standard Engle and Granger (1987) residual-based and Johansen and Juselius (1990) VAR approaches are contrasted with Gregory and Hansen (1996) sequential breaking test procedure. The Gregory- Hansen test is developed within the framework of the Engle-Granger residual-based cointegration analysis and can be viewed as a multivariate extension of the endogenous break univariate tests of Zivot and Andrews (1992). The null hypothesis of no-cointegration is tested against the alternative of cointegration with a break in the cointegrating relationship. Three models to take account of different types of structural changes in the cointegrating relationships are considered:

- 1) the *C* level shift model;
- 2) the *C/T* level shift with trend model;
- 3) the C/S regime shift model which allows the cointegrating vector parameters to shift.

These three models are estimated for sequential break points varying from 15 % to 85 % of the sample, with the residuals reserved from each iteration. These residuals are then utilized to evaluate ADF statistics for all possible break points. The Gregory-Hansen test statistic is the smallest value of ADF across all possible break points. The approximate critical values for the cointegration tests in the presence of structural break, obtained through simulation methods, are reported in Gregory and Hansen (1996, Table 1, p.109).

Finally, to evaluate structural changes in the energy demand cointegration relationship and the ECM, we will consider the stability of the estimated parameters by employing CUSUM tests (Brown, Durbin and Evans, 1975) and plotting the recursive parameter estimates (Harvey 1981). The cointegration relationship under consideration is:

$$E_t = \beta_0 + \beta_1 Y_t + \beta_2 P_t + \beta_3 M_t + u_t \qquad (1)$$

where, E is energy demand, Y is income, P is the energy price and M is structural variation. Equation (1) is estimated by ordinary least squares.

The general form of the ECM for Thai energy demand is:

$$\Delta E_t = \alpha_0 + \alpha_1 \Delta \hat{u}_{t-1} + \sum_{i=1}^{m} \alpha_{2i} \Delta E_{t-i} + \sum_{i=1}^{n} \alpha_{3i} \Delta Y_{t-i} + \sum_{i=1}^{o} \alpha_{4i} \Delta P_{t-i} + \sum_{i=1}^{p} \alpha_{5i} \Delta M_{t-i} + \varepsilon_t \qquad (2)$$

where $\hat{u}_{t-1}$ is the error-correction term, $\hat{u}_t$ is the estimated the residual term from eqn (1) and ε_t is an error term. Equation (2) is estimated by ordinary least squares.

3. Results

In this section we present in turn, the results from unit root testing, cointegration testing and estimation of the ECM for Thai energy demand. In all cases conventional tests and models are compared to results which account for structural change.

After determining the lag truncation for each test by minimising the AIC, we follow the sequential testing procedure advocated by (Harris 1995, pp.31-32) to test for unit roots in each series. The sequential testing procedure can indicate whether each unit root test should include a time trend or a constant (drift), or both. If we fail to reject the null hypothesis, then testing continues to more restrictive specifications. Testing ceases when we are able to reject the null hypothesis that a unit root is present.

Tables 1 and 2 present test results for the levels of the data for three types of processes: with trend and drift, without trend and with drift, and without trend and without drift. The ADF and PP test statistics indicate that E_t, Y_t, P_t and M_t are non-stationary at levels in the process with trend and drift. In contrast, the test statistics for the first difference of the variables are statistically significant (except only for Y_t in the PP test), leading to the rejection of the null hypothesis that the first differences are non-stationary. Therefore, all the variables are best described as I(1) variables.

Results for the Zivot and Andrews (1992) structural change unit root tests based on the levels in the data are reported in Table 3. Reference is made to the type of model, the break point suggested by the estimation method, and the minimum $t_{\hat{\alpha}}$ (test statistic on the lagged dependent variable). The AIC is employed to choose the best model for each series. At the 5% level of significance, both E_t and Y_t are stationary after recognizing structural changes in 1983 and 1984 respectively. While at the 10% level of significance both P_t and M_t are stationary after recognizing structural changes in 1974 and 1984 respectively. In contrast to the non-stationarity conclusions from the ADF and PP tests for the levels data, these results suggest that stationarity is apparent once structural changes are recognized. The Zivot-Andrews test results for data in first differences are presented in table 4 and as expected, consistent with the ADF and PP tests, suggest that all variables are stationary in first differences after recognising structural change.

For the levels data, the break for energy prices coincides with the first oil price shock. The break points for energy consumption, income and structural variation coincide with the lagged impact of the second oil shock occurring in 1979/80. Thailand experienced a number of economic instability problems as a result of the second oil shock. Rising world and US interest rates caused capital outflows from the country, and domestic overspending. On the external front the balance of payments deteriorated, international reserves declined sharply, and external debts accelerated. On the domestic front, consumer price inflation peaked in 1980 and slowed down considerably thereafter. Aside from restrictive fiscal and monetary policies that induced high interest rates and curbed the capital outflow problem, the Baht was devalued in 1981.

Given the results from the ADF and PP tests that all variables, E_t, Y_t, P_t, and M_t are *I(1)*, we initially employ the Engle-Granger residual based approach for cointegration testing using eqn (1). The residual based test statistics from this relationship are presented in table 5. The estimated ADF statistics are statistically significant at the 1 % level in all three types of processes whereas the estimated PP statistics are statistically significant at the 1 % level only in the process without trend and drift. Hence, in general the null hypothesis of non-stationarity of the residual can be rejected and it can be concluded that a cointegrating relationship exists.

To investigate the possibility of multiple cointegrating relationships we next employed the Johansen VAR approach. It is necessary to determine an appropriate lag length to use in the VAR model. The AIC and SBC statistics suggest optimal values of six lags and two lags, respectively. The values of SBC decline continuously until two lags. For reasons of parsimony we choose the VAR model with two lags.

The Johansen trace test and maximal eigenvalue tests are presented in table 6. The trace statistic indicates that at least one cointegrating vector ($r \geq 1$) exists in the system at the 95 pre cent confidence level (estimated LR statistic, 50.079 > 47.210, 95 % critical value). In order to cross check for identifying the specific number of cointegrating vectors, the maximal eigenvalue statistic is further employed. This statistic confirms the existence of only one cointegrating relationship at the 95 % confidence level in this system (estimated LR statistic, 28.640 > 27.067, 95 % critical value).

To examine structural changes in the cointegrating relationship we report the minimum value in the ADF series for the three model types as advocated by Gregory and Hansen. The ADF and break points for the models were: model C (-5.20*, 1973); model C/T (-5.43*, 1962) and model C/S (-5.38, 1968). Tests indicate (denoted by *) that the null hypothesis of no cointegration amongst the variables is rejected at the 10 % significance level for the level shift (C) and level shift with trend (C/T) models. The significant breaking periods are located in different time periods: 1973 for the first model and 1962 for the second model. In contrast, we can not reject the hypothesis of no cointegration for the regime shift (C/S) model. In general these results suggest that some, but not overwhelming, support for cointegration does exist even given the recognition of structural changes.

The estimated cointegration relationship (eqn(1)) is:

$$\begin{aligned}\hat{E}_t &= \underset{(-0.06)}{-269.6} + \underset{(14.9)}{0.504}Y_t - \underset{(-4.18)}{125.7}P_t + \underset{(5.91)}{898.9}M_t\end{aligned} \tag{3}$$

The t-statistics are presented in parentheses, all slope regressors are significant at the 1% level and $R^2 = 0.990$. To estimate the ECM, the AIC is applied to identify the optimal combination of lag (*m*, *n*, *o* and *p*) in eqn (2). The results of the preferred estimated regression of the ECM are presented in Table 7, the associated diagnostic tests (Table 8) indicate the absence of any particular regression problems.

In the ECM, the speed of adjustment coefficients suggests that within one year, about 58% of the disequilibrium between the actual and long-run aggregate energy demand will be decreased. The estimated coefficients of all variables have *a priori* expected signs, but not all coefficients are statistically significant. The changes in aggregate energy demand (ΔE_t) do not affect significantly the changes in E_t in the next year. Changes in income (ΔY_t), real energy prices (ΔP_t), and structural variation (ΔM_t) are all statistically significant at 1%, 5% and 10% level, respectively whereas their one-period lags are not. This implies that in the short run, the lag of those determinants does not significantly affect the changes in aggregate energy demand. In the short run, increases in income and structural variation will lead to an increase in the aggregate energy demand, whereas an increase in real energy prices will cause a reduction in aggregate energy demand.

Figures 1 and 2 present the results of the CUSUM stability test for the cointegrating regression model and the ECM. The tests indicate that the estimated lines of CUSUM of both models lie inside the critical values at the 5 % significance level. As a result, both models are appear to stable over time.

Next we plot the recursive estimates of each coefficient in the models to examine the stability of individual coefficients. The recursive coefficient estimates enable us to trace the evolution of estimates for any coefficient as more and more of the sample data are used in the estimation. If the coefficient displays significant variation as more data is added to the estimating equation, it is a strong indication of instability. Figures 3 and 4 report the results of the recursive coefficient estimates within their 95% confidence bands. In general, most coefficients appear to be reasonably stable. However, there appears to be some mild evidence of individual parameter instability, especially for the pre-1980 period and the Y_t coefficient in the cointegrating relationship, and ΔP_t and the disequilibrium error in the ECM.

The coefficient for the residual from the cointegration relationship is significant and has the expected sign in the estimated ECM.. Its negative sign suggests that when aggregate energy demand is higher than its long-run equilibrium value, there will be a mechanism to reduce actual aggregate energy demand in the next period so that aggregate energy demand is forced back toward the equilibrium. The estimate suggests that, within one year, about 58% of the disequilibrium gap between the actual and long-run aggregate energy demand will be decreased. Therefore, aggregate energy demand and its determinants have a stable long-run relationship.

The estimated ECM of energy demand can be used to explain the short-run movement of the energy demand in Thailand. The estimated coefficients of all variables have the *a priori* expected signs, but not all coefficients are statistically significant. The changes in aggregate energy demand (ΔE_t) do not affect significantly the changes in E_t in the next year. Changes in income (ΔY_t), real energy prices (ΔP_t), and structural variation (ΔM_t) are all statistically significant at 1%, 5% and 10% level, respectively whereas their one-period lags are not. This implies that in the short run, the lag of those determinants does not significantly affect the changes in aggregate energy demand. In the short run, increases in income and structural variation will lead to an increase in the aggregate energy demand, whereas an increase in real energy prices will cause a reduction in aggregate energy demand.

Evaluated at the means of the data, the results from the cointegrating regression imply long-run income, price, and structural variation elasticities of 0.568, -0.600 and 1.046

respectively. This indicates an inelastic demand in income and prices and elastic in structural variation. The results from the ECM imply a higher short-run income elasticity (0.788) but lower short-run price and structural variation elasticities (-0.522 and 0.491 respectively).

The ECM and cointegration relationship can be usefully employed to generate future energy demand projections for the period of 2006-2009. Assumptions for forecasting are made regarding major explanatory variables. We assume that the income will grow by 4.5 % per year (as assumed in the Ninth National Economic and Social Development Plan) and the share of manufacturing's output in GDP will increase by 1.9 percent (this is the average historical growth rate in manufacturing output share). Since our projections crucially depend on assumptions about world oil prices, we consider three alternative scenarios. This exercise has obvious merits in that world energy markets do face much uncertainty into the forecast period. The optimistic (scenario 1), moderate (scenario 2) and pessimistic (scenario 3) views assume real energy price increases of 2%, 4% and 6% respectively.

As shown in Figure 5, aggregate energy demand is projected to continue its prevailing upward trend as it approaches the year 2009. Our models forecast that the aggregate energy demand will reach the level of 81608 KTOE, 78606 KTOE and 75036 KTOE, respectively for the three scenarios, by the year 2009. The average annual growth rates for energy demand for the three scenarios are: 3.72%, 3.35% and 2.81% respectively. This highlights the need of an effective management policy in preparing plan for future energy development and ensuring the energy supply security.

4. Conclusion

This paper has contrasted results from standard tests for analysing energy demand with results which recognize the possibility of structural change. In general, the recognition of structural change has lead to some unique insights. Conventional unit root tests suggest that the variables are integrated of order one, however, Zivot-Andrews tests suggest that unit roots do not exist in the levels data once structural changes are recognised. Conventional cointegration tests suggest the presence of a cointegration relationship of order one, the Gregory-Hansen procedure confirms a cointegration relationship but only at a 10% level of significance. CUSUM tests indicate that the cointegrating relationship and the ECM appear to have stable coefficients, however closer inspection of individual parameter stability suggests instability for some coefficients in both the cointegration relationship and the ECM.

The elasticities stemming from the energy demand model are interesting and have important implications. Our estimates suggest the effect of a change in income on energy demand is greater in the short-run than in the long-run. This may follow from the inflexibility of firms' and households' energy-using capital and appliance stocks in the short-run. An increase in income will, therefore, bring about an immediate increase in the derived demand for energy in the short-term, but this derived demand is reduced in the longer term as more energy efficient machines are installed.

The effect of a change in the real price of energy is less in the short-run than in the long-run. This may also reflect the fixed nature of machine and appliance stocks in that a rise in the real price of energy produces a modest fall in energy consumption in the short-term. Energy consumption falls further in the longer term, however, as the price increase induces the installation of more energy-efficient domestic appliances and capital goods.

The low price elasticity for aggregate energy demand, even in the long run suggests that the over-pricing of energy as a policy instrument is not likely to be very influential on future energy demand. Additionally, taxes on energy prices are unlikely to achieve government goals for energy conservation and environmental improvement, although they may well be efficient for revenue raising. This indicates that the energy pricing policy is ineffective for energy demand management and for reducing dependency on energy sources from foreign countries.

The demand for energy is elastic in the long-run but inelastic in the short-run for structural variation. High levels for the long-run structural variation elasticity indicates that demand for energy in Thailand is likely to increase quite sharply as the share of manufacturing output in GDP increases. We offer two explanations. First, structural variation is measured as the share of manufacturing output in GDP and manufacturing has long been one of the major energy consumers of the country; for example, in 1999 manufacturing consumed around 34 % of total energy. Second, as a result of the industrialization of the Thai economy, the production process in manufacturing has become more capital-intensive. Under these circumstances, there may be a considerable degree of substitution of capital for labor. An increase in energy machine and power installation will, therefore, give rise to an increase in the derived demand for energy in the longer term.

The long-run elastic structural variation elasticity of energy demand implies that the changing industrial structure has involved a shift away from light manufacturing, with its low energy requirement toward heavy manufacturing industries, which demand relatively large amounts of energy. Consequently, the structural changes during the overall period 1956-1999, working in favor of industries with more energy intensive production, are likely to generate proportional increases in energy demand and hence environmental degradation. The enactment of energy policies to induce technical change toward more energy efficient production processes would be especially beneficial. It would be important to encourage a shift in the industrial mix with more output originating in the low energy intensive industries.

References

Asafu-Adjaye J. 2000. The relationship between energy consumption, energy prices and economic growth: time series evidence from Asian developing countries. *Energy Economics*. **22**, 615-625.

Atkinson J, Manning N. 1995. A survey of international energy elasticities. In *Global Warming and Energy Demand*, Barker T, Ekins P, Johnstone N.; Routledge: London.

Brown RL, Durbin J, Evans JM. 1975. Techniques for testing the constancy of regression relationships over time. *Journal of the Royal Statistical Society*, B. **37,** 149-192.

Christiano LJ. 1992. Searching for a break in GNP. *Journal of Business and Economic Statistics*. **10**(3), 237-250.

Dickey DA, Fuller WA. 1979. Distribution of estimators for autoregressive time series with a unit root. *Journal of American Statistical Association*. **84**, 427-431.

Dickey DA, Fuller WA. 1981. Likelihood ratio statistics for autoregressive time series with a unit root. *Econometrica*. **49**, 1057-1072.

Engle RF, Granger CWJ. 1987. Cointegration and error correction: representation, estimation and testing. *Econometrica*. **55**, 251-276.

Gregory AW, Hansen BE. 1996. Residual based tests for cointegration in models with regime shifts. *Journal of Econometrics*. **70**(1), 99-126.

Harris RID, 1995. *Using Cointegration Analysis in Econometric Modelling*. Prentice Hall: New York.

Harvey AC. 1981. *The Econometric Analysis of Time Series.* Philip Alan: Oxford.

Hoa TV. 1993. Effects of oil on output growth and inflation in developing countries: the case of Thailand from January 1966 to January 1991. *International Journal of Energy Research*. **17**(1), 29-33.

Johansen S, Juselius K. 1990. Maximum likelihood estimation and inference on cointegration with application of demand for money. *Oxford Bulletin of Economics and Statistics*. **52**, 169-209.

Masih AMM, Masih R. 1996. Energy consumption, real income and temporal causality: results from a multi-country study based on cointegration and error-correction modelling techniques. *Energy Economics*. **18**, 165-183.

Newey WK, West KD. 1987. A simple positive semi-definite heteroscadasticity and autocorrelation consistent covariance matrix. *Econometrica*. **55**, 1029-1054.

Ng S, Perron P. 1995. Unit root tests in ARMA models with data-dependent methods for the selection of the truncation lag. *Journal of the American Statistical Association*. **90**, 268-281.

Oh W, Lee K. 2004. Causal relationship between energy consumption and GDP revisited: the case of Korea 1970-1999. *Energy Economics.* **26**, 55-59.

Osterwald-Lenum, M. 1992. A note with quantiles of the asymptotic distribution of the maximum likelihood cointegration rank test statistics. *Oxford Bulletin of Economics and Statistics*. **54**(3), 461–472.

Perron P. 1989. The great crash, the oil price shock, and the unit root hypothesis. *Econometrica*. **57**(6), 1361-1401.

Phillips PCB, Perron P. 1988. Testing for a unit root in time-series regression. *Biometrica*. **75**, 335-346.

Zivot E, Andrews DWK 1992. Further evidence on the great crash, the oil-price shock, and the unit-root hypothesis. *Journal of Business and Economic Statistics*. **10**(3), 251-270.

Data Appendix

Summary data of the patterns of energy and national growth are from: Department of Energy Development and Promotion (DEDP) Ministry of Science, Technology and Environment, *Thailand Energy Situation*, various issues and National Economic and Social Development Board (NESDB), Office of the Prime Minister, *National Income of Thailand*, various issues, and the National Energy Policy Office (NEPO).

Annual data covering the period 1956 to 1999 is employed for analysis. Energy demand (E_t) is sourced from the DEDP, and the National Energy Administration (NEA), Ministry of Science, Technology and Environment. Income (Y_t) is soured from the NESDB. Energy Prices (P_t) is sourced from the Department of Commercial Economics, Ministry of Commerce. The indicator of structural variation (M_t) is calculated from data in NESDB, *National Income of Thailand*.

Table 1: Augmented Dickey-Fuller (ADF) Tests: Thai Energy Demand

Variables	Levels			First Differences		
	W/trend & drift	w/o trend w/drift	w/o trend &drift	w/trend & drift	w/o trend w/drift	w/o trend &drift
E_t	0.163 [2]	2.722*[2]	4.017***[2]	-6.065***[1]	-2.775*[2]	-1.885*[2]
Y_t	-0.259 [2]	1.401 [2]	1.905*[2]	-3.710**[1]	-2.661*[1]	-2.034**[2]
P_t	-2.750 [1]	-3.014**[1]	-1.015 [2]	-4.653***[1]	-4.458***[1]	-4.425***[1]
M_t	-1.573 [1]	1.434 [1]	4.083***[1]	-4.879***[1]	-4.498***[1]	-2.759***[1]

Notes: Figures in square brackets represent the number of lagged dependent variables used in the ADF regression which is selected by minimizing AIC.
*** indicates significant at a 1% level
** indicates significant at a 5% level
* indicates significant at a 10% level

Table 2: Phillips and Perron (PP) Tests: Thai Energy Demand

Variables	Levels			First Differences		
	W/trend & drift	w/o trend w/drift	w/o trend &drift	w/trend & drift	w/o trend w/drift	w/o trend &drift
E_t	-1.695 [3]	0.529 [3]	2.143**[3]	-4.109***[3]	-4.023***[3]	-3.528***[3]
Y_t	-0.807 [3]	1.994 [3]	3.647***[3]	-2.002 [3]	-2.015 [3]	-1.664*[3]
P_t	-2.438 [3]	-2.964**[3]	-1.214 [3]	-4.128**[3]	-4.008***[3]	-4.020***[3]
M_t	-2.052 [3]	1.871 [3]	4.669***[3]	-7.121***[3]	–6.672***[3]	–5.052***[3]

Notes: Figures in square brackets represent appropriate lag truncation for Bartlett Kernel as suggested by Newey-West (1987).
*** indicates significant at a 1% level
** indicates significant at a 5% level
* indicates significant at a 10% level

Table 3: Zivot and Andrews Unit Root Tests (at Levels):Thai Energy Demand

Variable	**AIC Model A**	**AIC Model B**	**AIC Model C**	**Best Model**	**k**	**Min $t_{\hat{\alpha}}$**	**Break Point**
E_t	18.074	17.809	**17.764**	C	8	-5.31**	1983
Y_t	16.656	**16.571**	16.613	B	6	-5.07***	1984
P_t	5.320	**5.061**	5.118	B	8	-4.17*	1974
M_t	2.180	2.232	**2.029**	C	4	-4.82*	1984

Notes: *** indicates significant at a 1% level
** indicates significant at a 5% level
* indicates significant at a 10% level

Table 4:
Zivot and Andrews Unit Root Tests (at First Differences): Thai Energy Demand

Variable	**AIC Model A**	**AIC Model B**	**AIC Model C**	**Best Model**	**k**	**Min $t_{\hat{\alpha}}$**	**Break Point**
ΔE_t	**17.934**	18.332	17.939	A	1	-8.12***	1987
ΔY_t	16.659	**16.506**	16.572	B	8	-5.86***	1986
ΔP_t	**5.421**	5.606	5.468	A	1	-5.93***	1985
ΔM_t	**2.295**	2.373	2.333	A	0	-7.60***	1985

Note: *** indicates significant at a 1% level

Table 5: ADF and PP Tests of Estimated Residuals

ADF Statistic			**PP Statistic**		
w/trend & drift	**w/o trend w/drift**	**w/o trend &drift**	**w/trend & drift**	**w/o trend w/drift**	**w/o trend &drift**
-4.263***[1]	-4.245***[1]	-4.353***[1]	-2.633 [3]	-2.703*[3]	-2.776***[3]

Notes: Figures in square brackets in the ADF statistic column represent the number of lagged dependent variables used in the ADF regression which is selected minimizing AIC, while figures in square brackets in the PP statistic column represent appropriate lag truncation for Bartlett Kernel as suggested by Newey-West (1987).
*** indicates significant at a 1% level
** indicates significant at a 5% level
* indicates significant at a 10% level

Table 6: Johansen Multivariate Cointegration: Thai Energy Demand

Hypotheses H_0	H_1	Eigenvalue	Test Statistics Trace	Max Eigenvalue
r = 0	r ≥ 1	0.503	50.079**	28.640**
r ≤ 1	r ≥ 2	0.237	21.439	11.064
r ≤ 2	r ≥ 3	0.159	10.375	7.120

Notes: *r* indicates the number of cointegrating relationships.
Critical values are tabulated in Osterwald-Lenum (1992).
** indicates significant at a 5% level

Table 7: Estimated ECM of Energy Demand

Regressor	Dependent Variable: ΔE_t Parameter Estimate
Constant	-77.123 (-0.291)
ΔE_{t-1}	0.131 (0.919)
ΔY_t	0.699 (4.368)***
ΔY_{t-1}	0.430 (1.486)
ΔY_{t-2}	-0.494 (-2.865)***
ΔP_t	-109.485 (-2.287)**
ΔM_t	422.085 (1.728)*
ΔM_{t-1}	-393.720 (-1.403)
$\hat{u}_{t-1}$	-0.584 (-3.586)***
R-squared	**0.858**
D-W statistic	**2.018**
F-statistic	**24.248**

Notes: Figures within parentheses show t-statistics.

*** indicates significant at a 1% level
** indicates significant at a 5% level
* indicates significant at a 10% level

Table 8: Diagnostic Checking for the Estimated ECM of Energy Demand

Type of Diagnostic Tests	Computed Value	p-value
Serial Correlation: Breush-Godfrey LM Test[1/]	3.162	0.531
Heteroscedasticity: ARCH Test [2/]	0.482	0.487
Normality: Jarque-Bera LM Test[3/]	1.101	0.577
Misspecification: Ramsey RESET Test[4/]	1.021	0.319

Notes: Distributional properties of diagnostics are:
[1/] LM4 [$\chi^2(4)$] test for the fourth order serial correlation amongst the residuals.
[2/] ARCH1 [$\chi^2(1)$] test for first-order autoregressive conditional heteroscedastic effects.
[3/] LM2 [$\chi^2(2)$] test for normality of residuals.
[4/] Regression specification Error F-test with (1, 32) degrees of freedom.

Figure 1: The Plot of CUSUM Test for the Cointegrating Regression Model

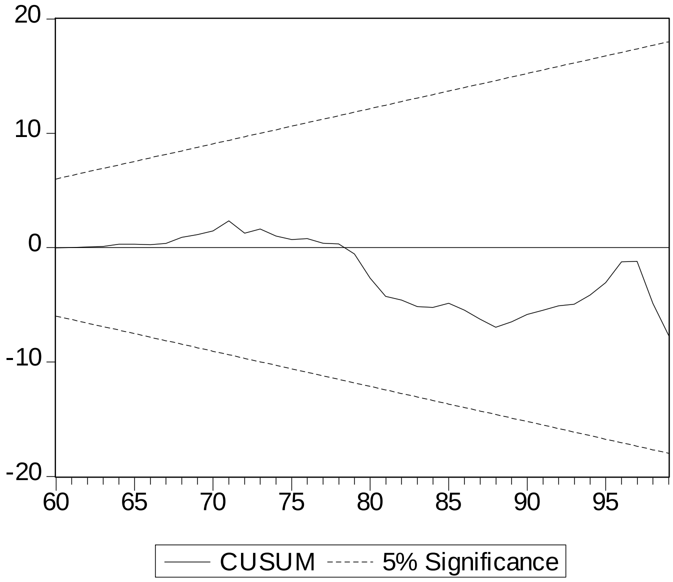


Figure 2: The Plot of CUSUM Test for the ECM

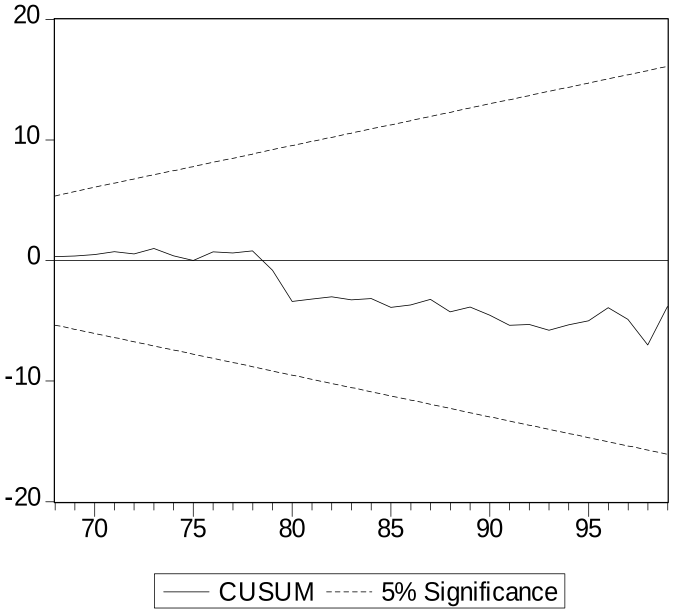
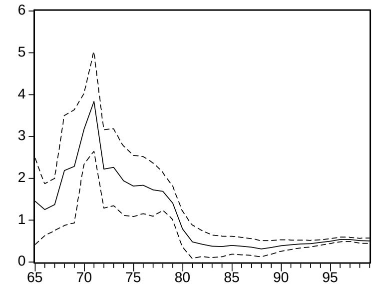
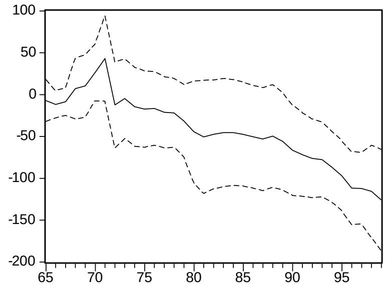


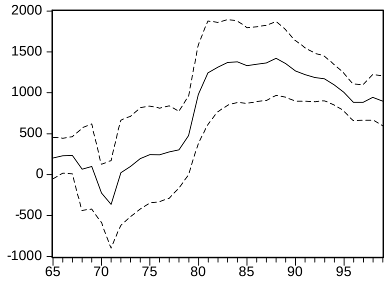
Figure 3:
Recursive Estimates for Coefficients in the Cointegrating Regression Model



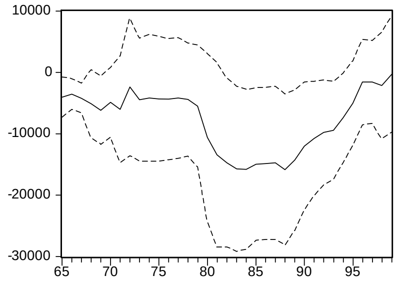
Y_t's Coefficient



P_t's Coefficient

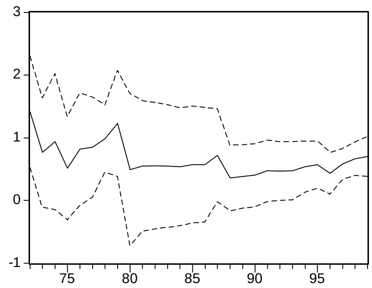


M_t's Coefficient

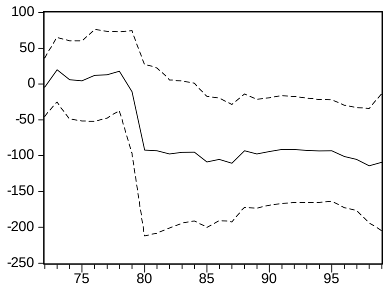


Constant Term

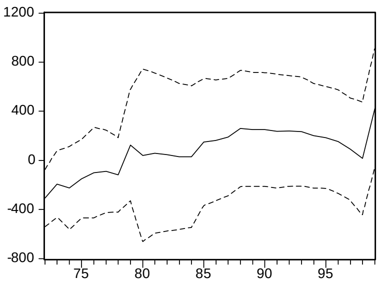
Figure 4: Recursive Estimates for Coefficients in the ECM



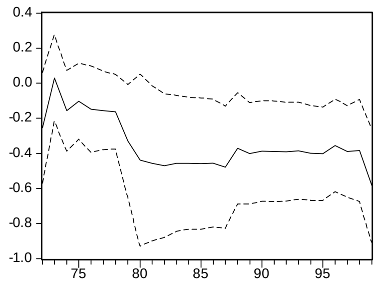
ΔY_t's Coefficient



ΔP_t's Coefficient



ΔM_t's Coefficient



Disequilibrium Error