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A reassessment of energy and GDP relationship: the case of Australia

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Abstract This paper investigates the long- and short-run relationships between energy consumption and economic growth in Australia using the bound testing and the ARDL approach. The analytical framework utilized in this paper includes both production and demand side models and a unified model comprising both production and demand side variables. The energy–GDP relationships are investigated at aggregate as well as several disaggregated energy categories, such as coal, oil, gas and electricity. The possibilities of one or more structural break(s) in the data series are examined by applying the recent advances in techniques. We find that the results of the cointegration tests could be affected by the structural break(s) in the data. It is, therefore, crucial to incorporate the information on structural break(s) in the subsequent modelling and inferences. Moreover, neither the production side nor the demand side framework alone can provide sufficient information to draw an ultimate conclusion on the cointegration and causal direction between energy and output. When alternative frameworks and structural break(s) in time series are explored properly, strong evidence of a bidirectional relationship between energy and output can be observed. The finding is true at both the aggregate and the disaggregate levels of energy consumption.

Keywords Energy consumption · GDP · Cointegration · Causality

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1 Introduction

The relationship between energy consumption and economic growth has been studied extensively in the empirical literature (see Ozturk 2010; Payne 2010a for a review). The aim of these studies has been to examine whether a relationship exists between energy consumption and output/income, and if yes, what is the direction of causality between them. The policy implications from the results of these studies are set forth as if energy consumption Granger causes output, direct measures on energy conservation would negatively affect economic growth (Karanfil 2009).[1] Early studies examined the hypothesis by using a bivariate model of energy and output and found no evidence of Granger causality running from energy consumption to output (Akarca and Long 1980; Kraft and Kraft 1978). Indeed, these studies found evidence of Granger causality in the opposite direction, i.e., output to energy use, coming to the conclusion that energy conservation may not be detrimental to economic growth. Subsequent studies adopted either a bivariate model (e.g. Chontanawat et al. 2008; Hu and Lin 2008) or a multivariate production function model constituting output, energy, capital and labour (e.g. Shahiduzzaman and Alam 2012; Stern 1993, 2000) or an energy demand model constituting energy, output and a measure of energy prices (e.g. Asafu-Adjaye 2000; Fatai et al. 2004). The superiority of a multivariate approach is that it reduces the potential omitted variable problem and, therefore, explores the additional channels through which energy and output are interlinked. Yet, results of the various studies are conflicting (Payne 2010c). It is argued that the differences in results could be attributed to the selection of the analytical tool, the econometric method, method of aggregation of energy types, sample size, resource endowment and so forth (Karanfil 2009; Payne 2010c; Stern 2011).

In this paper, we depart from the earlier studies in different ways. First, we extend the standard approach by employing both production and demand side models separately as well as a unified model consisting of both production and demand side variables for a same set of data. In this way, one can explore all possible channels through which energy can affect economic growth and vice versa. Secondly, we make more comprehensive treatment of structural break(s) in data as compared to previous studies. We employ both Zivot and Andrews (hereafter Z–A) (1992) and Lee and Strazicich (hereinafter L–S) (2003) endogenous structural break(s) unit root tests and incorporate the information in subsequent modelling and inference. This study finds that structural break(s) in data can significantly affect the estimation results. Traditionally, estimated models in most of the existing studies are based on the maintenance of the *priori* assumption of no structural break(s) in the time series. However, in view of historical changes, economic activities and energy consumption could go through one or more structural break(s) caused by domestic and global economic shocks (business cycle), changes in energy policy and fluctuations of energy prices (Chiou-Wei et al. 2008). Thirdly, we use data for unit energy prices for different energy types as opposed to the Consumer Price Index (CPI) used in most of the literature (Asafu-Adjaye 2000; Masih and Masih 1997). Note that energy constitutes only a very small weight in the CPI. Therefore, CPI may undermine the true movement of real energy prices. Once these factors are properly accounted for, strong evidence of a bidirectional relationship between energy and output can be observed. The finding is true both at aggregate and at disaggregate levels of energy consumption. This paper explores the issues by using Australian data for a period of last five decades.

[1] Empirical studies present four testable hypotheses based on the direction of causality [see Payne (2010a) for details].

The remainder of this paper is organized as follows. Following introduction, Sect. 2 illuminates the review of the literature. Section 3 describes the methodology and the data set. Section 4 provides the empirical results. Finally, Sect. 5 presents conclusion and policy implications.

2 Review of the literature

Recent literature such as Ozturk (2010) and Payne (2010a) provide comprehensive review of the existing literature. In this section, we narrow down our discussion on the literature based on the modelling frameworks utilized in the studies along with some attention on the structural break(s) in data. Two principal multivariate streams have emerged over time in the energy-output causality literature. One is the production side framework considering energy as a primary factor of production along with conventional inputs such as capital and labour. Another is the demand side framework which includes energy price as a third variable.

Stern (1993) first adopted the production side model to investigate the Granger causality between energy and output. The production side model is an amalgamation of the biophysical model, which considers energy as a sole factor of production, and the neoclassical growth model, which recognizes the role of only capital and labour as primary factors of production. Using the production framework, Stern (1993) found an evidence of Granger causality running from Divisia aggregate of energy use to economic growth in the United States. This result was different to that derived using the thermal aggregation of energy use, which indicated a causality that ran in the opposite direction, i.e., from economic growth to energy use. Stern (2000) subsequently found strong evidence of cointegration between energy and output in agreement with Stern's (1993) causality. Some other studies that have used such a production function approach to investigate the energy and GDP relationship include Yuan et al. (2008) for China, Ghali and El-Sakka (2004) for Canada, Soytas and Sari (2007) for Turkey, Warr and Ayres (2010), Payne (2011) for the United States and Shahiduzzaman and Alam (2012) for Australia. The findings of Ghali and El-Sakka (2004) proved the existence of cointegration and bidirectional causality between energy consumption and output in Canada. Yuan et al. (2008) found cointegration relationships at both aggregate and disaggregate levels of energy consumption in China and found Granger causality from electricity and oil consumption to output, but not from coal or total energy consumption to output. Soytas and Sari (2007) found the evidence of unidirectional causality running from electricity consumption to output in the Turkish manufacturing industry. Warr and Ayres (2010) redefined energy in terms of exergy (i.e. the amount of energy available for useful work) and the amount of useful work provided by energy inputs. They found evidence of unidirectional causality from both these measures of energy consumption to output. In contrast, Lee and Chien (2010) found no causality between these two variables in the United States and Germany; unidirectional causality from output to energy consumption in France and Japan; and unidirectional causality from energy consumption to output in Canada, Italy and the United Kingdom. They, therefore, concluded that energy conservation might hinder economic growth in Canada, Italy and the United Kingdom. At the disaggregated levels of fossil fuel consumption, Payne (2011) found evidence of no Granger causality from coal consumption to GDP, and unidirectional causality from real GDP to gas, and petroleum to GDP in the United States. Therefore, the results from the studies remain inconclusive.

Other stem of studies that have considered the demand side model, i.e., incorporating energy prices as a third variable in addition to energy and GDP, include, among others, Asafu-Adjaye (2000) for Asian developing countries, namely, India, Indonesia, Thailand and Philippines, Fatai et al. (2004) for Australia, Mahadevan and Asafu-Adjaye (2007) for a sample of 20 countries including Australia, Masih and Masih (1998) for Thailand and Sri Lanka and Levent (2007) for Turkey. The results of these studies are mixed (Payne 2010c). For example, Asafu-Adjaye (2000) found mixed results regarding the direction of causality for the sampled countries. Fatai et al. (2004) found evidence of unidirectional causality from GDP to coal, electricity and total final energy consumption in Australia. Nonetheless, most of the studies considered CPI or GDP deflator as a proxy for energy price. We overcome this limitation in this study by considering appropriate price series for different energy types.

Climent and Pardo (2007) extended the bivariate model by including employment, Brent oil price and CPI and found evidence of bidirectional causality between energy consumption and GDP in Spain. Oh and Lee (2004), on the other hand, found evidence of unidirectional causality from GDP to aggregate energy consumption in Korea using a demand side model that included real energy prices as a third variable. They found the same results in the case of the production side model. Considering a five-variable VAR model constituting GDP, energy, gross fixed capital formation, labour and real energy prices, Jin et al. (2009) found no evidence of any significant impacts of energy shock on real output growth in the United States. The use of gross fixed capital formation as a proxy of capital stock is problematic though as the former is a flow variable in contrast to the latter, which is a stock variable.

Major drawback of the aforesaid studies is the limited or no attention on the structural break(s) in the data even though macro-time series might undergo through such changes due adverse shocks. There are some other studies that considered the possibilities of structural break(s) in time series in the investigation of energy-GDP relationship, but these studies used a bivariate model. Altinay and Karagol (2004) employed Z–A (1992) and Perron's (1997) endogenous structural break tests to explore the possibility of structural break(s) in GDP and energy consumption series in Turkey for the 1950–2000 period. They found that both GDP and energy consumption series are trend stationary with structural break(s). The results were different in the case of conventional unit root tests such as Augmented Dickey–Fuller (hereinafter ADF) (Dickey and Fuller 1979, 1981) and Phillips–Perron (hereinafter P–P) (Phillips and Perron 1988), which indicated that the series are *I(1)*. Then, using the Hsiao (1981) version of the Granger causality test with de-trended series through the breakpoints obtained by Perron (1997), the study found no evidence of Granger causality between the two series. Lee and Chang (2005) employed Z–A and Perron (1997) tests along with the conventional unit root tests for GDP and aggregate and disaggregate energy consumption for Taiwan for the period 1954–2003. The need to consider the possibilities of structural break(s) in the cointegration test was further highlighted in this study.

Chiou-Wei et al. (2008) considered the case for some developing countries in Asia and the United States for the period 1954–2006 and conducted the Z–A structural break unit root test to detect the possible shift in regime. The study found evidence of stationarity with structural change in the first difference models. While the possible nonlinearity of the data was incorporated in the Granger causality tests, the cointegration analysis was carried out by implementing the methodology of Johansen (1991, 1995), which may have limitations in dealing with regime shifts (Esso 2010). The study found mixed results regarding the direction of causality for the countries in the sample. Hu and Lin (2008) also considered bivariate models for aggregated and disaggregate levels of energy consumption and GDP in Taiwan using quarterly data for the period 1982:1–2006:4. The null hypothesis of linear cointegration was strongly rejected by the threshold cointegration method developed by Hansen and Seo (2002). The study identified a significant asymmetric dynamic adjusting process between energy and GDP in Taiwan and the evidence of nonlinear cointegration between GDP and all disaggregate energy consumptions but oil. Esso (2010) employed the Z–A unit root test and Gregory and Hansen (1996a, b) threshold cointegration tests for seven Sub-Saharan countries for the period 1970–2007. The study found evidence of cointegration relationship with structural break in five countries in the sample. The model used in Esso (2010) is, however, a bivariate one.

As seen above, a common characteristic of the aforesaid models controlling for the possibilities of a structural break is the use of a bivariate framework comprised of only energy consumption and GDP. In addition, while Z–A and P–P tests for unit root have been widely used in the literature, they can only account for the possibility of one structural break. However, the time series processes under consideration may have gone through multiple structural break(s). In this case, Z–A and P–P tests suffer from loss of power in the estimation because of their restricted ability to account for only one break. Another criticism of Z–A and P–P unit root tests is that they do not allow the possibility of a structural break under the null, which could lead to a misinterpretation of the test results (Lee and Strazicich 2003). As argued by L–S (2003), the "rejection of null does not necessarily imply rejection of a unit root per se, but would imply rejection of a unit root without break" (p. 1082). The L–S test overcomes the limitation by proposing an endogenous two-break Lagrange multiplier (LM) unit root test that allows for breaks under both the null and alternative hypotheses.

For Australia, Narayan and Smyth (2005b) employed the Z–A unit root test for an annual time series data from 1966 to 1999. The Z–A test indicates the series in the study, i.e., electricity consumption per capita, real GDP per capita and an index of manufacturing sector employment series, are *I(1)*. The results are consistent with ADF and P–P tests. The study found statistically significant break(s) for electricity consumption in 1971 and 1990, employment in 1980 and real income in 1983. Considering a trivariate model comprising the above-mentioned three variables and applying the bound testing (Pesaran et al. 2001) approach, the study found evidence of cointegration only when electricity consumption is the dependent variable. Narayan and Smyth (2005c) estimated residential electricity demand model for Australia using annual data for the period 1969–2000. Using bound testing approach, the study found a cointegration relation among the variables. Previously, Akmal and Stern (2001) focused on residential energy demand in Australia and found a unique cointegration vector in the case of electricity and other fuels. Fatai et al. (2004) utilized a trivariate model by including energy prices along with GDP and aggregate and disaggregate levels of energy consumption for the period 1960–1999. Both Johansen (1991, 1995) and the bound testing approach utilized in this study indicated the evidence of unidirectional link from real GDP to total final energy consumption and electricity consumption.

In a recent study, Narayan et al. (2010) examined the unit root null hypothesis at both aggregate and sectoral levels of energy consumption in Australia for the period 1973–2007 using the L–S test. The study found aggregate energy series as stationary in most sectors of the economy both at the state and at the national levels. They concluded that energy conservation policies would only impact on energy consumption in the short-run. Shahiduzzaman and Alam (2012) utilized the production function model to examine the cointegration and causal relationship between aggregate energy consumption and output.

Using Divisia aggregation of energy use, the study found strong evidence of bidirectional causality between energy and GDP. The result is different in case of thermal measure of aggregate energy, which found the evidence of unidirectional causality from GDP to energy. The study, however, did not extend the analysis at disaggregated level of energy categories. Disaggregation of aggregate energy consumption to different energy vectors is important as each energy type possesses distinct qualities in performing useful economic tasks, and therefore, they are not easily substitutable (Berndt 1978; Cleveland et al. 2000). Accordingly, narrowing down the analysis to different energy types would be more informative, especially in the context of fuel-specific measures of energy conservation.

The present study differs from previous studies in a number of ways. We employ both Z–A and L–S endogenous structural break(s) unit root tests at both aggregate and disaggregate levels of energy consumption. Our modelling approach is also more comprehensive than the earlier studies as we consider production side and demand side models separately and a unified model in order to investigate the relationship between energy consumption at both aggregate and disaggregated levels and economic growth. Both long-run relationships and short-run dynamics are explored in this study. Finally, we make use of unit energy prices series for individual energy categories as opposed to the CPI used in most of the literature as a proxy for energy price.

3 Methodology

3.1 Models

The production function model constitutes energy as a separate input along with conventional inputs—capital and labour. We can write the aggregate production (Y_t) function at time t as:

$$Y_t = f(K_t, L_t, E_t) \tag{1}$$

where Y is real GDP, K is the capital input, L is the employment, and E denotes energy consumption. The subscript t symbolizes the time period.

The production function model used by Stern (1993) and majority of subsequent studies in this domain tested the existence of cointegration relationships by implementing the methodology of Johansen (1991, 1995) and then performed the Granger causality using the vector error correction (VEC) model approach (Ghali and El-Sakka 2004; Shahiduzzaman and Alam 2012; Soytas and Sari 2007; Stern 2000; Warr and Ayres 2010; Yuan et al. 2008). Other studies investigated Granger causality utilizing Toda and Yamamoto (1995) approach, which obviates the need of pretesting for cointegration (Bowden and Payne 2009, 2010; Payne 2009, 2010b; Payne and Taylor 2010; Shahiduzzaman and Alam 2012). A common focus of the studies investigating cointegration relationship using Johansen approach is to find the evidence of at least two long-run or cointegration relationships, of which one is clearly production function. Nonetheless, there is no consensus on the specification of the second cointegration relationship. For example, Stern (2000) and Shahiduzzaman and Alam (2012) modelled the second cointegration relationship as labour supply equation, while Ghali and El-Sakka (2004) normalized the other long-run relationship on energy consumption. To sum up, the second cointegration relationship can be interpreted by a factor demand/supply equation; however, in the absence of a relevant price variable, the model could be miss-specified. We overcome this deficiency by including prices of energy in the model, which can be expressed in terms of an energy demand equation, written as:

$$E_t = f(Y_t, K_t, L_t, P_E) \quad (2)$$

where, in addition to the previous notations, P_E is the real price of energy.

The third model we consider is a simplified demand model for energy and can be written as:

$$E_t = f(Y_t, P_E) \quad (3)$$

Equation (3) assumes the existence of a unique equilibrium demand for energy for a given price level, *ceteris paribus*. A time trend as a proxy technological progress is included in each estimated model from (1) to (3) when it is significant. Each model from (1) to (3) is estimated for aggregate and four major energy categories such as coal, oil, gas and electricity.

3.2 Data

The data for real GDP (*Y*) and capital stock (*K*) are collected from the Australian Bureau of Statistics (ABS). GDP series is the all industries chain value measures (2008 = 100) in AUS$ million. Capital is calculated as the chain volume measure of net capital stock (all industries) adjusted for capital utilization. Because time series data for capital utilization in Australia is not available, following Stern (1993) we use the employment rate as a proxy for capital utilization. Labour (*L*) is the civilian employment. The employment and labour data series are derived from the statistical database of the OECD. Energy consumption (*E*) is calculated as the total quantity (in energy units) of primary and derived fuels consumed minus the quantity of derived fuels produced. These data are collected from ABARE (2011b). Data for disaggregated energy categories such as, coal (*C*), gas (*G*), oil (*O*) and electricity (*E*) are collected from ABARE (2011b). Data for energy categories are in petajoules (PJ). Energy price for the total energy and oil consumption is proxied by the end-user price for diesel in Australian dollars obtained from the ABARES (2011a). Indexes for producer input prices for natural gas and electricity are collected from the ABS for the period 1970–2009 (cat no. 6427). Indexes for earlier years are estimated based on changes in the producer price index and consumer price index. The unit prices for gas in 2009 are taken from ABARES (2010), which is used to derive long-run unit prices for gas based on changes in the index numbers. Prices for Australian coal are collected online from IMF Commodity Prices (http://www.imf.org/external). Nominal energy prices are converted to the real by using GDP implicit price deflators (2009 = 100) collected from the ABS (cat no. 5204, Table 4). All the variables are expressed as natural logarithms. The full sample for the study is 1961–2009 with an exception for the model for gas, where the sample period considered is 1970–2009 based on the availability data.

3.3 Tests for stationarity

In order to assess the stationarity properties of the data, we employ both conventional and endogenous structural break(s) unit root tests. The conventional unit root tests are augmented Dickey and Fuller (ADF) (1979, 1981), Phillips and Perron (PP) (1988), and Kwiatkowski, Phillips, Schmidt and Shin (KPSS) (Kwiatkowski et al. 1992) (see Maddala and Kim 1998 for an application of these tests). Major limitation with the aforesaid unit root tests is that they fail to capture the possible structural break(s) in data. This is a serious problem as structural change in data may bias the test (Enders 2004). A common viewpoint about the macroeconomic time series is that they are first difference stationary (e.g. Nelson and Plosser 1982). However, re-examining the Nelson and Plosser (1982) data for unit root while capturing the effect of a known exogenous structural break, Perron (1989) found that most of the series (11 out of 14) were indeed stationary. Following the subsequent developments in the unit root tests with possible structural break, Z–A (1992) proposed a unit root test, where the breakpoint is endogenously estimated and, therefore, reduces the problem of data mining (Maddala and Kim 1998). However, the Z–A test has the limitation in that structural break is considered only under alternative hypothesis and can illustrate the possibility of only one break. Accordingly, critical values in Z–A test are derived assuming no break under the null hypothesis. Therefore, a spurious rejection of the null might occur (Maslyuk and Smyth 2008). As discussed above, the Lagrange multiplier unit root test introduced by L–S can overcome the limitation and also controls for multiple breaks in data. We employ both Z–A and L–S tests to explore the possibilities of structural break(s) as well as to test unit root hypothesis further alongside the conventional unit root tests.

3.4 Tests for cointegration

We adopt autoregressive the distributed lag (ARDL)-based bound testing approach as proposed by Pesaran and Pesaran (1997) and Pesaran et al. (2001) to test for cointegration. The ARDL model can be applied irrespective of whether the data are *I*(*0*) or *I*(*1*). The small sample properties of the bound testing approach for cointegration test are far superior to the conventional multivariate cointegration procedure (Fosu and Magnus 2006; Narayan and Smyth 2005a; Narayan and Narayan 2005; Pesaran et al. 2001; Sari et al. 2008). Pesaran and Shin (1999) have shown that the ARDL-based estimators of the long-run coefficients are super-consistent in case of small sample sizes. The ARDL approach also effectively corrects for possible endogeneity of the variables by allowing simultaneous estimation of the long- and short-run components within a VEC model (Squalli 2007). Moreover, ARDL approach allows to apply general-to-specific modelling technique to estimate consistent parameters of the model.

We follow the procedure as outlined in Pesaran and Pesaran (1997) to test for cointegration and to estimate the ARDL model. Apart from applying two alternative lag selection criteria such as the Akaike information criterion (AIC) and Schwarz bayesian criterion (SBC), a further investigation of the residual properties is made to decide about the optimal lags. Typically, the SBC selects a shorter optimal lag than the AIC, but the model for a shorter lag may not illustrate adequate residual properties. The optimal lag is chosen by the criteria that reflects relatively better residual properties. In case the residual properties are found adequate in both AIC and SBC models, the results from the shorter lag model are reported. The residual properties for each model are investigated through the Lagrange multiplier test of residual serial correlation, tests for heteroscedasticity and standard errors of the regression. A maximum lag length of four is considered for the estimation. The sub-samples are determined based on the results for structural break(s) as found in Z–A and L–S tests.

The null hypothesis of no cointegration among the variables can be tested by applying general *F* statistics and comparing them with critical values in Narayan (2005). The critical *F* values provided by Narayan (2005) are preferred than the values provided by Pesaran et al. (2001) as the former is suited for a sample size of 30–80 as compared to 500–1,000 observations for the latter. The bound testing approach implies rejection of the null if the computed F statistics is higher than the upper bound of the critical values in case of $I(1)$ variable (see Pesaran et al. 2001 for details on bound testing procedure).

4 Empirical results

4.1 Results of the unit root tests and structural break(s)

Considering the sample size (i.e. $T = 49$), we set the maximum lag length for the unit root tests to four using a $T^{1/3}$ formula, as suggested by Lütkepohl (1993). The estimation results of the ADF, PP and KPSS tests generally support the notion that the variables are non-stationary in level and stationary in the first difference form.[2] The results of the Z–A test indicate that none of the series (with the exception of gas) is trend stationary in level with the consideration of one structural break (Table 1). On the other hand, the first difference series are clearly stationary. GDP growth seems to have structural breaks in the mid-1970s and 1980s, both of the dates correspond to the business cycle downturn of the Australian economy (ECRI 2011). The break in 1984 is consistent with Narayan and Smyth (2005b). Growth in energy consumption appears to be affected by the oil price shocks in the mid-1970s and business cycle downturn in the mid-1980s. The results from the Z–A test indicate that the energy, GDP and other variables in Australia have possibly been affected by two major changes—one in the mid-1970s, possibly linked to the oil crisis, and another in the 1980s—over the last five decades. Given the information, we now proceed with the L–S test to examine the unit root hypothesis with the possibility of two structural breaks, which is presented in Table 2.

As seen in Table 2, the L–S test also indicates that the series are first difference stationary in both *crash* and *break* models. The finding on the first break (i.e. 1971) for GDP growth in the L–S test is roughly consistent with that of Z–A model with intercept only. The economic slowdown in the early 1970s might reflect the declining share of trade to the Australian economy due to the fall in terms of trade. By the mid-1970s, terms of trade improved again as supply constraints pushed up prices in the world commodity market (Stevens 2008). The L–S test indicates a second break in the GDP growth in the early 1990s. This break might correspond to the monetary policy-induced recession in the Australian economy during 1990–1992 (ECRI 2011; Narayan and Smyth 2005a). Note that both the recession in the mid-1970s and the recession in the early 1990s describe the periods of economic downturn affecting much of the world, including major economies like the United States, Germany, France, the United Kingdom, Italy, Japan, Australia and so forth (ECRI 2011). The significance of the *crash* and *break* dummies indicates that the energy consumption at both aggregate and disaggregate levels is significantly affected by oil price shocks in the mid-1970s.

4.2 Estimation results

4.2.1 Aggregate energy consumption

Table 3 presents the results of ARDL bound tests for the models including aggregate energy consumption. The computed F statistics for the full sample model is 3.87 for Eq. 1,

[2] The results are not reported here to conserve space, but can be available from the authors upon request.

Table 1 Z–A test for unit root with one endogenously determined structural break

Series	Intercept				Intercept and trend			
	Level		First diff		Level		First diff	
	Break	*t*-stat	Break	*t*-stat	Break	*t*-stat	Break	*t*-stat
Y	1983	−2.88	1973	−6.32[a]	1975	−3.68	1984	−6.10[a]
K	1990	−4.79	1974	−5.50[a]	1969	−3.65	1974	−5.56[b]
L	1981	−4.09	1974	−5.43[a]	1974	−4.44	1984	−5.46[b]
E	1969	−2.37	1975	−5.70[a]	1974	−3.93	1985	−5.83[a]
C	2002	−1.09	1985	−4.93[a]	2002	−2.92	1997	−5.42[a]
O	1980	−4.87[a]	1975	−4.86[a]	2002	−3.72	1984	−6.77[a]
G	1977	−6.13[a]	1983	−6.44[a]	1981	−6.88[a]	1993	−5.70[a]
EL	1969	−2.53	1976	−6.96[a]	1969	−2.54	1986	−7.05[a]
P_O	1987	−3.31	1984	−5.17[b]	1987	−3.52	1982	−5.38[b]
P_G	1971	−4.36	1980	−5.87[a]	1982	−4.67	1980	−6.10[a]
P_C	1983	−3.14	1982	−7.45[a]	1982	−3.01	1982	−7.62[a]
P_{EL}	1981	−3.46	1977	−4.85[b]	1981	−5.89[a]	1977	−4.95

Superscripts a and b denote the significance level at 1 and 5 %, respectively. Test results are for one augmented lag

which does not provide evidence of cointegration. Including a trend dummy to incorporate structural break in the early 1970s (DT_{72})[3] results in the rejection of the null hypothesis of no cointegration at 1 % significance level. Considering a sub-sample from early 1970s with and without the dummy variable provides the similar results for cointegration. Time trend is found significant in Eq. 1.

In a variant of Eq. 1, where energy is considered as dependent variable and GDP, capital and labour as explanatory variables, the null hypothesis of no cointegration is not rejected (Table 3). Estimating Eq. 2 turns out similar result (Table 3). A sub-sample model for 1976–2009 to reflect post-oil crisis scenario does not cointegrate either (results not reported). However, strong evidence of cointegration is found between energy consumption, GDP and energy prices as shown from the estimation results for Eq. 3. The results are for only intercept as the time trend was found insignificant, therefore excluded from the model. We also report results for the bound testing where capital and labour are taken as dependent variables in Eq. 1. The computed F statistics unanimously indicate the existence of a cointegration between energy and these two conventional factors of production. The estimated ARDL models including the aggregate energy consumption are presented in Table 4.

The first panel in Table 4 presents the estimation results of the production function model for aggregate energy consumption (Eq. 1). The full sample (1961–2009) ARDL model does not pass through the heteroscedasticity test (results not reported). An investigation of the residual properties of the estimated model shows that it crosses two standard error bands in 1966 and 1971 (Fig. 1). This is in line with the results of structural break(s) as shown by the L–S test, which indicates a significant structural break in Y series in 1971 (Table 2). Note that the full sample model without incorporating structural break in the early 1970s did not show a cointegration among the variables (Table 3). Estimating

[3] The structural break (s) dummies are included in a model based on the Z–A and L–S tests.

Table 2 L–S test for unit root with two endogenously determined breaks

Series	Crash (intercept)						Break (intercept and trend)					
	Level			First difference			Level			First difference		
	Break 1	Break 2	t	Break 1	Break 2	t	Break 1	Break 2	t	Break 1	Break 2	t
Y	1964	$1969^{b(D)}$	−1.69	1972	1991	-6.57^{a}	1970	1995	−4.45	$1971^{a(T)}$	$1993^{a(T)}$	-6.57^{a}
K	$1970^{c(D)}$	$1983^{a(D)}$	−2.05	1974	1991	-4.69^{a}	1970	$1991^{a(T)}$	-5.27^{c}	1974	$1990^{a(D),\ c(T)}$	-5.71^{b}
L	$1982^{a(D)}$	1991	−2.47	1973	$1985^{b(D)}$	-5.38^{a}	1967	$1977^{a(T)}$	−4.85	$1983^{a(D,\ T)}$	$1992^{b(T)}$	-5.79^{b}
E	1964	$1968^{c(D)}$	−.97	$1977^{c(D)}$	$1984^{c(D)}$	-5.31^{a}	1974	$1994^{a(T)}$	−3.51	$1975^{a(D),\ c(T)}$	$1983^{a(D,\ T)}$	-7.11^{a}
C	$1976^{c(D)}$	$1984^{b(D)}$	−2.15	$1992^{c(D)}$	2002	-4.89^{a}	$1975^{b(D),\ a(T)}$	2002	−3.95	$1992^{c(T)}$	1998	-10.53^{a}
O	$1973^{a(D)}$	$1982^{c(D)}$	−1.69	1969	1977	-5.32^{a}	$1974^{c(D),\ a(T)}$	$1982^{a(D)}$	-5.37^{b}	$1980^{b(D)}$	1989	-6.25^{a}
G	1983	1989	1.23	1977	1982	−.81	$1977^{b(D),\ a(T)}$	$1989^{a(D,\ T)}$	−3.98	$1978^{a(T)}$	$1990^{a(D,\ T)}$	-5.40^{a}
EL	1983^{b}	$2002^{a(D)}$	−1.42	$1977^{b(D)}$	1989	-3.76^{c}	$1976^{b(D),\ a(T)}$	1986	−4.83	1975	$1990^{c(D)}$	-7.13^{a}
P_O	1975	$1988^{b(D)}$	−2.88	1974	1994^{b}	-5.12^{a}	$1978^{a(T)}$	$1988^{b(D),\ a(T)}$	−4.62	$1979^{b(D,T)}$	$1988^{a(DT)}$	-7.03^{a}
P_G	1975	1983	−1.99	1978	$1982^{b(D)}$	-5.60^{a}	$1975^{a(T)}$	$1984^{a(T)}$	−4.72	$1969^{b(T)}$	$1982^{b(D)}$	-6.24^{a}
P_C	$1982^{b(D)}$	$2004^{b(D)}$	−2.25	1986	$2002^{a(D)}$	-3.60^{c}	$1981^{a(T)}$	$2004^{a(D,\ T)}$	−4.63	$1973^{a(D,\ T)}$	$1984^{a(D,\ T)}$	−4.76
P_{EL}	1978	$1982^{b(D)}$	−2.94	1976	$1983^{c(D)}$	-4.92^{a}	$1975^{b(D)}$	$1984^{a(T)}$	−4.62	$1979^{b(D),\ a(T)}$	$1983^{c(D),\ a(T)}$	-6.79^{a}

Superscripts a, b and c denote the significance level at 1, 5 and 10 %, respectively. Critical values 1 %—4.545; 5 %—3.842; and 10 %—3.504 for the crash (intercept) model and 1 %—5.823; 5 %—5.286; and 10 %—4.898 for the break (intercept and trend) model. Critical values for the dummy variables follow standard normal distribution (1 %—2.575; 5 %1.96; and 10 %—1.645). Superscripts D and T in the parentheses represent break in the intercept and break in the slope and superscripts a, b and c preceding them show the significance level. As for example, $1969^{b(D,T)}$ indicates that the break is significant in 1969 in both intercept and slope. Optimal lag lengths are determined by AIC criterion as decided in the ADF test for unit root

Table 3 Bound tests of cointegration: aggregated energy consumption

Model	Sample	Break dummy	Calculated F statistics	Decision
$Y\|K, L, E$	1961–2009	No	3.87	Not cointegrated
$Y\|K, L, E$	1961–2009	DT_{72}	9.53[a]	Cointegrated
$Y\|K, L, E$	1972–2009	No	9.65[a]	Cointegrated
$Y\|K, L, E$	1972–2009	DT_{84}	10.04[a]	Cointegrated
$E\|Y, K, L$	1961–2009	No	2.76	Not cointegrated
$E\|Y, K, L, P_E$	1961–2009	DT_{84}	1.67	Not cointegrated
$E\|Y, P_E$	1961–2009	No	10.53	Cointegrated
$K\|Y, L, E$	1961–2009	No	5.77[b]	Cointegrated
$L\|Y, K, E$	1961–2009	No	8.32[a]	Cointegrated
$L\|Y, K, E$	1961–2009	DT_{84}	7.03[a]	Cointegrated

Superscripts a, b and c denote the significance level at 1, 5 and 10 %, respectively. Critical values are taken from Narayan (2005). DT_{tb} is trend dummy, where tb denotes the year when the dummy starts. $DT_t = t-tb$, if $t \geq tb$, 0 otherwise

the model for 1972–2009 removes the heteroscedasticity problem along with satisfying the serial correlation property. The results are reported in Panel I in Table 4. An appropriate model selection criterion is AIC as the model selected by alternative SBC does not pass through the serial correlation chi-square (χ^2) test [$\chi^2 = 3.42(.064)$].[4] The estimation results show significant and expected negative coefficient of the error correction term. The coefficient value of the *ecm* (-1) is $-.35$, indicating that about 35 per cent of the deviations from long-run equilibrium is corrected for 1-year time. The estimation results indicate that energy and capita derive short-run output, but in the long-run, capital comes with unexpected negative sign. However, capital comes with expected positive and significant coefficient when a trend dummy from 1984 is included in the model (Panel II, Table 4) to capture the possible structural break in the mid-1980s as indicated by the Z–A test (Table 1). SBC is the preferred lag selection criterion in this case as the shorter lags SBC (1,1,0,0) model as compared to longer lags AIC (1,1,3,1) model saves some degrees of freedom and also found adequate in terms of serial correlation and heteroscedasticity properties. The sign and significance of the *ecm* (-1) in both Panels I and II in Table 4 further confirms the existence of cointegration among energy, output and other primary factors of production as found in Table 3. Figure 2 shows a reasonable fit of the estimated model with actual GDP growth series.

Panel III in Table 4 reports the estimated ARDL model for Eq. 3 for the period 1961–2009. The model satisfies adequately with the diagnostic properties. The income and price effects show expected positive and negative signs, respectively, in both short-run and long-run. Estimation for long-run coefficients indicates that income drives long-run energy demand in Australia. In the short-run, both income and price affect the demand for energy. In this model, while the *ecm*(-1) coefficient comes with an expected negative sign, it is found to be insignificant. Estimation of the model for post-oil crisis period (1975–2009) shows expected negative sign and significance of the error correction term (Panel IV, Table 4).

We also estimated the ARDL model for K and L as dependent variables to explore their interaction with energy consumption. The results are reported in Panels V and VI in

[4] Detailed results can be obtained from the authors upon request.

Table 4 Estimated ARDL models: aggregate energy consumption

Particulars	Long-run coefficients			Short-run : selected coefficients		
	Regressor	Coeff	t (p value)	Regressor	Coeff	t (p value)
Panel I Equation 1	*K*	−36	−1.85 (.074)	Δ*K*	.53	3.79 (.001)
Sample 1972–2009	*L*	35	1.23(.229)	Δ*L*	.12	1.02 (.317)
Dep. variable *Y*	*E*	.29	1.78(.087)	Δ*E*	.25	2.32 (.027)
AIC-ARDL (2,2,0,1)	Constant	9.60	2.30(.029)	*ecm(−1)*	−.35	–3.74 (.001)
	Trend	.03	4.39(.000)			
[*Diagnostics tests* R^2 = .82; serial cor. $\chi^2(1)$ = .011(.92); hetero. $\chi^2(1)$ = .38(.54)]						
Panel II Eq. 1	*K*	.44	3.31 (.002)	Δ*K*	.87	6.44 (.000)
Sample 1972–2009	*L*	−.07	−.29 (.773)	Δ*L(−1)*	−.03	−.29 (.775)
Dep. variable *Y*	*E*	.45	3.44 (.002)	Δ*E*	.21	4.08 (.000)
SBC-ARDL (1,1,0,0)	Constant	4.42	1.68 (.103)	*ecm(−1)*	−.46	−4.04(.013)
	DT84	.01	4.52 (.000)			
[*Diagnostics tests* R^2 = .76; serial cor. $\chi^2(1)$ = .93E−3(.98); hetero. $\chi^2(1)$ = 1.1(.30)]						
Panel III Eq. 3	*Y*	.64	2.93 (.006)	Δ*Y*	.56	5.16 (.000)
Sample 1961–2009	E_p	−.58	−.88 (.383)	ΔE_p	−.02	−2.91 (.006)
Dep. variable *E*	Constant	6.83	.68 (.498)	*ecm(−1)*	−.04	−1.03 (.310)
AIC-ARDL (1,1,0)						
[*Diagnostics tests* R^2 = .70; serial cor. $\chi^2(1)$ = 2.36(.12); hetero. $\chi^2(1)$ = .98 (.33)]						
Panel IV Eq. 3	*Y*	.68	16.27 (.000)	Δ*Y(–1)*	.27	1.65 (.111)
Sample 1975–2009	E_p	−.14	−1.71 (.099)	ΔE_p	−.03	−3.17 (.004)
Dep. variable *E*	Constant	.88	.68 (.504)	*ecm(–1)*	−.19	−2.09 (.046)
AIC-ARDL (2,3,0)						
[*Diagnostics tests* R^2 = .68; serial cor. $\chi^2(1)$ = .004(.95); hetero. $\chi^2(1)$ = .36 (.55)]						
Panel V Eq. 1	*Y*	.12	−.50 (.618)	Δ*Y*	.26	3.12 (.004)
Sample 1961–2009	*L*	1.49	4.87 (.000)	Δ*L*	.73	6.81 (.000)
Dep. variable *K*	*E*	.44	3.99 (.000)	Δ*E*	−.14	−1.86 (.071)
AIC-ARDL (2,1,4,4)	Constant	−13.98	−4.50 (.000)	Δ*E(−1)*	−.29	−3.46 (.002)
	Trend	−.008	−2.26 (.031)	*ecm(−1)*	−.28	−4.70 (.000)
[*Diagnostics tests* R^2 = .94; serial cor. $\chi^2(1)$ = .129 (.72); hetero. $\chi^2(1)$ = .54(.46)]						
Panel VI Eq. 1	*Y*	.20	2.21 (.035)	Δ*Y*	−.02	−.16 (.873)
Sample 1961–2009	*K*	.53	7.53 (.000)	Δ*K*	.82	6.92 (.000)
Dep. variable *L*	*E*	−.35	−4.39 (.000)	Δ*E(−1)*	.32	3.75 (.001)
AIC-ARDL (1,1,4,4)	Constant	8.36	11.15 (.000)	Δ*E(−2)*	.31	−3.68 (.001)
	Trend	.003	−1.40 (.172)	*ecm(−1)*	−.64	−5.96 (.000)
[*Diagnostics tests* R^2 = .90; serial cor. $\chi^2(1)$ = .052 (.82); hetero. $\chi^2(1)$ = 1.17 (.28)]						

R^2 values refer to the error correction model, while the other diagnostic test statistics refer to the underlying ARDL equation

Table 4. The estimation results indicate that energy drives both capital and labour in both short-run and long-run. Estimation of the long-run models suggests the possible complementarity between energy and capital and substitutability between energy and labour. The error correction terms are highly significant along with expected negative signs in both cases.

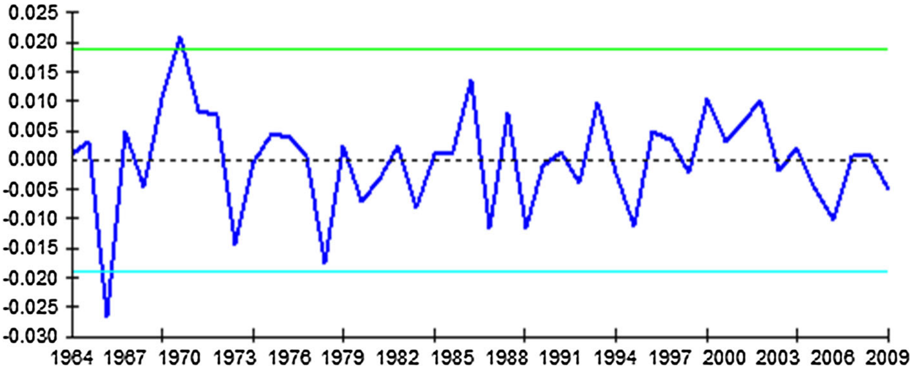

Fig. 1 Plot of residual and two standard errors bands

To sum up, the estimation results for the bound testing and subsequent ARDL models with aggregate energy consumption indicate that energy is a key driver of economic growth in both short-run and long-run in Australia. On the other hand, economic growth stimulates energy consumption. In the same vein, the estimation results indicate that energy seems to interact significantly with capital and labour in the production process. Both income and price determine the energy demand in Australia, whereas the estimated short-run elasticity for energy prices is relatively lower than that of long-run elasticity. What follows next is the repetitions of the above analytical steps in the case of disaggregate energy categories.

4.2.2 Disaggregate energy consumption

Each energy vector varies substantially in terms of performing useful economic works and generating environmental outcomes. The use of aggregate energy data does not capture the effects of these differences and fails to identify the impact of a specific type of energy on output. It is, therefore, value adding to extend the analysis at the disaggregate levels. Table 5 reports the results for the bound testing approach of cointegration for the models for various disaggregate levels of energy consumption. Using the production function model (Eq. 1) and considering coal as a measure of energy consumption, the null

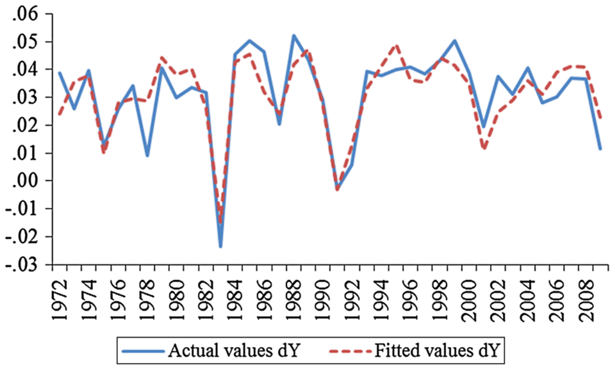

Fig. 2 Actual and fitted values of the model: Eq. 1

Table 5 Bound tests of cointegration: disaggregate energy consumption

Tests	Sample	Break dummy	Calculated F statistics	Decision
$Y\|K, L, C$	1961–2009	DT_{72}	10.88[a]	Cointegrated
$Y\|K, L, C$	1972–2009	No	5.85[c]	Cointegrated
$C\|Y, K, L$	1961–2009	No	4.34[c]	Cointegrated
$Y\|K, L, O$	1961–2009	DT_{72}	10.56[a]	Cointegrated
$Y\|K, L, O$	1975–2009	No	4.74[c]	Cointegrated
$O\|Y, K, L, Po$	1961–2009	DC_{80}	6.86[a]	Cointegrated
$O\|Y, K, L, Po$	1961–2009	No	6.55[a]	Cointegrated
$O\|Y, K, L$	1961–2009	No	5.09[b]	Cointegrated
$O\|Y, Po$	1961–2009	No	8.59[a]	Cointegrated
$Y\|K, L, G$	1970–2009	No	2.27	Not cointegrated
$Y\|K, L, G$	1976–2009	No	2.35	Not cointegrated
$G\|Y, P_G$	1970–2009	No	64.71[a]	Cointegrated
$G\|Y, P_G$	1976–2009	No	6.97[a]	Cointegrated
$Y\|K, L, EL$	1961–2009	No	1.95	Not cointegrated
$Y\|K, L, EL$	1961–2009	DT_{72}	7.34[a]	Cointegrated
$Y\|K, L, EL$	1978–2009	DC_{91}	5.36[b]	Cointegrated
$EL\|Y, P_{EL}$	1978–2009	No	6.34[a]	Cointegrated

Superscripts a, b and c denote the significance level at 1, 5 and 10 %, respectively. Critical values are taken from Narayan (2005). DT_{tb} and DC_{tb} are trend intercept dummies, respectively, where tb denotes the year when the dummy starts. $DC_t = 1$, if $t \geq tb$, 0 otherwise; $DT_t = t - tb$, if $t \geq tb$, 0 otherwise

hypothesis of no cointegration among the variables is rejected for the full sample model including a dummy for structural break in the early 1970s (DT_{72}). The result is same for the sub-sample 1972–2009 without incorporating a break dummy. An evidence of cointegration is found when coal is considered as dependent variable using a production function framework. We do not find any cointegration when coal is a dependent variable in the energy demand models (results are not reported here).

In the case of the models including oil as a measure of energy consumption, the null hypothesis of no cointegration is rejected in all three models (Eqs. 1, 2 and 3). The result is same when considering a structural break in GDP in the early 1970s. For the production function model including gas, the null hypothesis of no cointegration is not rejected for the full sample model (1970–2009) or for the sub-sample (1976–2009) to account for a post-oil crisis period. This could be due to a shorter sample for gas based on the availability of data and considerably a lower proportion of gas consumption to total energy consumption as compared to other fuels in Australia. The result for cointegration is, however, different in the case of energy demand model for gas, where the null of no cointegration is rejected for either of the samples (Table 5). Note that the evidence of cointegration is not evident when capital and labour are included in the energy demand model (Eq. 2) for coal and gas (not reported here). For electricity, the null of no cointegration is not rejected in the full sample model without controlling for any structural break. Incorporation of trend dummy from 1972 (DT_{72}) rejects the null hypothesis of no cointegration at 1 % significance level. An evidence of cointegration is also found in the case of production function model or energy demand models for electricity in the post-oil crisis period (Table 5).

The first panel in Table 6 reports the ARDL estimation results for production function model (Eq. 1) which includes coal consumption as a measure of energy use. An appropriate model selection criterion here is the SBC as the model selected by the AIC seems to exhibit heteroscedasticity problem in the residual. The sample period of the model is 1972–2009 as the L–S test indicated a structural break in GDP in 1971. As seen in the table, coal seems to play an important role on economic growth in both short-run and long-run. The error correction term shows highly significant and the expected negative sign. The feedback effects from GDP to coal are shown in Panel II in Table 6. The model is for full sample period as no significant structural break in growth in coal consumption is observed in the 1970s. The appropriate model selection criterion here is the AIC, as the serial correlation and heteroscedasticity properties of this criterion are similar to those of the SBC, the model based on the AIC criterion poses lower standard error of regression and residual sum square (not reported). The results indicate that the GDP is a forcing variable to determine coal consumption in both short-run and long-run. The error correction term comes with significant and an expected negative sign.

Estimation of the ARDL models for oil consumption is presented in Panels III and IV in Table 6. Panel III in the table shows the results for the production function model (Eq. 1).

Table 6 Estimated ARDL models for coal and oil

Particulars	Long-run coefficients			Short-run : selected coefficients		
	Regressor	Coeff	t (p value)	Regressor	Coeff	t (p value)
Panel I Eq. 1	K	−.50	−3.05 (.005)	ΔK	.55	4.01 (.000)
Sample 1972–2009	L	.55	2.81 (.009)	$\Delta K(-1)$	.51	2.66 (.012)
Dep. variable Y	C	.27	2.38 (.024)	ΔL	.22	2.32 (.027)
SBC-ARDL(2,2,0,0)	Constant	8.92	2.76 (.010)	ΔC	.10	2.88 (.007)
	Trend	.03	5.53 (.000)	$ecm(-1)$	−.39	−5.02 (.000)
[*Diagnostics tests* $R^2 = .82$; serial cor. $\chi^2(1) = .26(.61)$; hetero. $\chi^2(1) = 1.66(.20)$]						
Panel II Eq. 1	Y	2.91	2.11 (.045)	ΔY	.64	1.87 (.072)
Sample 1961–2009	K	−.76	−1.08 (.290)	$\Delta Y(-3)$	.61	2.13 (.042)
Dep. variable C	L	2.77	−1.46 (.156)	$\Delta K(-3)$	−1.92	−3.41 (.002)
AIC-ARDL (3,4,4,4)	Constant	23.48	1.35 (.188)	$\Delta L(-3)$	1.75	3.69 (.001)
				$ecm(-1)$	−.20	−2.21 (.035)
[*Diagnostics tests* $R^2 = .68$; serial cor. $\chi^2(1) = .129(.72)$; hetero. $\chi^2(1) = .54(.46)$]						
Panel III Eq. 1	K	−.217	−1.31 (.201)	ΔK	.58	3.99 (.000)
Sample 1972–2009	L	.42	2.22 (.034)	ΔL	.20	1.92 (.064)
Dep. variable Y	O	.11	1.75(.090)	ΔO	.05	1.55 (.131)
AIC-ARDL (1,2,0,0)	Constant	8.08	2.69(.010)	$ecm(-1)$	–.47	−4.87 (.000)
	Trend	.03	5.26(.000)			
[*Diagnostics tests* $R^2 = .77$; serial cor. $\chi^2(1) = .132(.72)$; hetero. $\chi^2(1) = .59(.44)$]						
Panel IV Eq. 3	Y	.56	8.47 (.000)	ΔY	.63	3.09 (.004)
Sample 1961–2009	Po	−.39	−3.18 (.003)	ΔPo	−.02	−.72 (.475)
Dep. variable O	Constant	4.43	3.52 (.001)	$\Delta Po(-1)$	.06	2.20 (.034)
SBC-ARDL (1,1,2)				$ecm(-1)$	−.16	−4.74 (.000)
[*Diagnostics tests* $R^2 = .73$; serial cor. $\chi^2(1) = .65(.42)$; hetero. $\chi^2(1) = 1.69(.19)$]						

R^2 values refer to the error correction model, while the other diagnostic test statistics refer to the underlying ARDL equation

A time trend is included in the model as it is found significant. Similar to the case of aggregate energy, estimation of the full sample model exhibits a heteroscedasticity problem. Estimation of the Eq. 2 does not eliminate the problem, nor does it show an improvement of results. A sub-sample model for Eq. 1 is found to remove the problem with heteroscedasticity in residuals and provides robust estimation. Panel III in the table indicates the positive role of oil consumption in economic growth in the long-run. No short-run impact from oil consumption to economic growth is observed in this estimation. The error correction term comes with the expected negative sign and is found to be significant at 1 % level. Estimation results for the oil demand equation are presented in Panel IV, Table 6. Both income and price effects are found to be significant in driving oil demand in the long-run, while income effects play a major role in driving oil demand in the short-run. Both the AIC and SBC criteria suggest the same lag length for the model estimation.

Panel I in Table 7 presents the estimated ARDL gas demand model for the full sample. While the estimated model exhibits adequacy in terms of residual properties, the error correction term comes with expected negative sign, but with insignificant coefficient. Re-estimating the model for post-oil crisis period shows expected negative sign and significant coefficient (Panel II). Both price and income effects are found to derive the demand for gas in long- and short-run. Unlike other energy categories, the long-run price elasticity of demand for gas is more than unity reflecting an elastic demand. The ARDL estimation results for the models for electricity consumption are presented in Panels III and IV in Table 7. Panel III shows the estimation of Eq. 1 by including electricity as a measure of energy consumption. An appropriate lag selection criterion is the AIC as the estimated model based on the SBC criterion exhibits heteroscedasticity problem in residuals. As seen in the Panel III, Table 7, electricity drives economic growth in Australia in both short-run and long-run. The error correction term is found significant at 1 % level along with the expected negative sign. The estimation results of the demand side model are presented in Panel IV in Table 7. The error correction term for the full sample model (1961–2009) shows insignificant coefficient despite having expected negative sign (results not reported here). As both Z–A and L–S tests indicate structural break in electricity consumption in the mid-1970s, a sub-sample model is estimated for the period 1978–2009. The error correction term now becomes significant along with the expected negative sign. The estimated results indicate that the GDP drives electricity consumption in Australia in both short-run and long-run, while price is the short-term driving force of electricity consumption in the short-run.

5 Conclusion and policy implications

This paper investigates the long-run relationship and short-run dynamics between energy consumption and GDP in Australia using the bound testing and the ARDL approach. For the first time in the literature, we employ both production side and demand side models and a unified model comprising both production and demand side variables for a single set of data. In addition, unit root hypothesis was tested by applying recent advances in techniques as proposed by Z–A (1992) and L–S (2003) accounting for the possibilities of one or more structural break(s) in time series determined endogenously. We also make use of the unit energy price data as opposed to the CPI used as a proxy of energy prices in the literature. The relationship was investigated at aggregate as well as at major energy categories such as coal, oil, gas and electricity.

Table 7 Estimated ARDL model for gas and electricity

Particulars	Long-run coefficients			Short-run : selected coefficients		
	Regressor	Coeff	t (p value)	Regressor	Coeff	t (p value)
Panel I Eq. 3	Y	.78	4.72 (.000)	ΔY	.92	3.06 (.005)
Sample 1970–2009	Pg	−1.34	−2.47 (.021)	$\Delta Y(-1)$	.51	1.59 (.123)
Dep. variable G	Constant	7.17	1.23 (.231)	ΔPg	−.31	−3.88 (.000)
AIC-ARDL (3,2,1)				$ecm(-1)$	−.06	−1.46 (.156)
[*Diagnostics tests* R^2 = .88; serial cor. $\chi^2(1)$ = 1.59(.21); hetero. $\chi^2(1)$ = .51(.48)]						
Panel II Eq. 3	Y	.64	3.93 (.001)	ΔY	.63	2.36 (.026)
Sample 1976–2009	Pg	−1.42	−3.19 (.004)	$\Delta Y(-1)$	.57	1.95 (.062)
Dep. variable G	Constant	9.69	1.82 (.080)	ΔPg	−.22	−4.54 (.000)
AIC-ARDL (3,3,0)				$ecm(-1)$	−.16	−4.99 (.000)
[*Diagnostics tests* R^2 = .84; serial cor. $\chi^2(1)$ = 1.59(.21); hetero. $\chi^2(1)$ = .51(.48)]						
Panel III Eq. 1	K	−.55	−1.57 (.126)	ΔK	.85	3.93 (.000)
Sample 1961–2009	L	1.30	3.46 (.002)	$\Delta L(-3)$	−.65	−3.21 (.003)
Dep. variable Y	EL	.53	2.48 (.019)	ΔEL	.48	3.78 (.001)
AIC-ARDL (1,4,4,1)	Constant	−3.10	−.80 (.433)	$ecm(-1)$	−.30	−3.25 (.003)
	Trend	.01	2.37 (.024)			
[*Diagnostics tests* R^2 = .84; serial cor. $\chi^2(1)$ = 2.73(.10); hetero. $\chi^2(1)$ = 2.16(.14)]						
Panel IV Eq. 3	Y	.92	6.95 (.000)	ΔY	.67	5.79 (.000)
Sample 1978–2009	P_{EL}	.05	1.22 (.235)	ΔP_{EL}	−.02	−1.76 (.090)
Dep. variable EL	Constant	−6.16	−3.16 (.004)	$ecm(-1)$	−.13	−2.68 (.012)
SBC-ARDL (1,1,1)						
[*Diagnostics tests* R^2 = .72; serial cor. $\chi^2(1)$ = .01(.94); hetero. $\chi^2(1)$ = 1.87(.17)]						

R^2 values refer to the error correction model, while the other diagnostic test statistics refer to the underlying ARDL equation

The empirical evidence presented herein indicates the data on GDP, energy consumption and other variables in the model experienced structural break(s) over the last five decades caused by both internal and external economic shocks. Accordingly, the results on the cointegration relationship could be affected by these break(s) in data. It is, therefore, crucial to incorporate the information on structural break(s) in the subsequent modelling and inferences. Moreover, neither the production side nor the demand side framework alone is sufficient to draw a fruitful conclusion. When alternative model specifications and sample periods based on the evidences of structural break(s) are explored, strong evidence of cointegration is found between GDP and energy consumption in Australia at both aggregate and disaggregate levels. The empirical results also indicate the existence of bidirectional link between GDP and aggregate energy and between GDP and major energy categories in Australia.

Our empirical evidence indicates that the cointegration results between energy, GDP and other relevant variables are more robust when controlling for a structural break in the early 1970s. Note that the Australian economy experienced an economic downturn in the early 1970s, which could be a possible reason for the differences in results. The cointegration between energy and output found in the present study is consistent with the findings of Fatai et al. (2004), Narayan and Smyth (2005b) and Shahiduzzaman and Alam (2012). The estimated results indicate that the aggregate energy use drives GDP, capital and labour in both long- and short-run. With respect to the energy demand models, our empirical evidences reinforce the dominant role of GDP in the long-run energy demand in Australia.

The empirical evidence found in this paper poses significant implications in terms of energy-environmental policies in Australia. Australia experiences relatively high level of energy and CO_2 emissions intensity as compared to other competing economies. While the country is currently implementing various energy conservation policies to reduce emissions, the estimation results as found in this study provide a useful guide for intervention strategies. The evidence of a bidirectional relationship between GDP and both aggregate and disaggregate levels of energy consumption indicates that any external shock to one of these will readily be transmitted to the other and that the process will continue through the feedback process. Direct measures in reducing energy consumption would, therefore, negatively impact the economic growth and energy consumption per se; however, alternative options such as improvement of energy efficiency and technological changes would be beneficial to economic growth and environmental outcomes. The low values of energy price elasticities indicate that a large change in relative price would be required to encourage energy efficiency measures and/or to shift towards low-emitting technologies.

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