

Long-Run Price and Income Elasticities of Namibian Aggregate Electricity Demand: Results from the Bounds Testing Approach

¹Emmanuel Ziramba and ²Kasnath Kavezeri

²Department of Economics, University of Namibia,
Private bag 13301 Windhoek, Namibia.

¹Department of Economics, Polytechnic of Namibia, Namibia.

Corresponding Authors: Emmanuel Ziramba

Abstract

This paper estimates the short-run and long-run price and income elasticities of aggregate electricity demand in Namibia over the period 1993q1 to 2010q1. We make use of the bounds testing approach to cointegration within an autoregressive distributed framework. Our empirical results suggest that electricity demand is income elastic, but price inelastic in the long-run.

Keywords: bounds testing, cointegration, long-run elasticities

INTRODUCTION

Namibia is a middle income country and one of the most industrializing countries in Africa. The Namibian economy is heavily dependent on energy, especially in the mining sector. The economy is driven mainly by diamond and uranium exports. Electricity is used for a number of purposes which include industrial, commercial and household purposes. It is supplied to consumers by a public utility known as NamPower.

The main objective of this paper is to estimate long-run price and income elasticities for aggregate electricity demand in independent Namibia. We also use most recent data, covering the period 1993q1 to 2010q1. Our study makes a methodological contribution to the literature on electricity demand in Namibia. We use the bounds testing approach to cointegration, developed by Pesaran et al. (2001), within an autoregressive distributed lag (ARDL) framework, to test for a long-run level relationship in the demand for aggregate electricity.

The remainder of the paper is organized as follows: the next section gives an overview of the energy sector in Namibia. Section 3 reviews a selection of some of the empirical studies on the demand for electricity. Section 4 outlines the empirical model specification used in this paper. The econometric techniques which are employed in this study are discussed in the same section. Section 5 presents the empirical results of the study. The final section summarizes the main findings of the paper and gives some policy implications.

An Overview of Namibia's Energy Sector

The Electricity Supply Industry (ESI) in Namibia falls under the Ministry of Mines and Energy

(MME), which is responsible for the overall national energy policy to ensure sustainable development and utilization of the country's energy resources. This industry has been governed by South African electricity acts from the colonial period until the year 2000 when the new Electricity Act of 2000 for Namibia came into existence. This Act made provision for the creation of an Electricity Control Board (ECB) to regulate the electricity industry and ensure customer protection (Ministry of Mines and Energy, 1998). The electricity Act of 2000 was later repealed in its entirety and replaced by Electricity Act 4 of 2007 (Electricity Control Board, 2010).

The ECB is, thus, a statutory regulatory body mandated to restructure the ESI in collaboration with the MME and other stakeholders in the industry. The responsibilities of the ECB include, inter alia, setting up the regulatory authority internally, implementing tariff methodologies and the issuing of licenses for different types of activities in the industry. For instance, all generation operations with a capacity of 500kW and more, need to be licensed (MME, 2008).

The other significant role players in the ESI that also came into existence as a result of government restructuring policy are the Regional Electricity Distributors (REDs). The purpose of the ESI restructuring was mainly to create a more efficient and competitive electricity sector. At that time, the distribution of electricity to end-users was the responsibility of about 46 different municipalities. This often led to inefficiency as many of the smaller distributors faced a lot of challenges, such as the lack of sufficiently skilled personnel, limited financial capacity, etc. As a result, the REDs were established to ensure efficient service delivery at cost-reflective prices. Three of the five proposed REDs are already

in full operation. The functional ones are NORED Electricity, which covers the North and North-eastern regions; CENORED Electricity, covering the central Northern region, while ERONGO RED distributes electricity in the Western part of Namibia (MME, 2008 and Bannon, 2006).

The Namibia Power Corporation, NamPower, is a state-owned power utility that generates and supplies power in bulk to the mines and industries, the Regional Electricity Distributors (REDs) as well as the municipalities and local authorities, who in turn, distribute electricity to end-users. The power utility has a total generation capacity of 384 MW from its three internal power sources, with an additional diesel powered source of 21.5MW capacity still under construction at Anixas. The current sources comprise the 120MW Van Eck coal-fired thermal power station in Windhoek, the 24MW Paratus diesel-powered station at Walvis Bay, and the hydroelectric power station at Ruacana with a 240MW installed capacity (Bannon, 2006).

In addition to the local installed capacity, Namibia also imports electricity from other members of the Southern African Development Community (SADC). These include ESKOM, the South African power utility; ZESA the power utility of Zimbabwe; ZESCO of Zambia; and the EDM of Mozambique. During the 2008/2009 financial year, sixty percent (60%) of Namibia's total electricity demand was met through imports. ESKOM and ZESA each contributed 48% of total imports, while equal portions of the remaining 4% came from ZESCO and EDM (NamPower, 2009).

An Overview of Electricity Demand

Electricity demand has received considerable attention in the empirical literature. Most of the empirical literature on electricity demand is on household or residential demand (see among others, Donatos and Mergos, 1991; Beenstock et al., 1999; Filippini and Pauchauri, 2004; Høltedahl and Joutz, 2004; Narayan and Smyth, 2005; Narayan et al., 2007; Yoo et al., 2007; Ziramba, 2008). For Namibia, there are three earlier studies (Lundmark, 2001 (as cited in De Vita et al. 2006), De Vita et al., 2006 and Kavezeri, 2009) which have looked at the demand for electricity among other energy sources. There are a number of empirical studies that have estimated the elasticities of electricity demand in other countries (among others, Bose and Shukla, 1999; Beenstock et al., 1999; Nasr et al., 2000; Al-Faris, 2002; Khan and Qayyum, 2009). In this study we review some of the studies on the subject starting with the ones on Namibia.

De Vita et al. (2006), using the autoregressive distributed lag (ARDL) bounds testing approach to cointegration, estimate long-run energy demand

elasticities of Namibia at both aggregated level and by energy type (electricity, petrol and diesel) for the period 1980q1-2002q4. Their results for national electricity consumption indicate that both price and income variables were statistically significant and had expected signs. Their respective income and price elasticities are 0.589 and -0.298.

Kavezeri (2009) analyzes the demand for electricity in Namibia over the period 1993:Q1 to 2006:Q4. The author expresses electricity demand as a function of income and the price of electricity. Her study employs the Engle and Granger (1987) approach to cointegration and an error correction model (ECM). The author's results indicate the presence of a long-run equilibrium relationship among the variables. Income was found to be a significant determinant of electricity demand in both short- and long-run periods. Electricity demand was found to be income elastic with a long-run elasticity of 1.02. The price coefficient was found to be negative but was not statistically significant.

Bose and Shukla (1999) estimate price and income elasticities of electricity consumption for distinct consumer groups (residential, commercial, agriculture, small and medium industries, and large industries). Their estimates at the national level were obtained by pooling data across 19 states over 9 years (1985/6-1993/4). Their results indicate that electricity consumption in commercial and large industrial sectors are income elastic, while residential, agricultural and small and medium industries are income inelastic. Their short-run price elasticities vary from -1.35 in agriculture to -0.26 in commercial and insignificant in small and medium industries.

Khan and Qayyum (2009) estimate short-run and long-run price and income elasticities for the national level and for households, industry and agriculture sectors in Pakistan. Their results suggest that the price and income elasticities possess expected signs at both the aggregate and disaggregate levels in both periods. Their long-run results for aggregate demand were as follows. Aggregate demand for electricity is significantly determined by real income, real price of electricity and average temperature. Their income elasticity was as high as 4.7 while their price elasticity estimate was as high as -1.64.

METHODOLOGY

A number of determinants for electricity or energy demand have been considered in the empirical literature. In the simplest form, the demand for electricity or energy in general, has been modelled as a function of a single variable, such as real income (Dincer and Dost, 1997). Some studies have expressed the demand for electricity as a function of own price, price of a substitute and real income (Al-Faris, (2002), Narayan and Smyth, (2005)). Nasr et

al. (2000) model electricity demand in Lebanon as a function of imports (because there were no figures for GDP) and temperature, but do not include any of the price variables.

The demand for electricity has been modelled in a variety of ways in the literature. The most common variables that have been included in the estimation of aggregate electricity demand models include income, price of electricity, price of substitute energy source and temperature variables. We specify the electricity demand as a function of real income and the real tariff (price) of electricity. We do not include the price of a substitute source of energy and temperature variables because data on these variables was not available for the study period. The most common specification of the aggregate electricity demand takes the form of a linear double-logarithmic form (Ang et al., (1992), Donatos and Mergos (1991), Filippini (1999), Beenstock et al., (1999), Hondroyannis (2004)). Our model takes the following form:

$$\ln EC_t = \beta_0 + \beta_1 \ln Y_t + \beta_2 \ln P_t + \varepsilon_t \quad (1)$$

Where $\ln EC_t$ is the natural logarithm of electricity consumption in kilowatt hours (kWh), $\ln Y_t$ is the natural logarithm of real gross domestic product (GDP), $\ln P_t$ is the natural log of the real electricity tariff (N\$/kWh), and ε is a random error term. According to economic theory β_1 and β_2 are expected to be positive and negative respectively.

In the literature a number of studies have applied the bounds test approach to co-integration in analyzing electricity demand (among others Narayan and Smyth, 2005; De Vita et al., 2006; Ziramba, 2008; Khan and Qayyum, 2009). In the present study we specifically follow these studies in applying the bounds testing approach to co-integration analysis in estimating price and income elasticities of electricity demand for independent Namibia.

The bounds testing approach to cointegration can identify the long-run relationship with a dependent variable followed by its *forcing variables* (Sari et al., 2008). It does not require precise knowledge of the order of integration or cointegration ranks of the variables. Without prior information about the direction of the long-run relationship between electricity consumption, real gross domestic product and real electricity price, we constructed the following unrestricted error correction model (UECM):

$$\Delta \ln EC_t = \alpha + \lambda_1 \ln EC_{t-1} + \lambda_2 \ln Y_{t-1} + \lambda_3 \ln P_{t-1} + \sum_{i=1}^k \delta_i \Delta \ln EC_{t-i} + \sum_{i=0}^l \phi_i \Delta \ln Y_{t-i} + \sum_{i=0}^m \pi_i \Delta \ln P_{t-i} + \mu_t \quad (2)$$

Where Δ denotes first difference, μ_t is the error term which is assumed to be white noise and the other

variables are as defined above. The bounds test methodology suggests analysing the null hypothesis of no co-integration through a joint significance test of the lagged variables $\ln Y_{t-1}$, $\ln P_{t-1}$, and $\ln EC_{t-1}$ based on the Wald or F -statistic:

$$H_0: \lambda_1 = \lambda_2 = \lambda_3 = 0$$

$$H_1: \lambda_1 \neq 0, \text{ or } \lambda_2 \neq 0, \text{ or } \lambda_3 \neq 0.$$

We test the null hypothesis of no co-integration by means of the *Wald* test. By adopting Pesaran et al.'s (2001) approach for co-integration analysis, a pre-test for unit root (degree of integration) of the interested series is not necessary. The co integration equation is defined as

$$\hat{\lambda}_1 \ln EC_{t-1} + \hat{\lambda}_2 \ln Y_{t-1} + \hat{\lambda}_3 \ln P_{t-1} = 0 \quad (3)$$

From the estimated UECM, the long-run coefficients are the coefficients of the one-lagged explanatory variables (multiplied by a negative sign) divided by the coefficient of the lagged dependent variable (Bardsen, 1989). The respective long-run income and price coefficients are $\Phi_1 = -(\lambda_2/\lambda_1)$ and $\Phi_2 = -(\lambda_3/\lambda_1)$. These coefficients can be interpreted directly as elasticities. They are the focus of our empirical analysis. The variances of the respective elasticities are given by,

$$\hat{V}(\hat{\Phi}_1) = \left(\frac{\partial \hat{\Phi}_1}{\partial \hat{\lambda}_2} \right)^2 \hat{V}(\hat{\lambda}_2) + \left(\frac{\partial \hat{\Phi}_1}{\partial \hat{\lambda}_1} \right)^2 \hat{V}(\hat{\lambda}_1) + 2 \left(\frac{\partial \hat{\Phi}_1}{\partial \hat{\lambda}_2} \right) \left(\frac{\partial \hat{\Phi}_1}{\partial \hat{\lambda}_1} \right) \hat{COV}(\hat{\lambda}_2, \hat{\lambda}_1) \quad (4)$$

and

$$\hat{V}(\hat{\Phi}_2) = \left(\frac{\partial \hat{\Phi}_2}{\partial \hat{\lambda}_3} \right)^2 \hat{V}(\hat{\lambda}_3) + \left(\frac{\partial \hat{\Phi}_2}{\partial \hat{\lambda}_1} \right)^2 \hat{V}(\hat{\lambda}_1) + 2 \left(\frac{\partial \hat{\Phi}_2}{\partial \hat{\lambda}_3} \right) \left(\frac{\partial \hat{\Phi}_2}{\partial \hat{\lambda}_1} \right) \hat{COV}(\hat{\lambda}_3, \hat{\lambda}_1) \quad (5)$$

A number of advantages of the bounds testing approach to cointegration have been highlighted in the empirical literature. First, it has been argued that the UECM is likely to have better statistical properties because it does not push the short-run dynamics into the residual term as in the Engle-Granger (1987) technique (Pattichis, (1999). Second, the procedure is applicable irrespective of whether the underlying explanatory variables are integrated of order zero ($I(0)$) or one ($I(1)$) (Mah, 2000), provided none of the variables is integrated of order two, $I(2)$. In our case this is confirmed by the unit root tests in Table 2. However, it must be noted that this method is inappropriate if there are two or more cointegrating relationship involving the dependent variable.

EMPIRICAL ANALYSIS AND RESULTS

Data, sources and univariate properties

Quarterly data for the period 1993q1 to 2010q1 for electricity consumption measured in kilowatt hours (KWh) and average electricity prices were obtained from the national electricity supplier, NamPower. Real income is proxied by real gross domestic

product (GDP). Real GDP and consumer price indices were obtained from National Planning Commission (NPC). Nominal electricity tariffs (prices) were deflated using the consumer price index. All variables are expressed in natural logarithms. Some descriptive statistics for the variables are presented in Table 1.

Table 1. Summary of basic statistics

	lnY	lnP	lnEC
Mean	9.130681	3.268602	6.261389
Maximum	9.504953	3.602225	6.602317
Minimum	8.762427	2.991140	5.886895
Std. Deviation	0.218150	0.164077	0.205190

Source: Own calculations based on data from NamPower and the National Planning Commission.

Figure 1 plots the individual time series of the data employed in this study in their natural logarithms. A stationary series is characterized by a time-invariant mean and a time-invariant variance. There are alternative methods that are used to test for non-stationarity of a time series. In order to confirm that none of the variables is integrated of order two we used the augmented Dickey-Fuller test to examine the unit root properties of the data. The results reveal that the null hypothesis of a unit root is not rejected in levels for all variables but was rejected in first differences. The unit root tests suggest that all the variables are integrated of order one and these results are reported in Table 2.

Table 2 ADF unit root tests

Variable	Levels			First differences		
	none	intercept	Trend & intercept	none	intercept	Trend & intercept
lnEC	3.607 (-1.614)	-0.360 (-2.591)	-2.863 (-3.167)	-9.303 ^a (-2.600)	-10.773 ^a (-3.533)	-10.686 ^a (-4.103)
lnY	4.094 (-1.614)	-0.442 (-2.590)	-2.681 (-3.167)	-10.643 ^a (-2.600)	-12.547 ^a (-3.532)	-12.450 ^a (-4.101)
lnP	1.458 (-1.613)	-0.788 (-2.591)	-2.387 (-3.169)	-2.719 ^a (-2.602)	-3.115 ^b (-2.908)	-3.103 (-3.169)

Where a, b denote significance at the 1% and 5% levels respectively. The numbers in the parentheses for the insignificant tests are the critical values at 10% level.

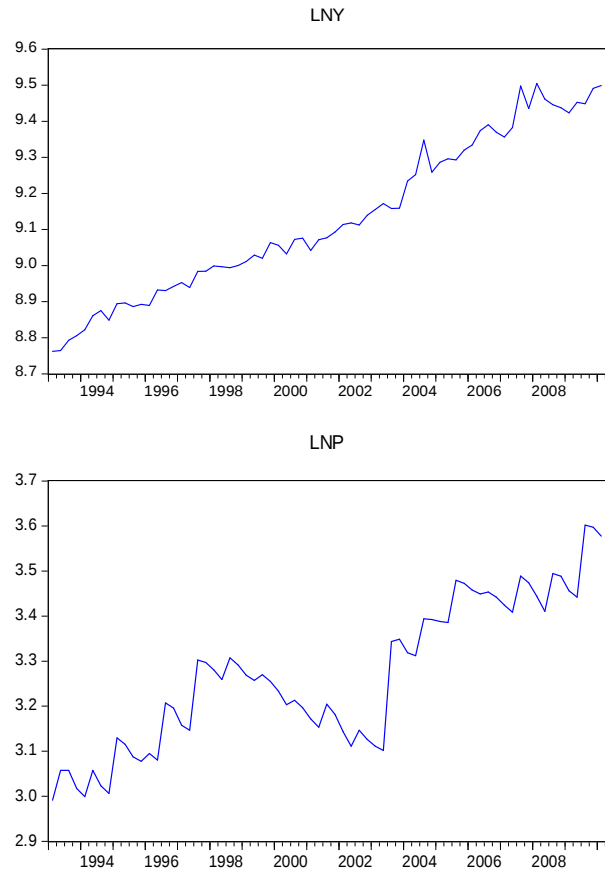
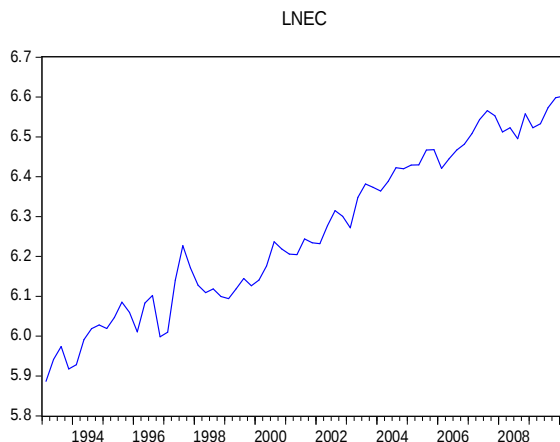


Figure 1. Individual time series plots.

Cointegration Test

In order to perform bounds test we estimate a parsimonious UECM. Given our sample size, we start by introducing a lag length of five for the differenced variables. Following Hendry et al. (1984), we obtained the preferred parsimonious UECM is presented in Table 3. The adequacy of this UECM is checked through a set of specification and diagnostic tests. The results of such tests are reported in Table 3. As is reflected in Table 3, the parsimonious UECM passes all specification and diagnostic tests. The Ramsey RESET (2) test does not reject the null hypothesis of no misspecification in the functional form. The Jarque Bera test confirms the normality of the residuals. The Breusch-Godfrey LM test does not reject the null hypothesis of no serial correlation in the residuals. The ARCH (2) test confirms that there is no evidence of autoregressive conditional heteroscedasticity. White's test does not reject the null hypothesis of no heteroscedasticity.

Based on the estimated parsimonious UECM, we performed the bounds test using the standard Wald test to assess for the significance of λ_1 , λ_2 and λ_3 in equation (2). The results of the bounds test are presented in Table 4. Thus, the null hypothesis of no co integration, that is, $\lambda_1 = \lambda_2 = \lambda_3 = 0$, is rejected and

we conclude that there is a stable long-run relationship among aggregate electricity demand, income and price when electricity consumption is the dependent variable.

Table 3: UECM for electricity demand (Dependent variable: $\Delta \ln EC_t$)

Regressor	Coefficient	t-statistic
$LNP_{(t-1)}$	-0.106 ^b	-2.102
$LNY_{(t-1)}$	0.372 ^a	2.838
$LNEC_{(t-1)}$	-0.332 ^b	-2.523
ΔLNY_t	0.235	1.781
$\Delta LNY_{(t-5)}$	0.330	2.914
$\Delta LNEC_{(t-1)}$	0.292	2.387
$\Delta LNEC_{(t-2)}$	-0.367	-3.027
$\Delta LNEC_{(t-5)}$	-0.150	-1.454
intercept	-0.965	-2.454
R^2	0.502	
DW	2.0607	

Diagnostic tests

Ramsey	F-statistic: 1.176 (0.317)
RESET (2)	
Jarque-Bera test	JB-statistic: 1.103 (0.576)
Breusch-Godfrey LM test	F-statistic: 0.293 (0.747)
White's test	Test statistic: 1.620 (0.134)
ARCH test (2)	F-statistic: 0.583 (0.448)

Where a, and b indicate 1%, and 5% significance level, respectively. The numbers in the parentheses are the p-values.

Table 4: Bounds tests co-integration test results

Dependent variables	Wald statistic
$\Delta \ln EC_t$	3.157 ^b (0.032)

Where b represents statistical significance at the 5% level.

Based on the variance-covariance matrix of coefficients in the UECM drawn from Eviews 7, the variances of each long-run elasticities are computed using the method suggested by Bardsen (1989) as outlined in equations (4) and (5). The t-statistics of income and price elasticities are, respectively 8.366 and -1.431.

Income and Price Elasticity of Elasticity Demand

Table 5: Long-run elasticities of electricity demand

Income elasticity	Price elasticity
1.121(8.366) ^a	-0.320

Numbers in the parentheses indicate t-statistics. The elasticities and variances (hence t-statistics) are computed using the methodology developed by Bardsen (1989), and a indicates significance at the 1% level.

The computed long-run income elasticity is in line with the theoretical expectations and is statistically significant. The income variable is statistically significant at 1 per cent level of significance. The income elasticity value is 1.121. The results indicate that electricity consumption is a normal good as it increases with income. The magnitude of the long-run income elasticity estimate is within the range of previous studies in other countries. Kavezeri (2009) found a long-run income elasticity estimate of 1.02. An increase in income will bring about a gradual increase in the demand for aggregate electricity in the long-run as households become wealthier and industries produce more goods and services. The price elasticity is negative but statistically insignificant. The long-run price elasticity is -0.320. This result is not surprising as electricity does not have close substitutes in the form of alternative energy sources in Namibia.

Constancy of the Cointegration Space

After estimating the parsimonious model, the cumulative sum of recursive residuals (CUSUM) and the CUSUM of squares (CUSUMSQ) tests were applied to test for parameter constancy. Figure 1 plots the CUSUM and CUSUM of squares statistics for equation (2). The results clearly indicate the absence of any instability of the coefficients during the investigated period because the plots of the two statistics are confined within the 5% critical bounds of parameter stability.

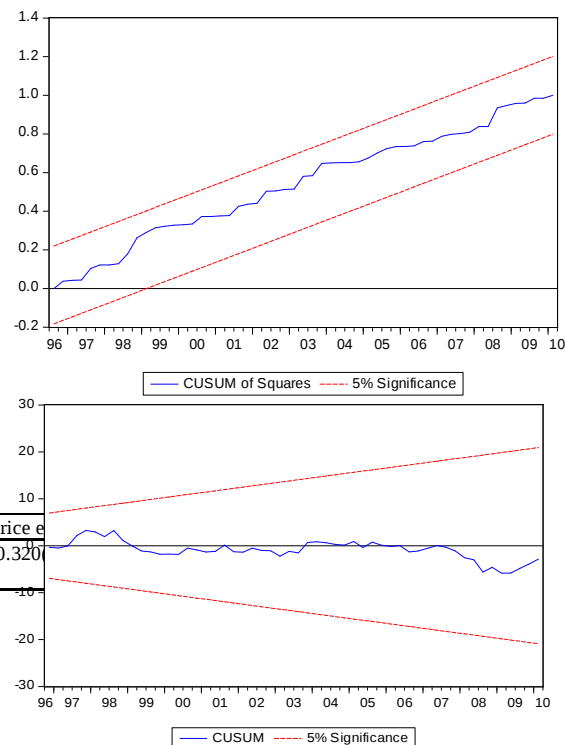


Figure 1. Plots of the CUSUM and CUSUM of Squares

SUMMARY AND CONCLUDING REMARKS

The main purpose of this study was to estimate long-run price and income elasticities for aggregate electricity demand in Namibia for the period 1993q1 to 2010q1. This was achieved by employing the bounds testing approach to co-integration analysis. This information will be useful to the designing of electricity pricing policy in Namibia.

The findings in this study for the long-run income elasticity of demand are consistent with previous studies for other countries. The income elasticity of demand has a positive sign as is expected and is statistically significant. The computed long-run income elasticity is 1.121. Our results indicate that a 10 percent increase real GDP will increase electricity consumption by about 11.2 percent. The computed long-run price elasticity has the expected negative sign but is statistically insignificant. Its magnitude is -0.32, indicating that aggregate elasticity demand is price inelastic. The stability tests performed demonstrate that the long-run aggregate demand for electricity in Namibia remained unchanged throughout the estimation period.

Electricity demand studies have important practical implications. The results indicate that the estimated aggregate demand for electricity can be used for policy purposes since it is stable. The fact that a stable aggregate electricity demand function seems to exist, would make forecasting of electricity demand at the national level possible. The estimated price and income elasticities of -0.32 and 1.121 imply that electricity demand in Namibia is price inelastic and but income elastic. It shows that tariff increases alone will not reduce electricity consumption and that the increase in income induces a significant increase in electricity demand. These up-to-date income and price elasticity estimates of electricity demand will prove useful to the design of electricity pricing policy. This study is also expected useful information is expected to help policy-makers involved in the provision of electricity in Namibia.

REFERENCES

- Al-Faris, A. R., 2002. The demand for electricity in the GCC countries. *Energy Policy* 30, 117-124.
- Ang, B. W., Goh, T. N., Liu, X. Q., 1992. Residential electricity demand in Singapore. *Energy* 17(1), 37-46.
- Bannon, J., 2006. *NamPower: Ten years and beyond* [Pamphlet]. Parklands, GDP Associates, United Kingdom.
- Bardsen, G., 1989. Estimation of long-run coefficients in error correction models. *Oxford Bulletin of Economics and Statistics* 51(3), 345-350.
- Beenstock, M., Goldin, E., Nabot, D., 1999. The demand for electricity in Israel. *Energy Economics* 21, 168-183.
- Bose, R.K., Shukla, M., 1999. Elasticities of electricity demand in India. *Energy Policy*, 27, 1377-146.
- De Vita, G., Endresen, K., Hunt, L.C., 2006. An empirical analysis of energy demand in Namibia. *Energy Policy*, 34, 3447-3463.
- Dincer, I., Dost, S., 1997. Energy and GDP. *International Journal of Energy Research* 21, 153-167.
- Donatos, G.S., Mergos, G.J., 1991. Residential demand for electricity: the case of Greece. *Energy Economics* pp. 41-47.
- Electricity Control Board, 2010. Annual report 2010. Windhoek, Electricity Control Board.
- Engle, R.F., Granger, C.W.J., 1987. Cointegration and error correction: representation, estimation and testing. *Econometrica* 55, 251-276.
- Filippini, M., Pachauri, S., 2004. Elasticities of electricity demand in urban Indian households. *Energy Policy* 32, 429-436.
- Holtedahl, P., Joutz, F.L., 2004. Residential electricity demand in Taiwan. *Energy Economics* 26, 201-224.
- Hondroyannis, G., 2004. Estimating residential demand for electricity in Greece. *Energy Economics* 26, 319-334.
- Hendry, D.F., Pagan, A., Sargan, J.D., 1984. Dynamic specification, in Z. Griliches and M. Intriligator (Eds), *Handbook of Econometrics* 2. Amsterdam: North Holland.
- Kavezeri, K., 2009. An analysis of the demand for electricity in Namibia. Unpublished Msc Thesis.
- Khan, M. A. and Qayyum, A., 2009. The demand for electricity in Pakistan. *OPEC Energy Review*, 70-96.
- Mah, J.S., 2000. An empirical examination of the disaggregated import demand of Korea- the case of information technology products. *Journal of Asian Economics* 11, 237-244.
- Ministry of Mines and Energy, 1998. White paper on energy policy. Windhoek, Ministry of Mines and Energy.

NamPower, 2009. Annual report 2006. Windhoek, NamPower.

Narayan, P.K., Smyth, R., 2005. The residential demand for electricity in Australia: an application of the bounds testing approach to cointegration. *Energy Policy* 33, 467-474.

Narayan, P. K., Smyth, R., Prasad, A., 2007. Electricity consumption in G7 countries: A panel cointegration analysis of residential demand elasticities. *Energy Policy*, 35, 4485-4494.

Nasr, G.E., Badr, E.A., Dibeh, G., 2000. Econometric modelling of electricity consumption in post-war Lebanon. *Energy Economics* 22, 627-640.

Pattichis, C. A., 1999. Price and income elasticities of disaggregated import demand: results from UECMs and an application. *Applied Economics*, 31, 1061-1071.

Pesaran, M. H., Shin, Y., Smith, R. J., 2001. Bounds Testing Approaches to the Analysis of Level Relationships. *Journal of Applied Econometrics* 16 (3), 289-326.

Sari, R., Ewing, B.T., Soytas, U., 2008. The relationship between disaggregate energy consumption and industrial production in the United States: An ARDL approach. *Energy Economics* 30, 2302-2313.

Yoo, S. H., Lee, J. S., Kwak, S. J., 2007. Estimation of residential electricity demand function in Seoul by correction for sample selection bias. *Energy Policy* 35, 5702- 5707.

Ziramba, E., 2008. The demand for residential electricity in South Africa. *Energy Policy* 36, 3460-3466.