

RESIDENTIAL ELECTRICITY DEMAND IN TAIWAN

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This paper examines the residential demand for electricity in Taiwan as a function of household disposable income, population growth, the price of electricity and the degree of urbanization. Short- and long-term effects are separated through the use of an error correction model. In the long run, the income elasticity is unitary elastic. The own-price effect is negative and inelastic. In an error-correction framework, the short run income and price effects are small and less than the long-run effects. We have used a proxy variable, urbanization, to capture economic development characteristics and electricity-using capital stocks not explained by income. The variable provides significant explanatory power to the model both in the short- and long-run. We have not seen this variable used in previous studies and we think it will be useful in explaining residential electricity consumption in other developing countries.

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This paper examines the residential demand for electricity in Taiwan as a function of household disposable income, population growth, the price of electricity and the degree of urbanization. Short- and long-term effects are separated through the use of an error correction model. In the long run, the income elasticity is unitary elastic. The own-price effect is negative and inelastic. In an error-correction framework, the short run income and price effects are small and less than the long-run effects. We have used a proxy variable, urbanization, to capture economic development characteristics and electricity-using capital stocks not explained by income. The variable provides significant explanatory power to the model both in the short- and long-run. We have not seen this variable used in previous studies and we think it will be useful in explaining residential electricity consumption in other developing countries.

Introduction

In the forty years spanning the period 1955 through 1995, residential consumption of electricity in Taiwan increased from 257 GWh to 26,144 GWh. During the same period the economy underwent a rapid process of economic development and modernization. Household disposable income increased from 22,438 to 5,003,970 NT\$ million. In real per capita terms that is equivalent to a ten-fold increase. The Taiwanese economy was predominantly rural in the 1950s, while today it relies on a world class manufacturing sector to produce technology-intensive goods. The proportion of the population living in cities has grown by more than 150% since the early 1950s. Nearly 60% of the population live in cities with populations greater than 100,000 inhabitants. The purpose of this paper is to examine the demand for residential electricity during this period of economic development.

The Taiwan Power Company (TPC) was established on May 5th, 1946, as the sole electric utility for Taiwan. Initially the system had a capacity of 275 MW from 28 power plants. At that time, 80% of the power was generated by 21 hydroelectric plants and the remaining 20% percent came from 7 thermal plants (coal, oil, and LNG). Today, the total installed capacity is 26,679 MW from 69 power plants. The share of hydroelectricity has fallen to 16.7%, thermal plants supply about 63%, and three nuclear plants provide the remaining 20%. The share of petroleum fired electricity generating plants has declined since the mid 1970s. Total generating capacity is expected to reach 42,000 MW in 2006, an increase of nearly 60% as compared with today.

In 1996, Taiwan embarked on a privatization policy of government-owned enterprises. TPC is slated to become a private enterprise in 2001 (Taiwan Power Company, Annual Report 1997). The utility will continue to hold licenses for all existing plants for ten years, but independent power producers are expected to provide a significant amount of the new planned capacity. At the end of 1997, TPC had already signed five power purchase agreements with total capacity of 5,200 MW (nearly a third of the expected increase in capacity by 2006).

There have been numerous empirical studies of electricity demand in industrialized (Western)

countries. However, energy demand models from a developing country perspective may require a different framework. One potential difference is that economic growth and structural change associated with rapid development suggest that income and price elasticities will not be stable. This study modifies the electricity demand model traditionally used in the literature to the case of Taiwan

Electricity demand studies such as this one have important practical applications. The estimation of consistent and stable price- and income elasticity estimates are important. Government planners and private investors need to be well informed regarding the value of the market in privatization programs. Admittedly, aggregate level and low frequency data are less preferred to their counterparts, but it is precisely the latter data that are the most difficult to obtain in developing countries. The estimates obtained in this study can therefore be viewed as providing good rules of thumb in the initial stage of investigation.

The paper is organized into five sections. The first section discusses the modification of typical electricity demand models to a developing country case. It attempts to explain both short-run and long-run behavior. Second, we present the data used in the study. Third, the econometric issues and time series properties of the data are examined. The model of electricity demand and econometric results are presented next, while the fifth section summarizes the findings of the paper.

I. A General Model of Electricity Demand for Developing Countries:

I.1 A Brief Review of Residential Electricity Demand Models

Residential electricity consumption can be expressed in general as a function of the stock of electrical energy using equipment, K_t , and economic factors X_t .

$$Kwh_t = F(X_t, K_t(X_t)) \quad (1)$$

The two components can have independent and interdependent impacts on electricity demand. The capital stock of energy using equipment can be broken down into two types. The first relates to the

demand for daily energy services: lighting, refrigeration, cleaning, and entertainment. The second relates to seasonal weather patterns which can affect the demand for heating and cooling services.

The dependence of capital stock on economic factors only holds in the medium to long-run. This is because in the long run, the stock of appliances is flexible and can respond to changes in relevant prices. In the short run, the demand for electricity will be constrained to changes in the utilization rates given the fixed stock of electricity-using appliances.

Fisher and Kaysen, (1962) suggested a two-stage model where consumption in the short run (the first stage) depends on two components: income, Y_t , and the price of electricity, PE_t , and is given by

$$Kwh_t = u_t * K_t = u_t(Y_t, PE_t) * K_t. \quad (2)$$

The component u_t is the utilization rate(s) of the appliance stocks. In the second stage (the long run), Fisher and Kaysen tried to explain the factors affecting the capital stock. Their (saturation) model used the growth rate in appliance stocks regressed on population, expected income, marriages, expected energy prices, and the number of wired households. The measurement of capital stock was problematic. Fisher and Kaysen issued warnings that the quality of the data ranged "from somewhat below the sublime to a bit above the ridiculous." and "No results can be better than the data on which they are based."(p.27).

Taylor, Blattenburger, and Rennhack (1984) summarized the results from research using the two stage framework and found that there was an improvement in the results relative to a static model. However, they noted that until data on appliance stock improved, additional research would be hampered.

The estimation problems associated with the two-stage approach led researchers to develop an alternative procedure which avoided the use of equipment stock. In this alternative approach a distinction is made between actual consumption, Kwh_t , and long-run desired or equilibrium consumption, Kwh_t^* . The latter is affected by the level of income, relative prices and other factors.

To empirically implement the procedure, a partial adjustment framework is specified. There is a correspondence between the equipment stock and the desired consumption: Since actual equipment stocks will seldom be at their long-run equilibrium levels, actual consumption will differ from desired consumption. Consumers are hypothesized as attempting to bring actual consumption in line with their desired level. They are only partially able to accomplish this in any time period. Nelson, Peck, and Uhler (1985 and 1989) provide excellent examples of this time series econometric approach.

1.2 A Modified Model of Residential Electricity Demand in Developing Countries

The problem of correctly capturing electricity-using capital stock is magnified when working with developing countries, since obtaining estimates of the quantities and vintages of electricity-using equipment is more difficult. We propose an alternative means for conditioning demand on the capital stock. We believe that the degree of urbanization is a reasonable proxy for electricity-using equipment since cities are electrified sooner than rural areas and are on the forefront of adopting modern household appliances. We can rewrite (1) as follows:

$$Kwh_t = F(X_t, Urban_t(X_t)) \quad (3)$$

where X represents disposable income, energy prices, and the number of customers. Thus we modify the Fisher Kaysen model using the following specification:

$$Kwh_t = u_t * K_t = u_t(Y_t, PE_t, Urban_t) * Urban_t(Y_t, PE_t). \quad (4)$$

In this paper, a general model for electricity demand in developing countries will be formulated as follows:

$$Kwh = f(\text{Population}, \text{Income}, \text{Price/Kwh}, \text{Price of Oil}, \text{Urbanization}) \quad (5)$$

where the dependent variable *Kwh*, is residential electricity consumption; *Income*, is real disposable income; *Price/Kwh*, is the real price of electricity per KWh; *Price of Oil*, is the world price of oil and is used as a proxy for substitutes for electrical energy; *Urban* is the degree of urbanization, expressed as the percentage of people living in urban areas with 100,000+ population. We have not seen this variable or other proxies for economic development used in electricity demand models. Our model does not include weather variables for cooling degree or heating degree days. While the former is probably important for Taiwan, we could not obtain sufficient data for modeling purposes.

Higher disposable income is expected to increase consumption through greater economic activity and purchases of electricity-using appliances in the short- and long-run, since electricity is a normal good (service). A rise in electricity prices *ceteris paribus* will lead to a fall in consumption demanded. Economic theory suggests that electricity purchases will depend on the prices of its substitutes: natural gas and distillate fuel oil. However, the independent influences of alternate fuel prices may be rather small because of the response(s) by suppliers and the constraints on consumers to fuel switching. Natural gas is not appropriate in the case of Taiwan since its consumption has been comparatively small and recent. Thus, we follow Kokklenberg and Mount (1993) and take fuel oil prices as our measure of alternative prices.

Urbanization, our measure of economic development, is expected to increase the consumption of electricity. The reason is twofold. First, urbanization implies greater access to electricity, since households can be more easily connected to the grid. Second, those consumers that already had access to electricity before moving are likely to increase their consumption once they arrive in an urban setting. The increase in demand will take place through increased use of existing appliances (the utilization rate), and through the purchase of new ones (increased stock of capital-using equipment). Examples of purchases include air-conditioning and modern kitchen appliances, both of which may result from increased exposure to the media and advertising typical of large cities.

The specification of the model above treats the price of electricity, disposable income, and urbanization as exogenous variables. However, it is possible that these variables are jointly

determined. Our empirical approach tests whether these variables can be treated as weakly exogenous.

II The Data

Figure 1 reveals the dramatic (exponential) growth in residential electricity consumption. In 1955, consumption was about 250 GWh; by 1995 it had grown 100 times to over 25,000 GWh. Per capita consumption grew by nearly 45 times from 28 thousand KWh to 1,264 thousand KWh (see Figure 2).

The degree of urbanization, measured as the per cent of the population living in cities with populations greater than 100,000, demonstrates the large structural changes taking place in the country during this period. Figure 3 shows that in 1955, about 23% of the population lived in urban areas. This proportion had increased to nearly 60% by 1995; the fastest migration occurred from 1967 to 1977 when the proportion grew from 30% to over 45%. Real disposable income per capita grew nearly ten times between 1955 and 1995, from 19,000 NT\$ to 196,000 NT\$. Figure 4 shows a plot of disposable income. It took the first 15 years for real disposable income to double. During the first three years of the 1970s income increased a further 10,000 NT\$. However, the oil crisis and global recession in 1974-75 led to a decline in income by more than 5,000 NT\$ per person, about 10% of total disposable income. After that the economy grew at a fairly steady pace with per capita income increasing from 50,000 NT\$ to 190,000 NT\$ by 1995.

Figure 5 shows the price per kilowatt in real and nominal values. Prices fell in real terms by more than 50% from 1955 to 1979, including during the oil shock of 1974. At that time, more than 60% of the thermal generating capacity was from oil-fired systems. Taipower has since changed the fuel mix so that today less than 20% is from oil-fired systems (Tien-Sung Chien, 1997). In 1979 and 1980, the nominal price nearly doubled and the real price increased by more than 50%. This seems to have been part of the government's conservation measures taken in response to the sharp rise in international energy prices in the 1970s. Since 1982 the real price of electricity has steadily declined,

while the nominal price has stayed between 2.0 -2.5 NT\$ KWh (7-8 US cents). Real and nominal oil prices are plotted in Figure 6. The relative price of electricity to oil is plotted in Figure 7 (note that it does not include import tariffs or taxes). Initially, the price of electricity rose in relative terms until 1970, when it fell dramatically, by 66%. From then on the two prices have stayed close to parity.

III. Econometric Modeling Issues and Results

We employ the general-to-specific modeling approach advocated by Hendry (1986). The general-to-specific modeling approach is a relatively recent strategy used in econometrics. It attempts to characterize the properties of the sample data in simple parametric relationships which remain reasonably constant over time, account for the findings of previous models, and are interpretable in an economic and financial sense. Rather than using econometrics to illustrate theory, the goal is to "discover" which alternative theoretical views are tenable and test them scientifically.

The approach begins with a general hypothesis about the relevant explanatory variables and dynamic process (i.e. the lag structure of the model). The general hypothesis is considered acceptable to all adversaries. Then the model is narrowed down by testing for simplifications or restrictions on the general model. Below we consider a linear dynamic single equation model. The autoregressive distributed lag (ADL) model begins with a regression of the variable of interest on lagged values of itself and current and lagged values of the explanatory variables.

$$z_t = \alpha_0 + \sum_{i=1}^p \alpha_i z_{t-i} + \sum_{j=0}^m \beta_j x_{t-j} + \epsilon_t$$

In this case, the model is referred to as an ADL(p,m), because it contains p lagged dependent variables and m lagged explanatory variables. In the single equation case, we are implicitly assuming a conditional model where the x variables are assumed to be (weakly) exogenous. However in our model, z_t is a vector which includes electricity consumption per capita, the real price of electricity, real disposable income per capita, and the degree of urbanization. The real world oil price is the x

variable in our model. The constant term, α_0 , can include trend terms, (seasonal variables,) and dummy variables. At this point we have not made any assumptions about the order of integration for the z and x variables. The objective of specifying the general model is that it can encompass several (competing) hypotheses. Bentzen and Engsted (1993) develop the rationale for the use of cointegration analysis for energy demand.

The first step involves examining the time series properties of the individual data series. We look at patterns and trends in the data and test for stationarity and the order of integration. Second, we form a Vector Autoregressive Regression (VAR) system. This step involves testing for the appropriate lag length of the system, including residual diagnostic tests and tests for model/system stability. Third, we examine the system for potential cointegration relationship(s). Data series which are integrated of the same order may be combined to form economically meaningful series which are integrated of lower order. Fourth, we interpret the cointegrating relations and test for weak exogeneity. Based on these results a conditional error correction model of the endogenous variables is specified, further reduction tests are performed and economic hypotheses tested.

III.1 Time Series Properties of the Individual Series

Campbell and Perron (1991) provide rules (of thumb) for investigating whether time series contain unit roots. To begin, we estimate the following three forms of the augmented Dickey-Fuller (ADF) test where each form differs in the assumed deterministic component(s) in the series:

$$\Delta z_t = \alpha_1 z_{t-1} + \sum_{i=1}^P \gamma_i \Delta z_{t-i} + \varepsilon_t \quad (6)$$

$$\Delta z_t = \alpha_0 + \alpha_1 z_{t-1} + \alpha_2 * t + \sum_{i=1}^P \gamma_i \Delta z_{t-i} + \varepsilon_t \quad (7)$$

$$\Delta z_t = \alpha_0 + \alpha_1 z_{t-1} + \sum_{i=1}^P \gamma_i \Delta z_{t-i} + \varepsilon_t \quad (8)$$

The ε_t is assumed to be a Gaussian white noise random error; and $t=1,\dots,n$ (the number of observations in the sample) is a term for trend. In equation (6) there is no constant or trend. Equation (7) contains a constant but no trend. Both a constant and a trend are included in (8). The number of lagged differences, p , is chosen to ensure that the estimated errors are not serially correlated.

The results from the unit root tests are found in Table I. The first three rows test the null hypothesis that a series follows a unit root process or random walk. This implies it is non-stationary and (possibly) integrated of order one, $I(1)$, rather than $I(0)$. The second three rows test the null hypothesis that second differencing of a series follows a unit root. If true, the researcher must difference the series twice to obtain a stationary process.

We find that for all series in Table I the null hypothesis of a unit root in the first difference cannot be rejected. There is evidence that income per capita is stationary, $I(0)$, for the ADF regression including a constant and trend term, $I(1)$ -8. However further testing suggested that the model with only a constant was the appropriate choice, $I(1)$ -7. The slope coefficient of the trend term was insignificant. The tests for unit roots in the second differences are rejected, implying that the series are $I(1)$ stationary in their first differences.

III.2 The VAR System

The function for electricity demand is specified to reflect the economic determinants and developments from a predominantly rural economy to an urban industrial one.

$$LKwhpc_t = A_t \cdot LYdpc_t^{\beta_1} \cdot LRp_t^{\beta_2} \cdot LRPOil_t^{\beta_3} \cdot Urban_t^{\beta_4} \cdot e^{\varepsilon_t} \quad (9)$$

Where $LKwhpc_t$ is the quantity of electricity consumed in billions of kilowatt hours, $LYdpc_t$ is real disposable income in 1987 dollars, LRp_t is the real price of electricity per kilowatt hour, $LRPoil_t$ is the real world oil price, and $Urban_t$ is the population in cities of 100,000 or more. The A_t term can include deterministic elements like trend(s) and dummy variables.

In the econometric analysis, we will specify electricity consumption and disposable income in per capita terms. Also, the variables will be transformed to natural logarithms with the exception of the urbanization variable, since it is already in per cent. Plourde and Ryan (1985) note the pitfall of modeling in double log form with its impact on elasticities and specification of the utility function. We tried models in levels and semi-log form, but were unable to obtain useful results; this may be due to the non-stationarity of the data.

We specify the VAR as a four variable system with a maximum of three lags. The model includes the log of the real world oil price, a trend term and dummy variable. Between 1979 and 1981, there were sharp increases in the energy price variables. Based on correspondence with officials from Taipower and the Energy Commission of the Ministry of Economic Affairs, we constructed a dummy variable with a value of unity for those years and zero in all other years.

$$Z_t = \Pi_0 + \sum_i^p \Pi_i Z_{t-i} + \sum_i^m \Theta_i X_t + U_t$$

where

$$Z_t = \begin{bmatrix} Z_1 \\ Z_2 \\ Z_3 \\ Z_4 \end{bmatrix} = \begin{bmatrix} \text{Residential Kwh per Capita} \\ \text{Real Price of Electricity / Kwh} \\ \text{Real Disposable Income per Capita(NT\$ 1990)} \\ \text{Urbanization} \end{bmatrix} \quad (10)$$

where

$$X_t = \text{Real Price of Oil}$$

The selection criteria for the appropriate lag length use several criteria. The maximum possible lag length considered was three years. Since we are using annual data and a sample of 40 years, three years should capture the dynamic interactions between the dependent and independent variables.

We use the Wald F-test form of the test since we have a small sample. This compares the fit of the maintained hypothesis or restricted model against the null hypothesis or unrestricted model.

The maximum number of lags is specified and the model is estimated. Then, zero restrictions on the coefficients at lag three are tested. If the null hypothesis of no explanatory power is accepted, then a model with only two lags explains as well as one with three lags. That model is over-parameterized. The process is repeated for one lag. Now, the null hypothesis is that all the coefficients at two lags are zero and that a model with one lags explains as well as two lags.

The Bayesian Schwarz Criterion (BSC), the Akaike Information Criterion (AIC), and the Hannan-Quinn Criterion are used as alternative criterion for testing the appropriate lag length. They rely on information similar to the Chi-Squared tests and are derived as follows:

$$\begin{aligned} AIC &= \log(\text{Det } \hat{\Sigma}) + 2*c*T^{-1} \\ BSC &= \log(\text{Det } \hat{\Sigma}) + c*\log(T)*T^{-1} \\ HC &= \log(\text{Det } \hat{\Sigma}) + 2*c*\log(\log(T))*T^{-1} \end{aligned}$$

where $\text{Det } \hat{\Sigma}$ is the determinant of the covariance matrices of residuals. The number of observations is T , and c is a correction for the number of variables in the unrestricted equation. The adjustment is used to improve the small sample properties of the statistic. This is calculated as the number of estimated coefficients in all equations. If there are n variables (equations), p lags of each variable, and an intercept, we have $c = n^2 * p + n$ coefficients. Intuitively, the log determinant will decline as the number of regressors increases, just as in a single equation ordinary least squares regression. It is similar to the residual sum of squares or estimated variance. The second term on the right hand side acts as a penalty for including additional regressors; it increases the statistic.

We calculate these statistics for each lag length and choose the lag length based on the model(s) with the minimum value for the statistics. The three tests do not always agree the same number of lags. The AIC is biased towards selecting more lags than is actually needed. However, this is not necessarily a negative feature of the test since we are trying to remove any serial correlation.

Table II reports the results from the F and Related Statistic for the Sequential Reduction from a maximum of three lags of the endogenous and exogenous variables. We conclude that one lag of the endogenous variables and only the contemporaneous value of the exogenous variable provide the same explanatory power for the system as the models with more lags. The F-tests against models with three lags of all variables, three lags of the endogenous variables, and two lags of the endogenous variables are 0.79, 0.87, and 1.17 respectively. Similarly the BSC, HQ, and AIC statistics are the most negative using a VAR with a single lag in the endogenous variables and only the contemporaneous value of the exogenous variable.

III.3 Cointegration Testing

Cointegration tests are a multivariate form of integration analysis. Individual series may be $I(1)$, but a linear combination of the series may be $I(0)$. The error correction model provides a generalization of the partial adjustment model and permits the estimation of short-run and long-run elasticities.

The approach is based on the findings of Nelson and Plosser (1982), in which many macroeconomic and aggregate level series are shown to be well modeled as stochastic trends, i.e., integrated of order one, or I(1). Simple first differencing of the data will remove the nonstationarity problem, but with a loss of generality regarding the long run "equilibrium" relationships among the variables.

Engle and Granger (1987) solve this filtering problem with the cointegration technique. They suggest that if all, or a subset of, the variables are I(1), there may exist a linear combination of the variables which is stationary, I(0). The linear combination is then taken to express a long-run "equilibrium" relationship¹. The new series, ECM_t , the estimated residual from (3), represents the deviation between current and equilibrium levels:

$$ECM_t = LKwhpc_t - \alpha_0 - \alpha_1 LYdpc_t - \alpha_2 LRP_t - \alpha_3 LRpoil_t - \alpha_4 Urban_t \quad (11)$$

Series which are cointegrated can always be represented in an error correction model. The error correction model is specified in first differences, which are stationary, and represent the short run movements in the variables. When the ECM term is included in the model, the long-run, or equilibrium, relations are accounted for. For example, the growth rate of electricity consumption per capita would be a function of the growth rates in real disposable income per capita, real electricity prices, real world oil prices, the degree of urbanization, and the error correction term. Lags of the independent and dependent variables would be included to capture additional short- and medium term dynamics of electricity consumption. The advantage of the first difference model is that the specification is stationary so that estimation and statistical inference can be performed using standard statistical methods. The contemporaneous coefficients are interpreted as short run elasticities.

We begin by modeling all the variables in a VAR system

$$z_t = \Pi_0 + \Pi_1 z_{t-1} + \dots + \Pi_k z_{t-k} + \epsilon_t \quad (12)$$

In our case z is a $(p \times 1)$ vector of nonstationary I(1) variables, where $p=5$. The Π_i are $(p \times p)$ coefficient matrices at different lags. In this case we are only using a single lag of the variables, and

the real world oil price enters in restricted form as an exogenous variable. Π_0 is a (p*1) vector of constant terms; it includes a constant, a time trend, and a dummy variable for the 1979-1981 period. The vector of disturbances, ϵ_t , are assumed to be white noise. The system in levels can be transformed to one in differences and the error correction term without loss of generality.

$$\Delta z_t = \Pi_0 + \Gamma_1 \Delta z_{t-1} + \dots + \Gamma_{k-1} \Delta z_{t-k+1} + \Pi z_{t-k} + \epsilon_t,$$

$$\text{where } \Gamma_i = -I + \Pi_1 + \dots + \Pi_i ; \forall i=1, \dots, k-1, \quad (13)$$

$$\text{and } \Pi = -I + \Pi_1 + \dots + \Pi_k$$

The rank of the matrix for the level terms, Π , is of reduced rank when cointegrating relationships exists. In that case it can be partitioned as $\Pi = \alpha\beta'$ where α is the (p*r) matrix of speed of adjustment coefficients and β is the (p*r) matrix of cointegrating vectors or long-run relationships. Cointegration testing and the maximum likelihood estimation of β are derived from a series of regressions and reduced rank regressions.

Table III contains the results from the cointegration analysis. The first column gives the null hypotheses of more than p-r= 0, 1, 2, or 3 cointegrating vectors. When there are no cointegrating vectors, p-r=0 and the model in differences is appropriate. We compare our results against the updated values in Table 2 of Osterwald-Lenum (1992). The results from the maximum eigenvalue and trace statistics indicate the presence of a single cointegrating relation. There is marginal evidence of possibly a second vector using the trace statistic, but we could not find sufficient evidence of a stable relationship nor are we able to provide an economic justification for including it.

The implied cointegrating relationship is obtained from first row of the standardized beta eigenvector. It is interpreted as the desired level of per capita residential electricity consumption.

$$LKWhr_{pc} = 3.31 * Urban + 1.57 * LRYdpc - 0.15 * LRp + 0.18 * LRPoil \quad (14)$$

The tests for weak exogeneity reveal if the implied disequilibrium in the cointegrating relation feeds

back into a (set) of the current endogenous variable(s). Individually, $LKWhpc$, $Urban$, and $LRYdpc$ reject the null hypothesis of no feed-back. However, the real own price of electricity does appear to be weakly exogenous. Historically, the Taiwan Power Corporation has set prices according to economic and social/political criteria. These results imply that we can reduce the system to a three-variable VAR conditioning on the two energy price series.

Note that the electricity price and oil price coefficients are roughly equal and of opposite signs. We test for the relative price restriction along with the weak exogeneity of the real oil price and obtain the following reduction in the cointegration vector.

$$LKWhr_{pc} = 3.91 * Urban + 1.04 * LRYdpc - 0.16 * (LRp - LRpoil) \quad (15)$$

The chi square statistic with two degrees of freedom is 1.42 with a p-value of 0.42; thus we cannot reject the restriction. The income elasticity is slightly greater than unity. An increase of one percent in urbanization leads to nearly a four percent increase in residential electricity consumption. An increase in the relative price of electricity of one per cent leads to a decline in consumption of 0.16; demand is relatively inelastic. This estimate is reasonably close to the long run own price elasticity of demand in numerous other studies (see Bohi & Zimmerman (198?), Dahl (1993), Halvorsen (1975), Espey (1999), Nelson et al.(1985 and 1989).

Since the major focus of our paper is residential electricity demand and the possible relationship with urbanization, we test the restriction that income together with energy prices are jointly weakly exogenous. The chi-square statistic with three degrees of freedom is 5.32 and has a p-value of 0.15, and the null hypothesis of the joint weak exogeneity of $LRYdpc$ and $LRp-LRpoil$ cannot be rejected.. Thus, we will condition our model on these restrictions and reduce the VAR to a two-variable system.

IV. The Error Correction Model

The vector obtained in the cointegration analysis represents the long-run relationship between the

variables. To model the demand for electricity more generally, however, an Error-Correction Model is employed. The Error-Correction framework models the variables in differences and the coefficients on the differenced variables correspond to short-run elasticities. The model furthermore contains an error-correction term (ecm). This term is obtained from the long-run relationship and expresses deviations in electricity consumption from its long-run mean. The coefficient in front of the ecm term measures the speed of adjustment in current consumption to the previous disequilibrium demand value.

Boswijk and Franses (1992) emphasize that the dynamic specification can affect the size and power of the cointegration tests. We started by estimating a model in levels with three lags and used the Bayesian-Schwartz Criterion and Akaike Information Criterion to reduce the number of lags. The results suggested that a single lag was adequate. However, theoretical grounds and individual t-statistics suggested some additional explanatory power through three lags. This implies a model in first differences using only two lags. We retained two lags of each variable and then removed the insignificant regressors. The model in its most general form is as follows:

$$\Delta y_t = \alpha + \sum \delta_i \Delta y_{t-i} + \sum \beta_i \Delta x_{t-i} + \gamma ecm_{t-1} + e_t \quad i=1, \dots, t \quad (16)$$

where y are the dependent variables and x is a vector of independent variables and ecm is the error correction term. Using the general-to-specific modeling approach, we finally obtained the results found in Table IV.

The equation of interest is DLKWhr_{pc}, the change in the natural logarithm or growth of consumption. (The results for the change in urbanization are included as part of the system, but are not discussed because they are not the objective of the paper.) Household disposable income and urbanization have positive and significant effects on residential electricity consumption, even in the short run. The short-run vs. long-run elasticities are 0.25 vs. 1.04 for income, and 1.94 vs. 3.91 for urbanization. Thus, as expected, the long-run elasticities are consistently higher.

The own price of electricity was found to be inelastic, -0.14, but not the relative price or the price of

oil. The two latter variables did not provide any explanatory power to the model and were excluded from subsequent analysis. This may not be as surprising as it may first seem; electricity in Taiwan is the only source for indoor cooling/heating, thus substitution possibilities are small. This and more generally the fact that electricity is considered a basic consumption good, lend themselves as explanations for the low price elasticity.

The change in urbanization from a year ago has a positive effect on current consumption growth: A one percent increase in urbanization leads to an almost two percent increase in consumption - still less than half the long-run effect. The urbanization measure was lagged one period since electricity consumption is not likely to react immediately to changes in its value. The positive effect reflects Taiwan Power Corporation's requirement to serve customers, but also the greater propensity of urban residents to consume electricity than rural customers.

The error-correction term is significant and has a coefficient of -0.06, indicating that when demand is above or below its equilibrium level, consumption adjusts by less than one-tenth within the first year. There is a second "correction" type term in the lagged expenditure share on electricity consumption. If the share increased last year, there is a significant downward effect this year.

The model overall has coefficients with the expected signs and with magnitudes which seem reasonable. The model also performs well statistically: more than a third of the annual variation of the change is explained, and the residual summary statistics for autocorrelation, conditional heteroscedasticity, and normality do not reveal any problems.

The stability of the model is tested by examining model and parameter constancy through recursive estimation. Figures 8 and 9 present the results from the tests. In the top two graphs of Figure 8 we show the recursive residuals. The lower two contain one-step ahead and n-step-ahead Chow tests. The first is equivalent to a sequence of one year ahead forecast tests and the second is a sequence of break point Chow tests. The test statistics are normalized on the critical values at 5%. This implies values greater than unity represent rejections of the null hypothesis of structural breaks. There are

significant forecast errors in 1979 and 1981, but otherwise there does not appear to be any structural breaks (with 25-26 observations we might expect 1-2 rejections). In Figure 9, the recursive coefficient estimates and the associated two standard error bounds are plotted. Except for the change in the relative price of electricity in the mid 1970s, the estimates are stable.

Conclusion

The residential electricity demand model developed here is an attempt at understanding consumption in a (rapidly growing) developing country. We have used a proxy variable, urbanization, to capture economic development characteristics and electricity-using capital stocks not explained by income. The variable provides significant explanatory power to the model both in the short- and long-run. We have not seen this variable used in previous studies and we think it will be useful in explaining residential electricity consumption in other developing countries. Urbanization is an indirect measure of electricity-using capital stock and subject to measurement error. As electricity use penetrates residential markets, the value of this measure may decline. However in the case of Taiwan, the effect of urbanization and the model itself appear stable. The short-run income and price elasticities are smaller than the long-run elasticities, and thus are consistent with economic theory. The income elasticity appears to be unity in the long-run.

Taiwan is pursuing a privatization program for the electric utility sector. Thus, policy makers and private investors can benefit from the cointegration and error correction modeling techniques used in this paper. Note that this model's forecasting power has not been rigorously tested. This is an important topic for further research and a useful application of the elasticity estimates found here.

Figure 1

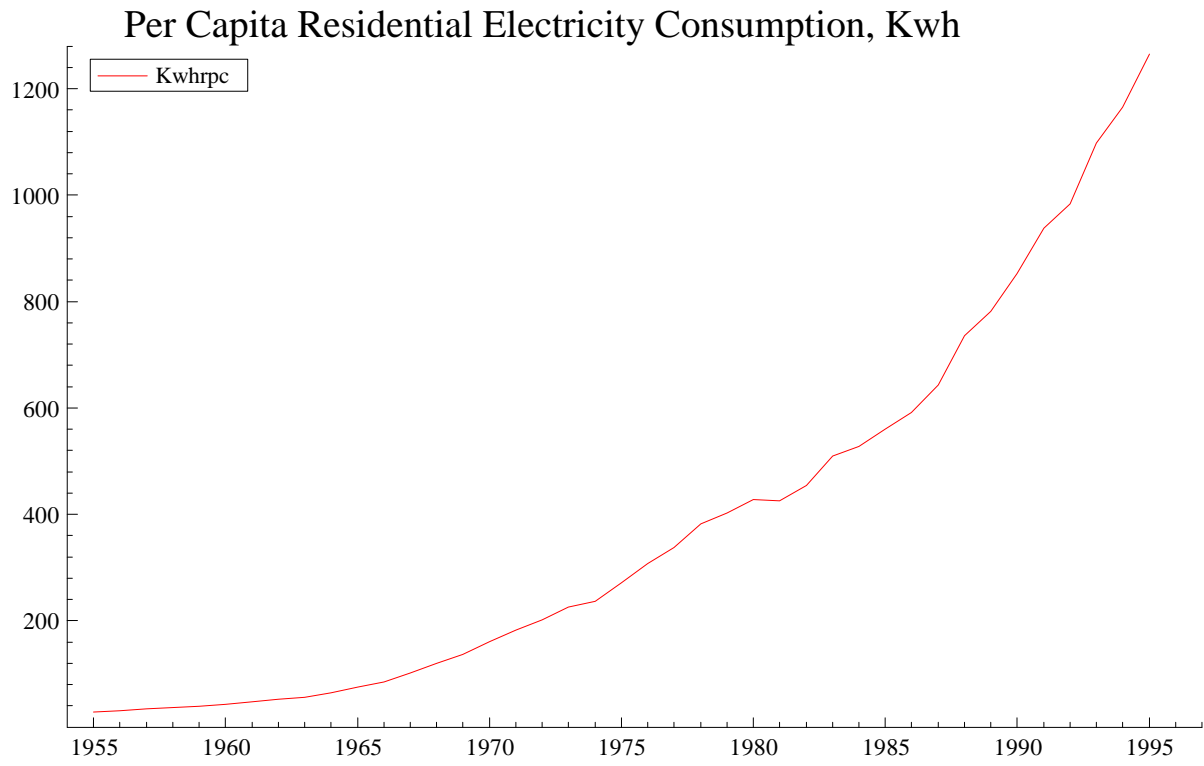


Figure 2

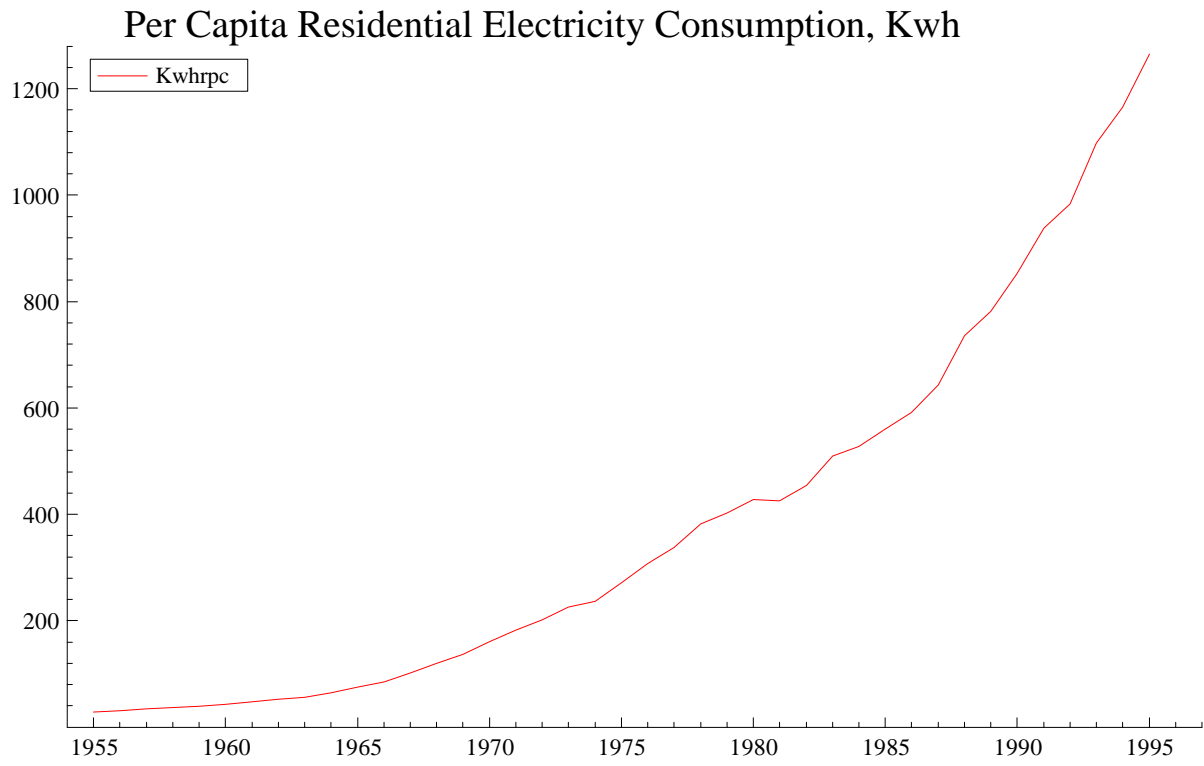


Figure 3

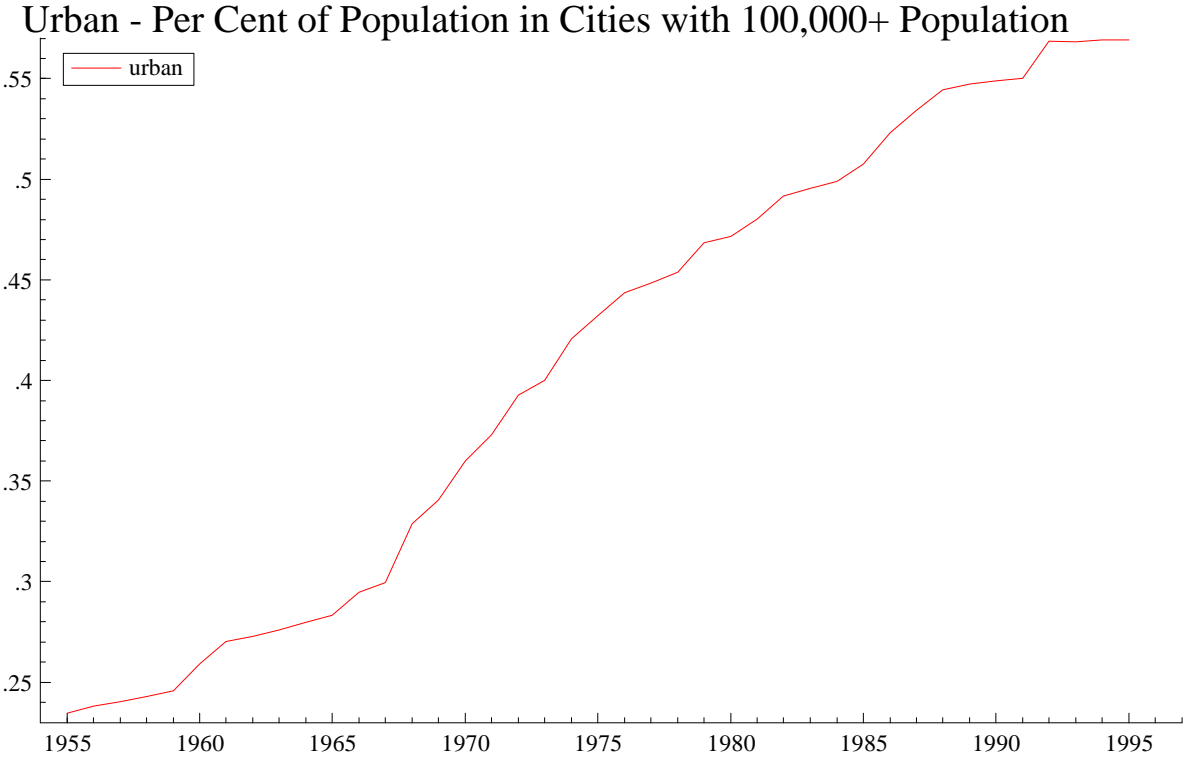


Figure 4

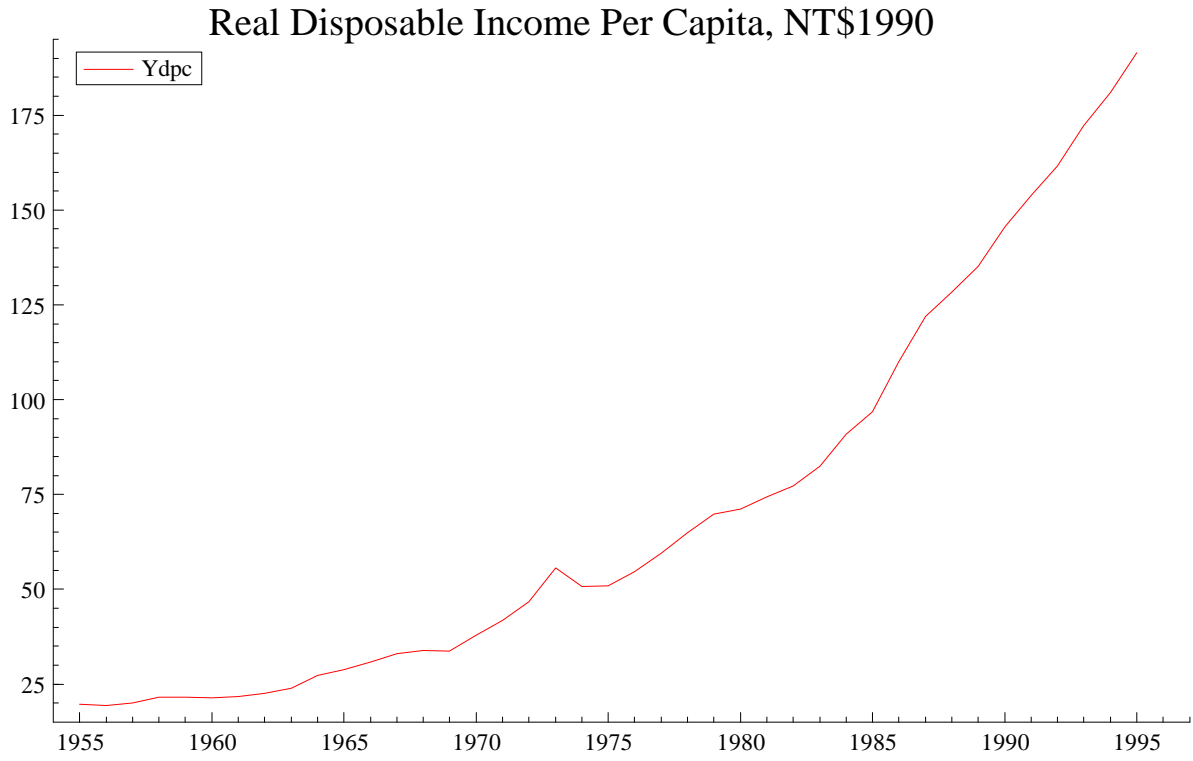


Figure 5

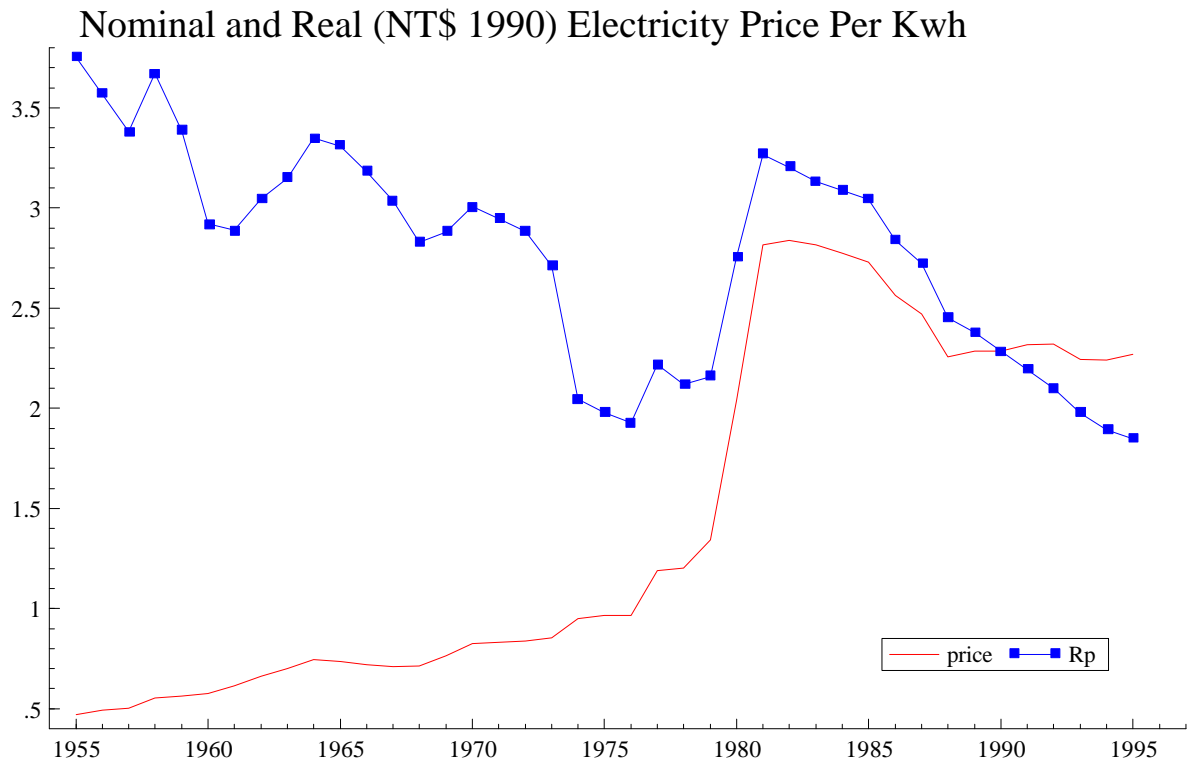


Figure 6

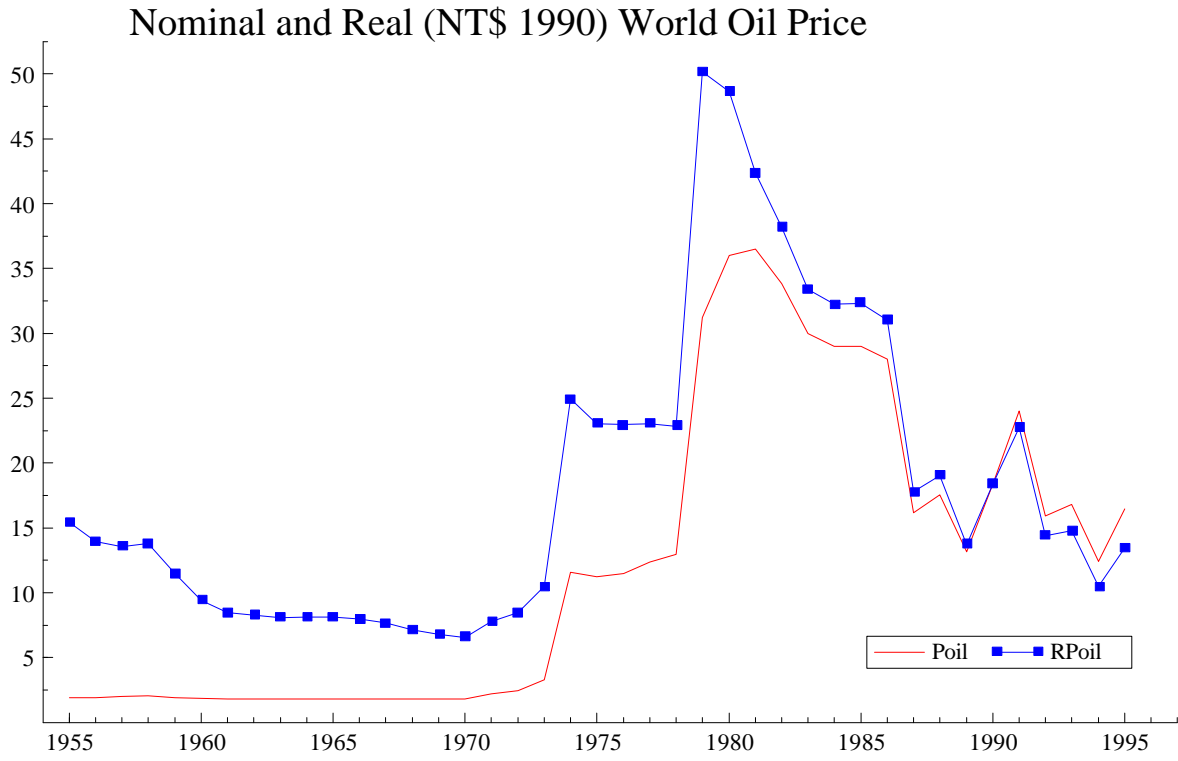


Figure 7

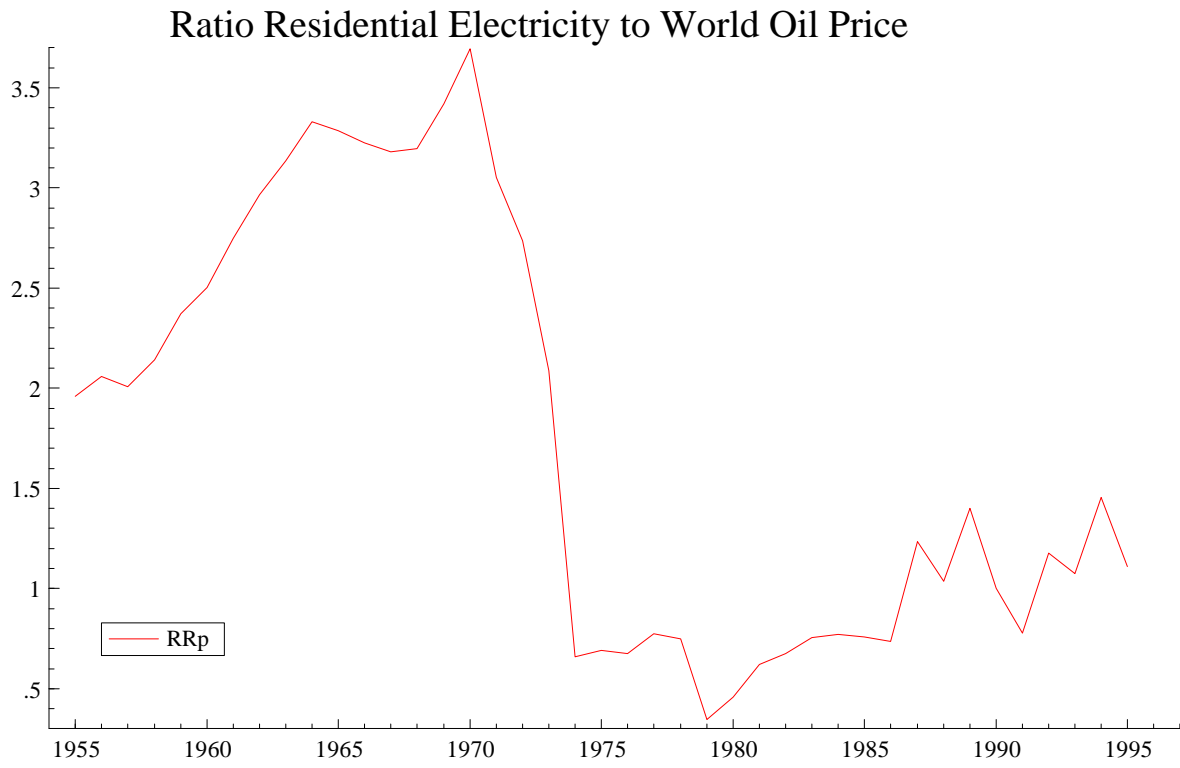
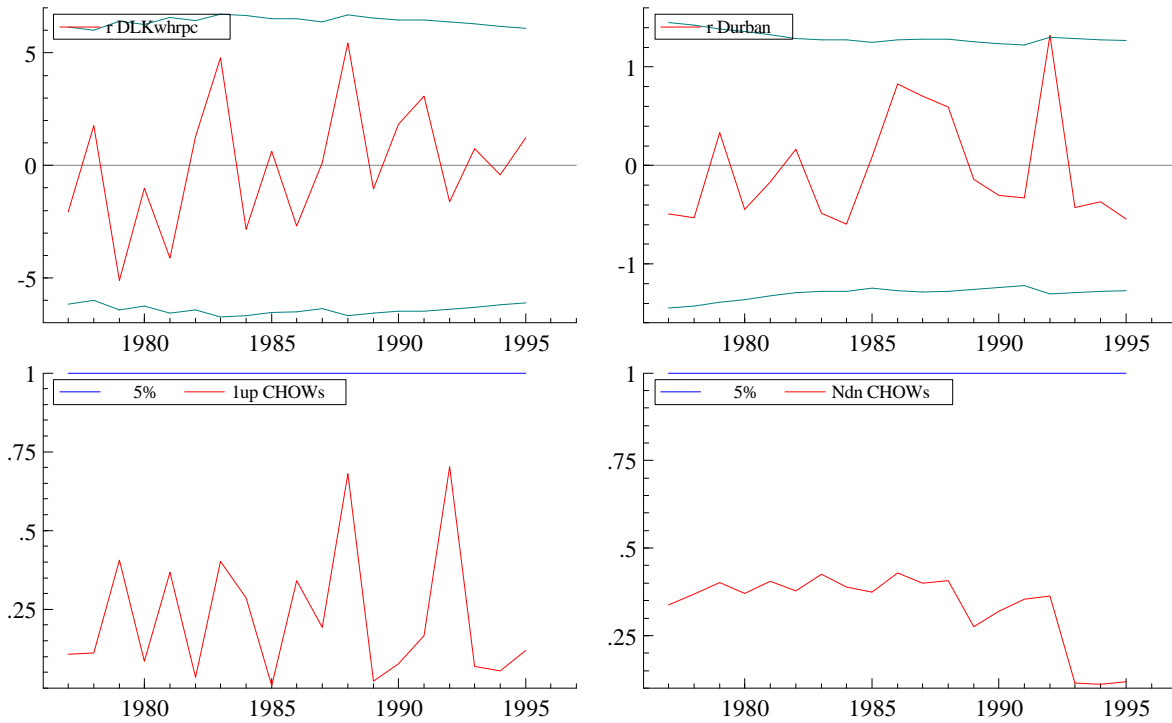
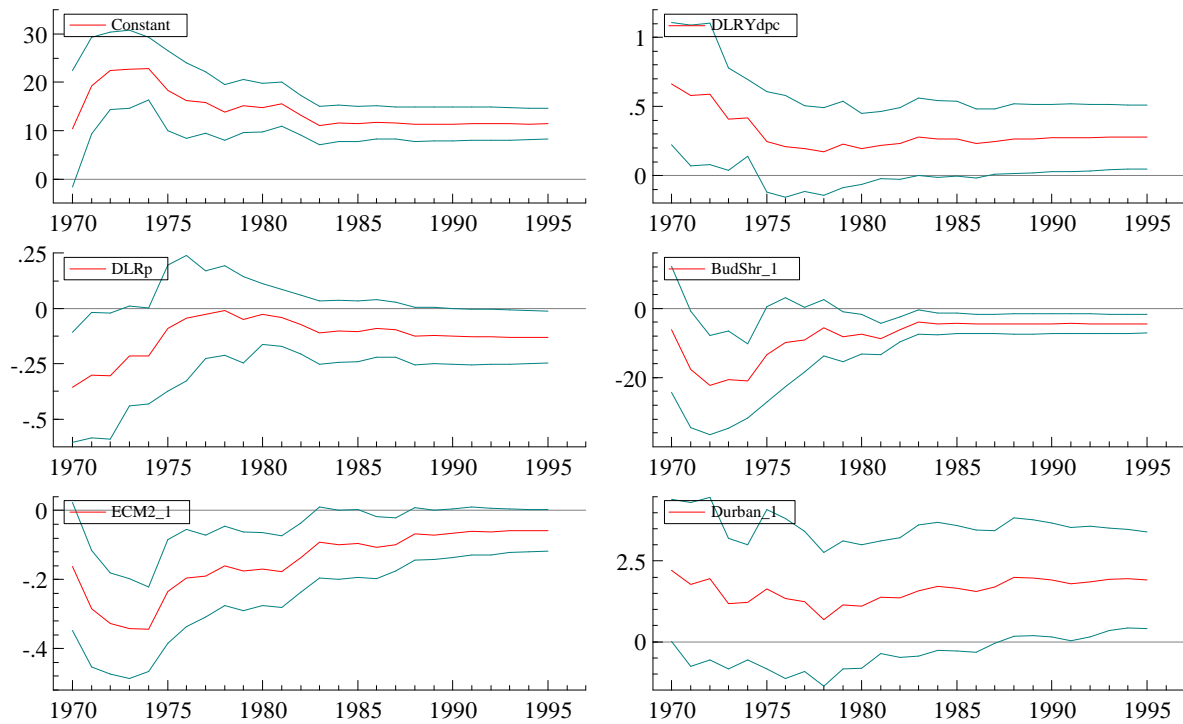


Figure 8
Stability Analysis of Final Error Correction Model (System)
of Residential Electricity Demand



The top two graphs present the recursive 1-step ahead errors with two standard error bounds. The bottom left graph plots out the system 1-up Chow test. The model stability test using the recursive Chow-test is given in the bottom right hand graph.

Figure 9
Stability Analysis of Final Error Correction Model
of Residential Electricity Demand
Recursive Coefficient Estimates with +/- 2 Standard Error Bounds



The recursive coefficient estimates with their plus or minus two standard error bounds are given for the constant, current change in income, current real price of electricity, the budget share for electricity last year, the error or equilibrium consumption correction term, and the change in urbanization last year.

References

Andersson, R. and L. Taylor, "The Social Cost of Unsupplied Electricity: A Critical Review," *Energy Economics*, Vol. 8, No. 3, July 1986, pp. 139-146.

Bentzen, J. and T. Engsted, "Short- and long-run elasticities in energy demand: A cointegration approach," *Energy Economics*, Vol. 15, No. 15, January 1993, pp. 9-16.

Berndt, E.: "The Practice of Econometrics", 1991

Branch, R.E. 1993: "Short Run Income Elasticity of Demand for Residential Electricity Using Consumer Expenditure Survey Data", The Energy Journal, Vol. 14, No.4.

Bohi, D.R. and M. Zimmerman, "An Update on Econometric Studies of Energy Demand," *Annual Review of Energy*, 9, pp. 105-154.

Boswijk, Peter and Philip H. Franses, "Dynamic Specification and Cointegration," *Oxford Bulletin of Economics and Statistics*, 54, 3, 1992, pp.369-81.

Campbell, J.Y. and P. Perron, "Pitfalls and Opportunities: What Macroeconomists Should Know About Unit Roots" in Macroeconomics Annual, Vol. 6, (eds.) O.J. Blanchard and S. Fischer (Cambridge, MA: MIT Press, 1991).

Dahl, Carol, "A Survey of Energy Demand Studies in Support of the Development of NEMS," prepared for the U.S. Department of Energy, October 1993.

Davidson, J.E.H., D.F. Hendry, F. Srba, and S. Yeo, "Econometric Modelling of the Aggregate Time-Series Relationship Between Consumers' Expenditure and Income in the United Kingdom," *Economic Journal*, Vol. 88, No. 4, October 1978, pp. 661-692.

Enders, W, "Applied Econometric Time Series", 1995

Engle, R. and C.W.J. Granger, "Cointegration and Error-Correction: Representation, Estimation and Testing," *Econometrica*, Vol. 55, No. 2, March 1987, pp. 251-276.

and J.J. Hallman, "Merging Short- and Long-Run Forecasts: An Application of Seasonal Cointegration to Monthly Electricity Sales Forecasts," *Journal of Econometrics*, 40, 1989, pp 45-62.

Espey, Molly, "Turning on the Lights: A Meta-Analysis of Residential Electricity Demand Analysis" Working Paper, University of Nevada at Reno, 1999.

Greene, D., "Vehicle Use and Fuel Economy: How Big is the 'Rebound Effect'?", *The Energy Journal*, Vol. 13(1), 1992, pp. 117-143.

Harris, J.L., and Liu, Lon-Mu 1993: "Dynamic Structural Analysis and Forecasting of Residential Electricity Consumption", International Journal of Forecasting, Vol. 9.

Halvorsen, R., "Residential Demand for Electric Energy," *The Review of Economics and Statistics*, Vol. 57, No. 1, February 1975, pp. 12-18.

Hendry, David F., "Econometric Modelling with Cointegrated Variables: An Overview," *Oxford Bulletin of Economics and Statistics*, 48, 3, 1986, pp 201-12.

Johansen, S. and K. Juselius, "Maximum Likelihood Estimation and Inference on Cointegration-With Applications to the Demand for Money," *Oxford Bulletin of Economics and Statistics*, Vol. 52, No. 1, March 1990, pp. 169-210.

Kokklenberg, E. and T. Mount, "Oil Shocks and the Demand for Electricity," *The Energy Journal*, 14(2), 1993, pp. 113-138.

Nelson, C.R. and C.I. Plosser, "Trends and Random Walks in Macroeconomic Time Series: Some Evidence and Implications," *Journal of Monetary Economics*, Vol. 10, No. 1, February 1982, pp. 139-162.

and S.C. Peck, "The NERC Fan: A Retrospective Analysis of the NERC Summary Forecasts," *Journal of Business and Economic Statistics*, Vol. 3, No. 3, pp 179-187, July 1985.

, S.C. Peck, and R.G. Uhler, "The NERC Fan in Retrospect and Lessons for the Future," *Energy Journal*, Vol. 10, No. 2, pp91-107, 1989.

Osterwald-Lenum, M., "A Note with Quantiles of the Asymptotic Distribution of the Maximum Likelihood Cointegration Rank Test Statistics," *Oxford Bulletin of Economics and Statistics*, Vol. 54, No. 3, August 1992, pp. 461-472.

Plourde, Andre and David Ryan, "On the Use of Double-Log Forms in Energy Demand Analysis," *The Energy Journal*, vol. 6, no. 4, 1985, pp 105-13.

Taiwan Power Company Annual Report: 1997, Taipei, Republic of China.

Taylor, L., "The Demand for Electricity: A Survey," *The Bell Journal of Economics*, Vol. 6, Spring 1975, pp. 74-110.

DATA APPENDIX

The electricity consumption and price data for this study came from officials at the Taiwan Power Corporation and the Energy Commission in the Ministry of Economic Affairs. Mr I.L. Wang and Stephen S.T. Lee, successive Directors of the General Planning Department, provided data from the Taipower Annual Report, 1996 and the Taipower Statistical Yearbook:1996. In some cases they translated consumption data tables on from Chinese. Ms. Cindy Hu collected historical and projected capacity and generation data from the Taiwan Power Company Annual Report, 1997. The demographic, disposable income, and consumer price data came from the Taiwan Statistical Data Book, 1996. The world oil price data came from the U.S. Annual Energy Review: 1998 published by the U.S. Energy Information Administration and the OECD International Energy Agency Statistics: Energy Prices and Taxes,Q2 1996.

Table I								
ADF(3) Statistics Testing for a Unit Root								
	Variable							
Order	KWh	KWhpc	Yd	Ydpc	Rp	Rpoil	Rpp	Urban
I(1) - 8	1.01	0.76	9.97	7.64	-1.04	-0.19	-1.44	6.32
I(1) - 9	-2.82	-2.31	-0.52	0.39	-1.78	-1.29	-1.37	-1.74
I(1) -10	-1.01	-1.86	-3.24	-4.19*	-1.49	-1.21	-1.48	-0.07
I(2) - 8	-0.62	-0.54	-2.00	-2.57**	-3.78**	-6.03**	-6.18**	-2.59*
I(2) - 9	-3.19*	-3.78**	-5.24**	-5.23**	-4.05**	-5.94**	-6.12**	-5.12**
I(2)-10	-4.35**	-4.51**	-4.63**	-4.76**	-4.00**	-5.89**	-6.05**	-5.46**

The sample period for the ADF test was from 1959-1995. All variables are in natural logarithms with the exception of the Urban variable. The first three rows, I(1)-i, present the ADF t-tests corresponding to tests for unit roots in the levels of the series. The last three rows, I(2)-i, report the ADF t-test results for testing whether the first difference has a unit root. A rejection implies that the first difference of the series is a stationary process.

The identifier i refers to equations 8, 9, and 10 in the paper which are ADF regressions with no constant, a constant, and a constant plus trend respectively. The critical values for the t-tests at 5% are -1.95, -2.95, and -3.54 respectively; at 1% they are -2.62, -3.62, and -4.23 respectively. Rejections at the 5% and 1% critical values are denoted as * and ** respectively. The critical values for this table are calculated from MacKinnon (1991).

A maximum of three lags was used in each test. The appropriate lag length was chosen by examining the autocorrelation function of the residuals.

Table II
Lag Length Specification Tests
***F* and Related Statistics for the Sequential Reduction**
from the Third-Order VAR to the First-Order
Sample Period 1958-1995

Null Hypothesis				Maintained Hypothesis		
System	k	<i>L</i>	SC	VAR(3,3)	VAR(3,0)	VAR(2,0)
VAR(3,3)	76	612.65	-24.97	n.a.		
VAR(3,0)	64	603.64	-25.64	0.69668 [0.7453] (12, 42)		
VAR(2,0)	48	594.23	-26.68	0.65113 [0.8923] (28, 59)	0.64572 [0.8327] (16, 58)	
VAR(1,0)	32	580.58	-27.49	0.79605 [0.7865] (44, 63)	0.87297 [0.6579] (32, 71)	1.1749 [0.3094] (16, 70)

The first four columns report the VAR models with the lags of the endogenous variables and the lags of the exogenous variable, the real oil price, the number of unrestricted parameters, *k*, the log likelihood statistic, *L*, and the Schwarz-Criterion, SC.

In the individual cells of the three columns containing the maintained hypothesis results are: the approximate F-statistic for testing the null hypothesis in the first column against the maintained model, the tail probability associated with the F-statistic in square brackets, and the respective degrees of freedom for the constrained parameters and the observations minus free parameters in parentheses.

Table III
Cointegration Analysis of the Taiwan Residential Electricity Data

Eigen value	0.598	0.433	0.266	0.102	
Null Hypothesis	$r = 0$	$r \leq 1$	$r \leq 2$	$r \leq 3$	
$\lambda \text{ max}$	36.48**	22.69	12.38	4.33*	
$\lambda \text{ max}$ adjusted	32.83**	20.42	11.15	3.9*	
95% critical value	30.3	23.8	16.9	3.7	
$\lambda \text{ trace}$	75.89**	39.41*	16.72	4.33*	
$\lambda \text{ trace}$ adjusted	68.3**	35.47*	15.05	3.9*	
95% critical value	54.6	34.6	18.2	3.7	
Standardized Eigenvectors β'					
Variable	LKWWhpc	Urban	LRYdpc	LRp	LRpoil
	1.00	-3.31	-1.57	0.15	-0.18
	-0.23	1.00	-0.29	-0.22 e-03	0.1 e-05
	0.51	-8.46	1.00	-0.42	0.11
	-3.49	11.81	14.38	1.00	-6.70
Standardized Speed of Adjustment Coefficients α'					
LKWWhpc	0.081	0.113	0.007	0.003	
Urban	0.015	-0.068	0.018	-0.0001	
LRYdpc	0.100	0.041	-0.178	-0.7 e-04	
LRp	0.084	1.055	-0.085	-0.003	
Weak Exogeneity Tests					
Variable	LKWWhpc	Urban	LRYdpc	LRp	
$\chi^2(\cdot)$	8.98**	5.07*	4.85*	1.27	
<i>p-value</i>	0.003	0.024	0.028	0.25	

The VAR is estimated over the period 1957-1995. It includes one lag of each variable and the contemporaneous value of the real oil price. A constant, trend and the dummy variable for 1979-1981 are entered in unrestricted form.

The statistics $\lambda \text{ max}$ and $\lambda \text{ trace}$ are Johansen's maximal eigenvalue and trace statistics for testing for cointegration. The null hypothesis is in terms of the cointegration rank r . For example rejection of $r=0$ is evidence in favor of at least one cointegrating vector. The adjusted statistics control for the degrees of freedom. The critical values are taken from Osterwald-Lenum (1992).

The weak exogeneity test statistics are evaluated under the assumption that $r=1$ and so are asymptotically distributed as chi-squares with 1 degree of freedom if weak exogeneity of that variable for the cointegrating vector is valid.

Table IV
Final Error Correction Representation
of Residential Electricity Demand
Sample is: 1957 to 1995

Equation 1 for DLKwhrpc

Variable	Coefficient	Std.Error	t-value	t-prob	HCSE
DLRp	-0.14755	0.056960	-2.590	0.0139	0.063253
DLRYdpc	0.25711	0.11072	2.322	0.0262	0.14792
ECM2_1	-0.059986	0.029290	-2.048	0.0481	0.032711
BudShr_1	-4.3307	1.3259	-3.266	0.0024	1.3145
Durban_1	1.9075	0.72127	2.645	0.0122	0.66076
Constant	11.429	1.5706	7.277	0.0000	---

std.err. = 2.87004

Equation 2 for Durban

Variable	Coefficient	Std.Error	t-value	t-prob	HCSE
ECM2_1	-0.021893	0.0056386	-3.883	0.0004	0.0059263
BudShr_1	0.37006	0.25641	1.443	0.1578	0.21285
Constant	0.50399	0.30942	1.629	0.1123	---

std. err. = 0.597863

loglik = -16.392804 $\log|\Omega| = 0.840657$ $|\Omega| = 2.31789$ T = 39
 LR test of over-identifying restrictions: $\chi^2(7) = 13.0686$ [0.0705]

correlation of residuals

	DLKwhrpc	Durban
DLKwhrpc	1.0000	
Durban	-0.15008	1.0000

DLKwhrpc:Portmanteau 2 lags= 1.3823
 Durban :Portmanteau 2 lags= 0.092677
 DLKwhrpc:AR 1- 2 F(2, 29) = 1.1116 [0.3426]
 Durban :AR 1- 2 F(2, 29) = 5.4573 [0.0097] **
 DLKwhrpc:Normality $\chi^2(2)$ = 0.76634 [0.6817]
 Durban :Normality $\chi^2(2)$ = 10.088 [0.0064] **
 DLKwhrpc:ARCH 1 F(1, 29) = 0.018695 [0.8922]
 Durban :ARCH 1 F(1, 29) = 0.073736 [0.7879]
 DLKwhrpc: χ^2 F(14, 16) = 1.1612 [0.3836]
 Durban : χ^2 F(14, 16) = 0.44776 [0.9310]
 Vector portmanteau 2 lags= 2.0882
 Vector AR 1-2 F(8, 60) = 0.272 [0.9727]
 Vector normality $\chi^2(4)$ = 10.867 [0.0281] *
 Vector χ^2 F(42, 54) = 0.80393 [0.7673]

¹ The issues are discussed in the introductory chapter in Doornik and Hendry (1992).