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On the causal dynamics between hydroelectricity consumption and economic growth in Latin America countries

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ABSTRACT

The aim of this study is to examine the relationship between hydroelectricity consumption and economic growth in seven Latin America countries including Argentina, Brazil, Chile, Colombia, Ecuador, Peru and Venezuela. The analysis is conducted within a neoclassical model involving capital and labour force, for the period of 1970–2012. We utilize a recently developed cointegration test to investigate the long run relationship in the variables. Having established that the variables are cointegrated, the results indicate long run bidirectional causality between hydroelectricity consumption and economic growth in Argentina and Venezuela. There is evidence for long run unidirectional causality from hydroelectricity consumption to economic growth in Brazil, Chile, Colombia, Ecuador and Peru. The long run coefficients from the regression analyses suggest that hydroelectricity consumption positively affect the economies of these Latin American countries. However, limited evidence of causality between the two variables is found in the short run. The ensuing policy implications of the findings are discussed.

Contents

1. Introduction

The increasing alarm over greenhouse gas emissions and unstable energy prices associated with the consumption of fossil fuels (such as oil and natural gas) have in part led to the rising importance of renewable energy. Unlike the fossil fuel usage, renewable energy usage has limited negative impact on its surrounding environment. Therefore, investing in renewable energy will significantly reduce the discharge of harmful or unwanted substance (commonly associated with fossil fuels) into the water, air, and soil. Furthermore, the adoption of renewable energy also reduces a country's vulnerability to supply disruption and price volatility that are found in the

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usage of fossil fuels. As such, renewable energy facilities can lead to the protection of the environment and security of energy supply.

There are numerous studies on the viability of renewable energy in an economy including the causal relationship between aggregate renewable energy and economic growth [1–3]. The composition of renewable energy such as hydroelectricity, solar, wind, biofuels is different in terms of their availability, flexibility, level of development, and return on investment. Considering aggregate renewable energy without adequately evaluating the different characteristics of its components could conceal the varying effects of the various types of renewable energy consumption, which could lead to wrong policy implications for each component, especially for hydroelectricity which is characteristically dissimilar from other types of renewable energy. For instance, hydropower is a more reliable source of energy than solar power because it provides steady supply of power. On the other hand, solar power is an intermittent energy source with periods of interruption in energy supply, especially during the overcast days. Unlike the other forms of renewable energies, hydroelectricity can further offer benefits beyond supplying energy such as irrigation, water supply, recreation and aquaculture [4]. Resulting from these differences, the impact of different sources of energy on economic growth is not the same, and different directions of causation exist between gross domestic product (GDP) and various kinds of energy consumption. Therefore, policy implications inferred from a causal examination of renewable energy and economic growth may only be applicable to some, but not all the components of renewable energy. Detailed causal analyses between different types of renewable energy consumption, especially the hydroelectricity consumption- and economic growth are limited in the literature.

The main purpose of this study is to examine the causal relationship between hydroelectricity consumption and economic growth in seven Latin American countries (Argentina, Brazil, Chile, Colombia, Ecuador, Peru and Venezuela) for the period, 1970–2012. One of the rationales for choosing these countries is that no regional bloc in the world generates as much energy from hydropower as the Latin American region. While the share of hydroelectricity consumption in total energy consumed in the world is estimated to be 16%, the shares of hydroelectricity consumption in total energy consumed in Brazil, Ecuador and Venezuela are 35%, 19%, 25% respectively [5–8]. Although Latin American countries generate so much hydropower, there is still a high potential for hydropower resources in the subregion. Hydroelectric plants provided for around 80%, 68%, 50% and 70% of the total electricity generated in Brazil, Colombia, Ecuador and Venezuela [5–8]. The unutilized hydropower resources are very vast and renewable energy is one of the region's most important assets. Brazil currently uses only about 30% of its hydropower potentials, and the remainder is located mostly in the Amazon region [9]. With a potential of 113 GB of hydropower resources, Colombia is second only to Brazil in South America [10]. In Chile, the estimated potential is around 9000 MW [11].

We contribute to the existing literature in the following four ways: firstly, we analyze the casual relationship between hydroelectricity and economic growth, which to the best of our knowledge has not been adequately explored by the extant literature. Secondly, we provide for structural breaks in the unit root, cointegration and causality analyses. Many of the past papers on the relationship between renewable energy consumption and economic growth have either ignored the presence of structural breaks altogether in their analyses or have only provided for breaks in either unit root testing or cointegration analysis [12]. However, ignoring the possibility of structural breaks in the estimation process may weaken the power of rejecting a false null hypothesis [13]. In particular, our choice countries have experienced several structural breaks over the years such as the Latin American debt fiasco of the early 1980s and the financial crisis of the late 1990s and the early 2000s.

Thirdly, we conduct the unit root testing, cointegration testing and causality analysis within the time series framework. Largely due to the fact that the sample periods are finite (less than 30 years), several existing multi-country studies on the link between renewable energy and economic growth have utilized panel techniques, especially panel causality tests. Although panel-based methods are less susceptible to the problems associated with short span of data; thus generating more degrees of freedom, more variability, and efficient estimates; time series methods on the other hand have an advantage over the other methods because they capture the individual characteristics of each country better. As such, the time series approach would provide more information and better policy guide for individual countries in the sample. Furthermore, utilizing a dataset with duration of 43 years, we reduce the problems of small sample size that is typically associated with time series techniques. Similarly, the relatively long period of coverage tends to cover more events and thus more informative than short sample size. Fourthly, we analyze the four hypotheses on the energy consumption-growth nexus in relation to their long run elasticities.

The remainder of the paper is organized as follows. Section 2 deals with the literature review of the multi-country studies on renewable consumption and economic growth. Section 3 illustrates the data and methodology adopted in this paper and Section 4 provides the empirical findings. Section 5 deals with discussion, while the last section involves the conclusion of the study.

2. Literature review

The causal relationship between renewable energy consumption and the economy is one of the areas that have received limited attention in the energy consumption–economic growth literature. Instead, most of the literature has primarily focused on either aggregate energy consumption or electricity consumption with very vital policy implications [2]. For instance, unidirectional causality flowing from economic growth to energy consumption suggests that the economy is less energy-dependent and conserving energy use is a vital policy option because energy conservation policies will not harm economic development. Similarly, the causality running from energy consumption (with or without feedback) to economic growth implies that energy consumption plays a vital role in the economic development process. Therefore, any attempt to limit energy consumption may impede economic growth and encouragement of energy use will promote economic growth. The nonexistence of causality between energy consumption and gross domestic products (GDP) is an indication that any initiative in the energy sector will have no impact on the output.

In this section, we review some multi-country papers on renewable energy and economic growth. The literature review is divided into two parts, with the first part concentrating on papers wherein the individual country's estimates are not provided, while the second part involves papers wherein the country specific estimates are provided. Sadorsky [14] examined the relationship between renewable energy and real GDP per capita, while providing for electricity prices for the period, 1994–2003. The author adopted the cointegration test of Pedroni [15] and the conventional error correction model in the estimation process. Using a sample that comprises Argentina, Brazil, Chile, China, Colombia, Czech Republic, Hungary, India, Indonesia, South Korea, Mexico, Peru, Philippines, Poland, Portugal, Russia, Thailand and Turkey: the results show the existence of bilateral causality between real income and real renewable energy consumption in the long run.

Apergis and Payne [16] examined the nexus for a panel of six Central American countries over the period of 1980–2006. The authors provided for labour force and capital in a multivariate setting.

The test statistics support the existence of long run equilibrium relationship in the series and further suggest bilateral long run causality between renewable energy consumption and economic growth. The fully modified ordinary least square (FMOLS) coefficients suggest that the use of renewable energy has positive effect on economic growth.

Apergis and Payne [17] explored the causal relationship between renewable energy consumption and economic growth for 80 countries in a multivariate model that includes real gross fixed capital formation and the labour force. A long run relationship between the variables was established in the variables. The findings further reveal long run bidirectional causality between renewable energy consumption and economic growth. Ben Aïssa et al. [18] used the data of 11 African countries to examine the relationship between renewable energy and economic growth in 1980–2008, while separately providing for imports and exports. The results reveal evidence in favour of unidirectional causality from renewable energy to economic growth. Apergis and Danuletiu [19] applied the Canning and Pedroni [20] causality test to conduct bivariate causal analysis between renewable energy consumption and economic growth in 80 countries for the period of 1990–2012. The findings provided evidence for bidirectional causality between renewable energy consumption and economic growth.

Pao et al. [21] examined the causal relation between renewable energy consumption (among other forms of energy) and economic growth in Mexico, Indonesia, South Korea, and Turkey for the period of 1990 to 2010. Using the Westerlund [22] cointegration test and the conventional panel causality test, the authors provide evidence for a unidirectional long-run causality from renewable energy consumption to economic growth with positive bidirectional short-run causality. Similarly, Salim et al. [23] explore the dynamic relationship between renewable consumption, non-renewable energy consumption, industrial output and GDP growth in 29 OECD countries for the period of 1980–2011. The study adopts the Westerlund [22] cointegration test and the conventional panel causality test in the estimation process. The results support the existence of cointegration in the variables. The panel causality analyses show bidirectional causality between renewable energy consumption and industrial output in the long and short run. The same result was observed for the relationship between GDP and renewable energy consumption.

One issue with the foregoing papers is that the results generated at a panel level may not be necessarily valid across the sample. Although the panel estimates may support a particular hypothesis, it does not necessarily imply that all the countries in the sample size will yield similar conclusions. The second part of the literature involves papers wherein country's specific estimates are provided. Tugcu et al. [12] examined the causal relationship between renewable energy consumption and economic growth in G-7 countries (Canada, France, Germany, Italy, Japan, the United Kingdom, and the United States) for the period of 1980–2009. In a specification that includes physical capital, labour, human capital, research and development, the autoregressive distributed lag (ARDL) method of Pesaran et al. [24] and the causality test of Hatemi-J [25] were applied. In the equation involving renewable energy consumption, the results unveil the nonexistence of causal relationship between renewable energy consumption and economic growth in Canada, France, Italy and the USA. There is bilateral causality between renewable energy consumption and economic growth in Japan and England, while the conservation hypothesis is supported for Germany.

Salim and Rafiq [26] examined the relationship between renewable energy consumption and income, while providing for oil price and employment in Brazil, China, India, Indonesia, Philippines and Turkey. The study shows bidirectional causality in Brazil, China, Philippines and Turkey and unidirectional causality from income to renewable energy in India and Indonesia. Bildirici and Özaksoy [27] used the data of 10 European countries-Austria, Finland, France, Hungary, Poland, Portugal, Romania, Spain, Sweden, and Turkey to examine the relationship between biomass consumption and economic growth. Using the ARDL cointegration test and VECM Granger causality test, the results indicate unidirectional causality from biomass energy consumption to economic growth for Hungary and Poland. Unidirectional causality from economic growth to biomass energy consumption is documented for Austria and Turkey, while bidirectional causality between economic growth and biomass energy consumption is noted in the case of Spain, Sweden, and France. The results further unveil the absence of long run causality in Finland, Portugal and Romania. Bildirici [28] examined the causal relationship between biomass energy consumption and economic growth in 10 developing and emerging countries – Argentina, Bolivia, Cuba, Costa Rica, El Salvador, Jamaica, Nicaragua, Panama, Paraguay and Peru. Using the ARDL and Granger causality test, the estimates show unilateral causality from biomass energy consumption to GDP in Argentina, Bolivia, Cuba, Costa Rica, Jamaica, Nicaragua, Panama and Peru. While there is bidirectional causality from biomass energy consumption to GDP for El Salvador, no causality is found in the case of Paraguay.

Sebri and Ben-Salha [29] utilize the data of Brazil, Russia, India China and South Africa to explore the causal relationship between economic growth and renewable energy consumption for the period of 1971–2010. Using a multivariate framework that includes carbon dioxide emissions and trade openness, the authors adopt the autoregressive distributed lag (ARDL) method of Pesaran et al. [24] and vector error correction model (VECM). The results support the bidirectional causality between renewable and economic growth in the countries.

Omri et al. [30] investigate the causal relationship between renewable energy and economic growth in 17 developed and developing countries for the period of 1990–2011. The test statistics provide support for unidirectional causality running from renewable consumption to economic growth in Hungary, India, Japan, Netherlands, and Sweden, while unidirectional causality running from economic growth to renewable consumption exists in Argentina, Spain, and Switzerland. Moreover, the authors established bidirectional relationship in Belgium, Bulgaria, Canada, France, Pakistan, and the USA, while no causality exists in Brazil, Finland, and Switzerland.

The foregoing review reveals that countries tend to follow different causality patterns, when individual country' estimate is estimated. Therefore, it is better to provide for each country's long run and short run elasticities in the estimation. Moreover, it is observed that the studies on hydroelectricity consumption–economic growth nexus are very scarce in the literature. Furthermore, the above review indicates that structural breaks have not been adequately treated in the analysis, especially in the causality tests, whereas the policy implications are usually generated from the causality analysis.

3. Methodology

3.1. Model and data

Consistent with the existing energy papers such as Lean and Smyth [31], we examine the causal relationship between hydroelectricity consumption and economic growth within the neoclassical model that includes capital and labour force. The general functional form of the model is given as follows:

$$\ln Y_t = f(H_t, K_t, L_t) \quad (1)$$

Here Y is real gross domestic product, H is hydroelectricity consumption (tonnes of oil equivalent), K is the gross fixed capital formation (which represents capital), and L is the labour force. The data for the real gross domestic product and gross fixed capital formation were obtained from *United Nations Database*, while data for hydroelectricity consumption were extracted from the *BP Statistical Review of World Energy*. The data for labour force is sourced from the *Total Economy Database* of the Conference Board for the period of 1970–2012. We transform all the variables into logarithmic form, which produces better result compared to the linear functional form. The empirical equation of the model is given as follows:

$$\ln Y_t = a_1 + a_2 \ln H_t + a_3 \ln K_t + a_4 \ln L_t + u_t \quad (2)$$

where $\ln Y_t$ is natural log of real GDP per capita, $\ln H_t$ is natural log of hydroelectricity consumption, $\ln K_t$ is natural log of gross fixed capital formation, $\ln L_t$ is natural log of labour force and u_t is error term with the assumption of normal distribution. We take a closer look at the hydroelectricity consumption in Table 1, where the descriptive statistics are reported. Brazil is the largest hydroelectricity consumer with an average usage of 49.775 million of tonnes equivalent (mtoe) per annum. The other large users in the region include Argentina, Colombia and Venezuela. The standard

Table 1
Descriptive statistics

Country	Mean	Std. dev.	Skewness	Kurtosis	Jarque–Bera
Argentina	5.188	3.045	−0.070	1.730	2.924 (0.232)
Brazil	49.775	25.411	0.088	1.982	1.912 (0.384)
Chile	3.209	1.614	0.268	1.769	3.227 (0.199)
Colombia	5.774	2.829	0.012	1.806	2.556 (0.279)
Ecuador	1.104	0.743	0.200	2.196	1.446 (0.485)
Peru	2.708	1.240	0.265	1.806	3.058 (0.217)
Venezuela	9.461	6.214	0.160	1.643	3.483 (0.175)

The probability values are in parenthesis

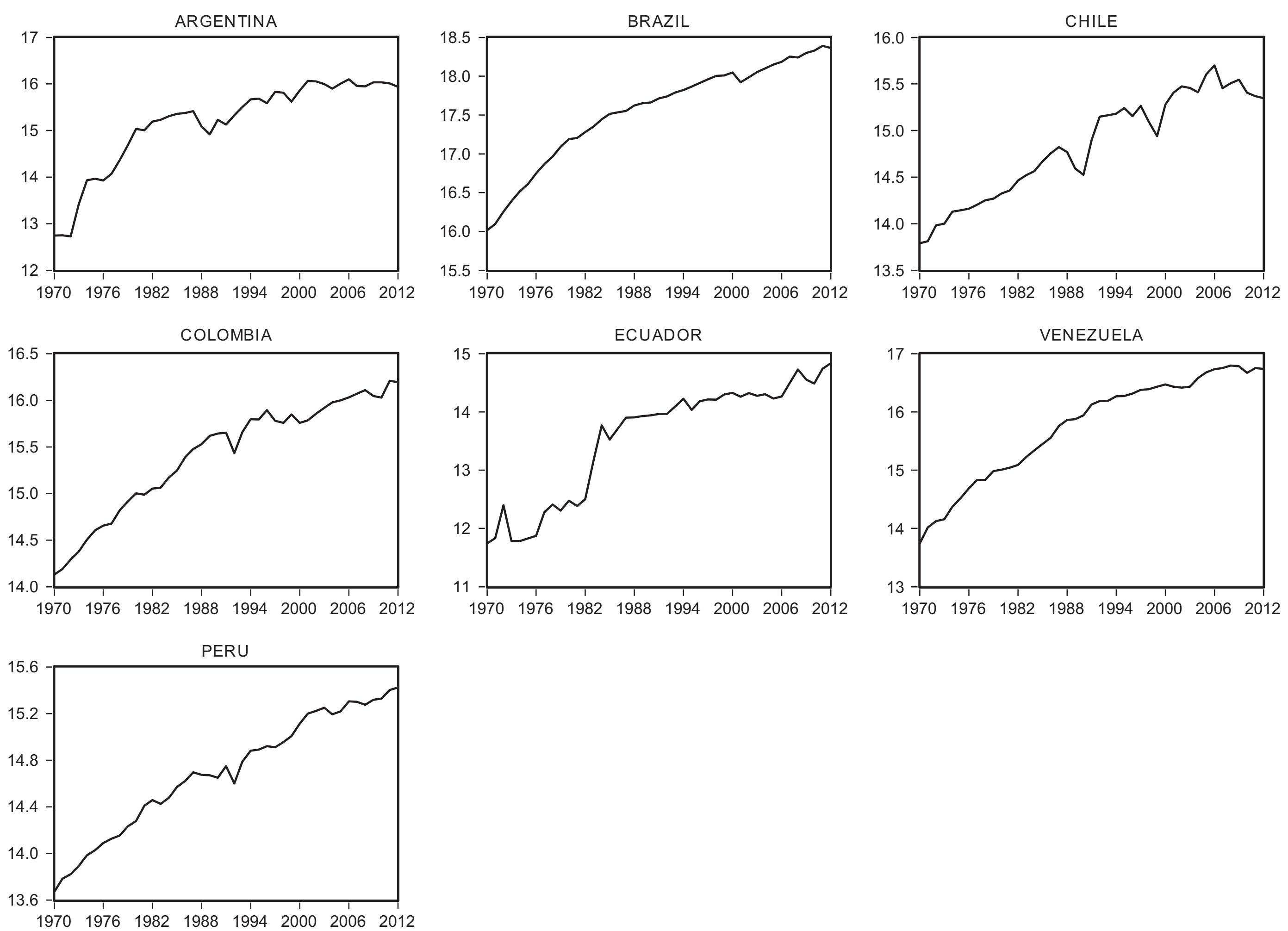

Fig. 1. Hydroelectricity consumption, 1970–2012.

deviation shows that the largest fluctuation in hydroelectricity usage is experienced in Brazil, which coincidentally is the top consumer of hydroelectricity. The normality test, as represented by the Skewness, Kurtosis and Jarque–Bera statistics indicates that the series are normally distributed. Fig. 1 shows that the hydroelectricity consumption is increasing in most of these countries.

Beyond the fact that the control variables (capital stock and labour force) are integral part of the neoclassical model, they are important in an energy-output model. For instance, capital stock is a significant factor in the production function. Moreover, the inclusion of more variables, such as capital in addition to hydroelectricity incorporates more information that affects aggregate output than in the bivariate case, in which hydroelectricity alone is regarded as the singular input of production.

In the developing countries, capital and energy complement each other. Private investment often enables hydroelectricity projects to be built in a shorter time frame. There is rising call for renewable energy use in developing countries, not only from local investors, but also from international investors [32]. Investment in hydropower projects is increasingly becoming more global with investors exploring new regions [10]. The past literature has controlled for capital formation in energy-growth equations [31]. Labour force is required in the production of several goods and services and it explains a substantial part of the total output in an economy. Besides, it is required in the construction of many hydropower plants especially at the beginning of the project cycle. Hydroelectric projects are complex construction projects at the best of times, with great demands for skilled labour. The past literature has controlled for labour force in energy-growth regressions [31].

3.2. Unit root tests

There is the need to conduct unit root tests on the variables before the estimation of the long run properties of the variables. The traditional unit root tests such as Dickey–Fuller are weak, if there are structural shifts in a variable. To tackle this issue, stationarity tests which allow for shifts have been suggested in the time series literature. In this paper, we adopt the Lee and Strazicich [33,34] tests, which are expressed as below:

$$\Delta S_t = \delta' \Delta Z_t + \phi \bar{S}_{t-1} + \sum_{i=1}^{p} \gamma \Delta \bar{S}_{t-i} + \mu_t \tag{3}$$

The null hypothesis of nonstationarity is valid when $\phi = 0$ and the alternative hypothesis is valid if $\phi < 0$. The structural shift is represented in the method by Z_t, which contains the exogenous variables. In a model that provides for a shift in both level and trend, $Z_t = [1, t, D_{1t}, DT_{1t}]'$ where $DT_{jt} = t$ if $t \geq T_B + 1$, and 0, otherwise. D_{1t} and DT_{1t} are the dummy variables that represent the period, when structural shift break is experienced in the level and trend respectively. To endogenously estimate the location of $\lambda_j = T_{Bj}/T, j = 1.$, the "minimum LM test" in Lee and Strazicich [34] is used. In a model that provides for two changes in both level and trend, $Z_t = [1, t, D_{1t}, D_{2t}, DT_{1t}, DT_{2t}]'$ where $DT_{jt} = t$ if $t \geq T_{Bj} + 1, j = 1, 2$, and 0, otherwise. In order to make sure that the issue of serial correlations is avoided, the augmented terms of $\Delta \bar{S}_{t-i}$ are introduced into the model.

3.3. Cointegration test

Although there is a large volume of papers that have addressed the issue of structural breaks in univariate root tests, few methods that provide for structural breaks in multivariate co-integration analysis are available [35,36]. One of the pioneering studies was Inoue [37] and thereafter Saikkonen and Lutkepohl [38], which proposed a two-stage method to compute co-integrated VAR functions with shift. A setback associated with the method of Saikkonen and Lutkepohl [38] is that it fails to incorporate all the necessary restrictions on the deterministic series in the first stage. Lutkepohl et al. [39] introduced a procedure, wherein the parameters of the deterministic series are computed through the Generalised Least Squares (GLS) approach in the first stage, and the Johansen [40] co-integration test is subsequently applied in the second stage. On the other hand, the Johansen et al. [41] tests incorporate the deterministic components into the co-integration VAR and also provide for several structural shifts. We apply the Johansen et al. [41] co-integration method to investigate the long run relationship in hydroelectricity consumption, capital stock, labour force and economic growth. This is plausible by extending the popular likelihood-based co-integration analysis as analyzed in Johansen [40,42]. Similar to the Johansen [42] test, the specification is conditionally based on the initial k of each subsample, $Z_{T(j-1)+1, \ldots\ldots} Z_{T(j-1)+k}$ for $j = 1, \ldots\ldots q$ and $T_{j-1+k} < t \leq T_j$ the VAR formulation representation is transformed into VECM framework. The multiple structural changes are presented as follows:

$$\Delta Z_t = \left(\prod, \prod_j\right) \begin{pmatrix} z_{t-1} \\ t \end{pmatrix} + u_j + \sum_{i=1}^{k-1} \Gamma_i \Delta z_{t-i} + \varepsilon_t$$

The expression is similar to Johansen [42] with the exception of the $(n \times 1)$ vectors $\prod_j$ and u_j that are associated with the deterministic component, which differ in each subsample for $j = 1, \ldots\ldots q$. Z_t is the relevant vector for the co-integration analysis. Γ and Π are associated with the stochastic elements of the time series, and they are similar in all subsamples. k is selected, so ε_t is assumed to fulfil the classical assumptions. There is long run relationship, if the coefficient matrix Π has a reduced rank such that $\Pi = \alpha\beta$, where α denotes adjustment speed and β contains the long run parameters [41,43]. Thus, $\prod_j = \alpha yj$ whereby γj denotes the coefficients of the long run trend in every subsample. Consistent with the Johansen et al. [41], this formulation is denoted as H_l.[2] The H_l method is suitable if the variables display some forms of linear trend across the periods. Coincidentally, all the series under investigation in this study (hydroelectricity consumption, capital stock, labour force and economic growth) follow such pattern (Fig. 1)[3]. Hence, this study focus on the H_l model (wherein only two breaks or three periods) are considered. The model is represented below[4]:

$$\Delta Z_t = \alpha \begin{pmatrix} \beta \\ \gamma_1 \\ \gamma_2 \\ \gamma_3 \end{pmatrix}' \begin{pmatrix} Z_{t-1} \\ tE_{1,t} \\ tE_{2,t} \\ tE_{3,t} \end{pmatrix} + \mu_1 E_{1,t} + \mu_2 E_{2,t} + \mu_3 E_{3,t}$$

$$+ \sum_{i=1}^{k-1} \Gamma_i \Delta z_{t-i} + \sum_{i=1}^{k} \sum_{j=3}^{q} \kappa_{j,i} D_{j,t-1} + \varepsilon_t \tag{5}$$

where the length of each period is of length Tj such that $T_0 < T_j < T_q$. T_{j+1} is the first element in $(j+1)$th; t represents the trend; $E_{j,t}$ denotes dummy variable, which is 1 if the observation i

[2] Johansen et al. [41] also introduce two other models. The first model involves a break in intercept in which the intercepts in the cointegrating relationship can be different across sub-samples. The second model allows some or all of the time series to follow a trending pattern in every sub-sample and the cointegrating links are stationary in each sub-sample (with possibly a broken constant level). However, Johansen et al. [41] argued and demonstrated that such specification is heavily burdened with nuisance parameters and as such, it is less attractive.

[3] The graphs of the other series are not reported here because of space concern.

[4] We focus on two-structural breaks as against one-structural break model because evidence suggests that the region experienced more than a break during the sample period.

Table 2
Conventional unit root tests

Countries	Variables	ADF unit root test	PP unit root test	DF-GLS unit root test
Argentina		*T*-statistic	Prob. value	*T*-statistic
	$\ln Y_t$	−1.490(1)	−1.014(1)	−1.801(1)
	$\Delta \ln Y_t$	−4.953***(0)	−4.875***(3)	−4.976***(0)
	$\ln H_t$	−1.940(4)	−1.932(4)	−1.262(0)
	$\Delta \ln H_t$	−5.730***(0)	−5.732***(4)	−5.577***(0)
	$\ln K_t$	−2.298(1)	−1.748(1)	−2.412(1)
	$\Delta \ln K_t$	−4.918***(0)	−4.760***(4)	−4.939***(0)
	$\ln L_t$	−1.806(1)	−1.289(3)	−1.938(1)
	$\Delta \ln L_t$	−4.481***(0)	4.365***(3)	−4.598***(0)
Brazil	$\ln Y_t$	−3.134(3)	−1.558(4)	−1.774(0)
	$\Delta \ln Y_t$	−4.546***(0)	−4.481***(1)	−4.363***(0)
	$\ln H_t$	−1.998(0)	−1.932(4)	−1.263(0)
	$\Delta \ln H_t$	−5.729***(0)	−5.732***(0)	−5.577***(0)
	$\ln K_t$	−2.669(0)	−2.851(2)	−1.817(0)
	$\Delta \ln K_t$	−4.518***(0)	−4.453***(3)	−4.399***(0)
	$\ln L_t$	−3.173(1)	−3.185(2)	−1.611(1)
	$\Delta \ln L_t$	−5.327***(0)	−5.280***(4)	−5.375***(0)
Chile	$\ln Y_t$	−2.992(3)	−2.360(1)	−2.361(1)
	$\Delta \ln Y_t$	−4.794***(0)	−4.824***(2)	−4.587***(0)
	$\ln H_t$	−2.125(0)	−2.130(3)	−2.272(0)
	$\Delta \ln H_t$	−6.096***(1)	−6.064***(4)	−6.220***(1)
	$\ln K_t$	−3.000(0)	−3.004(2)	−1.994(0)
	$\Delta \ln K_t$	−5.425***(0)	−5.369***(2)	−5.563***(0)
	$\ln L_t$	−2.078(0)	−2.318(2)	−2.049(0)
	$\Delta \ln L_t$	−6.046***(0)	−6.045***(1)	−6.198***(0)
Colombia	$\ln Y_t$	−3.079(1)	−2.693(2)	−2.280(1)
	$\Delta \ln Y_t$	−3.834**(0)	−3.833**(1)	−3.883***(0)
	$\ln H_t$	−1.967(0)	−1.793(4)	−1.482(0)
	$\Delta \ln H_t$	−7.331***(0)	−7.932***(4)	−7.474***(0)
	$\ln K_t$	−2.436(1)	−1.957(4)	−2.572(1)
	$\Delta \ln K_t$	−4.366***(0)	−4.294***(4)	−4.468***(0)
	$\ln L_t$	−2.078(0)	−2.318(2)	−2.049(0)
	$\Delta \ln L_t$	−6.046***(0)	−6.045***(1)	−6.198***(0)
Ecuador	$\ln Y_t$	−2.036(4)	−3.124(1)	−2.296(4)
	$\Delta \ln Y_t$	−5.080***(0)	−5.141***(3)	−5.196***(0)
	$\ln H_t$	−1.886(0)	−1.870(1)	−1.900(0)
	$\Delta \ln H_t$	−5.893***(1)	−7.078***(4)	−7.016***(0)
	$\ln K_t$	−1.602(0)	−1.762(2)	−1.559(0)
	$\Delta \ln K_t$	−6.114***(0)	−6.122***(1)	−5.486***(0)
	$\ln L_t$	−0.560(0)	−0.663(3)	−0.725(0)
	$\Delta \ln L_t$	−6.629***(0)	−6.625***(2)	−6.791***(0)
Peru	$\ln Y_t$	−1.179(1)	−0.647(0)	−1.551(1)
	$\Delta \ln Y_t$	−4.315***(0)	−4.120**(4)	−4.367***(0)
	$\ln H_t$	−2.851(0)	−2.745(3)	−2.171(0)
	$\Delta \ln H_t$	−7.982***(0)	−8.815***(4)	−7.797***(0)
	$\ln K_t$	−1.783(1)	−1.232(1)	−2.065(1)
	$\Delta \ln K_t$	−4.337***(0)	−4.122(4)	−4.349***(0)
	$\ln L_t$	−0.60(0)	−0.860(0)	−1.153(0)
	$\Delta \ln L_t$	−6.441***(0)	−6.469***(3)	−6.527***(0)
Venezuela	$\ln Y_t$	−2.150(0)	−2.150(0)	−2.751(1)
	$\Delta \ln Y_t$	−4.853***(0)	−4.737***(4)	−4.977
	$\ln H_t$	−1.389(0)	−1.395(1)	−0.302(1)
	$\Delta \ln H_t$	−6.533***(0)	−6.547***(2)	−5.423***(0)
	$\ln K_t$	−1.701(2)	−2.120(4)	−2.649(1)
	$\Delta \ln K_t$	−5.407***(1)	−4.382***(4)	−5.443***(1)
	$\ln L_t$	−2.492(1)	−2.059(0)	−1.984(1)
	$\Delta \ln L_t$	−4.059**(0)	−3.924**(4)	−4.144***(0)

With maximum lag set at 4, the optimal lags in ADF, DF-GLS are selected based on Akaike information criterion, whereas the Bartlett with Newey–West bandwidth is used for PP. For uniformity sake, the regressions in each test include a constant and trend.

** Indicate significance at 5% levels respectively.

*** Indicate significance at 1%.

is of *j*th period; *k* is the maximum lag length in the VAR equation; $D_{j,t-i}$ is a dummy for the *i*th observation in *j*th period.

Computing the rank *r* in the H_l model differs from the standard statistics in Johansen [40] as the asymptotic critical values rely on three factors. The first is the volume of shifts that are incorporated into the analysis. The second factor is the location of the shifts. The third factor is the amount of series under investigation minus the co-integrating rank, *r*. The trace critical values for Johansen et al. [41] are provided in Giles and Godwin [44]. Upon the confirmation of the existence of cointegration in the variables, the causality test can be conducted under the assumption of at least a long run relationship in the series.

3.4. Granger causality test

The pattern of hydroelectricity consumption, capital, labour force and economic growth can be investigated by using the VECM Granger causality method. Granger [45] posited that the VECM is more suitable to investigate the causality between series that are integrated at I(1). The model is based on the assumption that all the variables are not exogenous and also premised on the fact that the dependent variable is explained by the past values of the independent variables and the past values of the dependent variable. The VECM equations are specified as below:

$$\Delta \ln Y_t = \alpha_{10} + \phi_{11} D_1 + \phi_{12} D_2 + \sum_{i=1}^{k} \beta_{13} \Delta \ln Y_{t-i} + \sum_{i=1}^{k} \delta_{14} \Delta \ln H_{t-i} + \sum_{i=1}^{k} \kappa_{15} \Delta \ln K_{t-i} + \sum_{i=1}^{k} \gamma_{16} \Delta \ln L_{t-i} + \xi_{17} ECT_{t-1} + \mu_t \quad (7)$$

$$\Delta \ln H_t = \alpha_{20} + \phi_{21} D_1 + \phi_{22} D_2 + \sum_{i=1}^{k} \beta_{23} \Delta \ln Y_{t-i} + \sum_{i=1}^{k} \delta_{24} \Delta \ln H_{t-i} + \sum_{i=1}^{k} \kappa_{25} \Delta \ln K_{t-i} + \sum_{i=1}^{k} \gamma_{26} \Delta \ln L_{t-i} + \xi_{27} ECT_{t-1} + \mu_t \quad (8)$$

$$\Delta \ln K_t = \alpha_{30} + \phi_{31} D_1 + \phi_{32} D_2 + \sum_{i=1}^{k} \beta_{33} \Delta \ln Y_{t-i} + \sum_{i=1}^{k} \delta_{34} \Delta \ln H_{t-i} + \sum_{i=1}^{k} \kappa_{35} \Delta \ln K_{t-i} + \sum_{i=1}^{k} \gamma_{36} \Delta \ln L_{t-i} + \xi_{37} ECT_{t-1} + \mu_t \quad (9)$$

$$\Delta \ln L_t = \alpha_{40} + \phi_{41} D_1 + \phi_{42} D_2 + \sum_{i=1}^{k} \beta_{43} \Delta \ln Y_{t-i} + \sum_{i=1}^{k} \delta_{44} \Delta \ln H_{t-i} + \sum_{i=1}^{k} \kappa_{45} \Delta \ln K_{t-i} + \sum_{i=1}^{k} \gamma_{46} \Delta \ln L_{t-i} + \xi_{47} ECT_{t-1} + \mu_t \quad (10)$$

where α is the constant term and ϕ is the coefficient associated with the dummy variables included due to the break. In the model involving no break, ϕ is set to zero. β, δ, κ and γ are the coefficients of the lagged regressors. These parameters represent the short term impacts of the explanatory series on the dependent series. The *F*-test of joint significance of these lagged terms constitutes the short-run Granger causality. For instance, if all the coefficients δ in Eq. (7) are jointly significant, then there is causality flowing from hydroelectricity consumption to Granger cause economic growth in the short run.ξ is the symbol associated with ECT and captures the adjustment speed towards the long run equilibrium. The *t*-test for that coefficient represents the long-run Granger causality result with respect to ECT. A significant ECT coefficient means that past equilibrium errors are important determinants of the current outcomes. For instance, if the coefficient ξ is significant in Eq. (7), then hydroelectricity consumption, capital, and labour force Granger cause economic growth in the long term. The same analogy is applicable to the remaining equations in the system.

4. Results

The empirical analyses commence with the testing for the unit root properties of the series. As a benchmark, we first apply the

Table 3
LM unit root test

Country	Variable	*T*-statistics	TB1	TB2	DU1	DT1	DU2	DT2
Argentina	$\ln Y_t$	−4.450[4]	1989	2000	−0.128*** (−2.532)	0.026 (1.437)	−0.017 (−0.323)	−0.130*** (−3.542)
	$\Delta \ln Y_t$	−5.967[4]	1990	2001	0.144** (2.991)	0.002 (0.102)	−0.233*** (−4.415)	0.209*** (5.883)
	$\ln H_t$	−5.104[1]	1982	1991	0.015 (0.120)	−0.215*** (−3.742)	−0.006 (−0.050)	0.203*** (3.563)
	$\Delta \ln H_t$	−7.311***[1]	1982	1991	−0.267** (−1.945)	−0.135* (−1.941)	0.593*** (4.052)	−0.054 (−1.000)
	$\ln K_t$	−4.172[3]	1989	1999	−0.362** (−2.352)	0.133* (2.295)	0.079 ()0.517	−0.314*** (−3.458)
	$\Delta \ln K_t$	−6.378 [4]	1990	2001	0.582*** (4.012)	−0.073 (−1.325)	−0.833*** (−5.110)	0.682*** (6.255)
	$\ln L_t$	−3.966[0]	1993	2004	0.007 (0.355)	−0.016** (−1.989)	0.015 (0.843)	0.037*** (3.321)
	$\Delta \ln L_t$	−5.467[0]	1990	2001	0.005 (0.262)	−0.017** (−2.309)	0.033* (1.783)	0.001 (0.127)
Brazil	$\ln Y_t$	−3.388[0]	1978	2002	0.039 (1.312)	−0.084*** (−6.134)	−0.037 (−1.200)	0.022* (1.871)
	$\Delta \ln Y_t$	−7.419***[4]	1982	1991	−0.154*** (−4.227)	0.088*** (5.298)	−0.077*** (−2.582)	−0.033*** (−2.859)
	$\ln H_t$	−3.797[4]	1984	2000	0.107** (3.502)	−0.120*** (−5.623)	0.195*** (−7.563)	0.038*** (2.896)
	$\Delta \ln H_t$	−6.698[0]	1983	2002	0.085** (2.064)	−0.082*** (−3.945)	−0.314	0.074*** (3.655)
	$\ln K_t$	−5.235[4]	1980	1999	0.036 (0.464)	−0.299*** (−5.582)	0.131* (1.871)	0.057** (2.393)
	$\Delta \ln K_t$	−6.490[4]	1982	1991	−0.354*** (−3.953)	0.226*** (4.862)	−0.160** (−2.048)	−0.068**** (−2.182)
	$\ln L_t$	−4.535[4]	1985	2002	−0.015 (−0.763)	−0.043*** (−4.724)	−0.049** (−2.422)	0.009 (1.198)
	$\Delta \ln L_t$	−7.394***[2]	1983	1996	0.030 (1.665)	0.029*** (1.665)	0.042** (2.418)	−0.006 (−0.835)
Chile	$\ln Y_t$	−5.003[2]	1990	2000	−0.053 (−1.133)	0.155*** (5.260)	0.019 (0.427)	−0.060*** (−2.906)
	$\Delta \ln Y_t$	−7.376***[1]	1981	1988	−0.295*** (−9.706)	0.092*** (4.573)	0.063** (2.191)	−0.118*** (−7.149)
	$\ln H_t$	−4.704[1]	2007	–	0.220* (1.779)	−0.191*** (−3.236)	–	–
	$\Delta \ln H_t$	−6.054[1]	1993	–	−0.049** (−2.422)	−0.060*** (−2.906)	–	–
	$\ln K_t$	−4.981[1]	1987	1998	−0.116 (−0.909)	0.347*** (5.117)	−0.211* (−1.715)	−0.113** (−2.169)
	$\Delta \ln K_t$	−5.761[1]	1981	1986	−0.962*** (−7.863)	0.332*** (4.053)	0.321*** (2.797)	−0.429*** (−5.653)
	$\ln L_t$	−4.540[4]	1983	1996	−0.028* (−1.706)	−0.009 (−1.204)	0.008 (0.453)	−0.036*** (4.591)
	$\Delta \ln L_t$	−5.856 [3]	1980	1991	0.067*** (3.579)	−0.043*** (4.099)	−0.067*** (−3.734)	0.040*** (4.387)
Colombia	$\ln Y_t$	−3.776 [1]	1982	1998	−0.008 (−0.507)	−0.021*** (−2.715)	−0.053*** (−3.227)	−0.017 (−2.204)
	$\Delta \ln Y_t$	8.049*** [4]	1991	1999	0.008 (0.487)	0.026*** (2.657)	0.094*** (5.024)	−0.012 (−1.448)
	$\ln H_t$	−5.109[0]	1990	1996	0.029 (0.441)	−0.126*** (−3.611)	−0.118*** (−1.963)	0.024 (0.776)
	$\Delta \ln H_t$	−9.056*** [3]	1991	1997	−0.651*** (−8.624)	0.364 (7.980)	0.233*** (3.104)	−0.339*** (−7.533)
	$\ln K_t$	−5.265[1]	1997	2003	0.160* (1.750)	−0.244*** (−4.783)	−0.097 (−1.156)	0.330*** (5.580)
	$\Delta \ln K_t$	−6.567*** [0]	1991	1999	0.066 (0.886)	0.065** (1.807)	0.308**8 (4.547)	−0.022 (−0.625)
	$\ln L_t$	−4.540 [4]	1983	1996	−0.028* (−1.706)	−0.009 (−1.204)	0.008 (0.453)	−0.036*** (−4.591)
	$\Delta \ln L_t$	−7.436*** [0]	1984	1995	0.009 (0.532)	−0.001 (−0.283)	−0.047*** (−2.882)	0.001(0.031)
Ecuador	$\ln Y_t$	−3.607[4]	1980	1996	0.056* (1.768)	−0.107*** (−3.937)	0.037 (1.257)	−0.008 (−0.745)
	$\Delta \ln Y_t$	−8.104***[0]	1980	2000	0.016 (0.517)	−0.040** (−2.553)	−0.001 (−0.043)	0.025** (2.154)
	$\ln H_t$	−5.185[0]	1982	1995	0.345** (1.963)	0.254*** (3.088)	0.210 (1.244)	−0.379 (−4.153)
	$\Delta \ln H_t$	−8.356*** [1]	1980	1986	−0.954*** (−5.140)	0.492 (4.542)	−0.141 (−0.902)	−0.663*** (−6.381)
	$\ln K_t$	−4.769 [1]	1980	2002	0.109 (1.074)	−0.283 (−4.989)	−0.090 (−0.905)	0.218*** (4.607)
	$\Delta \ln K_t$	−7.062*** [0]	1984	2000	0.221** (2.206)	−0.028 (−0.670)	0.153 (1.522)	0.026 (0.667)
	$\ln L_t$	−4.267 [0]	1997	2004	−0.013 (−0.791)	−0.036** (−3.192)	−0.008 (−0.510)	−0.004*** (−3.113)
	$\Delta \ln L_t$	−9.016*** [0]	1994	2000	0.025 (1.358)	−0.07*** (−3.126)	0.060*** (3.473)	0.005 (0.585)
Peru	$\ln Y_t$	−4.658[3]	1986	1997	0.149*** (3.076)	−0.171*** (−4.442)	−0.0685 (−1.575)	0.108*** (4.919)
	$\Delta \ln Y_t$	−6.969***[4]	1982	1990	−0.291*** (6.598)	0.172*** (5.984)	0.218*** (4.365)	−0.179*** (6.357)
	$\ln H_t$	−3.588[0]	1983	–	−0.001 (−0.014)	−0.028** (−1.991)	–	–
	$\Delta \ln H_t$	−9.056***[3]	1991	2000	−0.586*** (−9.537)	0.335*** (8.529)	0.055 (1.173)	−0.224*** (−7.449)
	$\ln K_t$	−4.318	1986	2006	0.239* (1.923)	−0.274*** (−3.277)	−0.020(−0.160)	0.2123*** (3.439)
	$\Delta \ln K_t$	−5.773[2]	1991	1998	−0.264*(1.831)	0.251***(3.596)	−0.203(−1.574)	−0.036 (−0.593)
	$\ln L_t$	−5.169	1990	2000	0.057(4.437)	−0.007 (−1.373)	0.032(2.855)	−0.009*(−1.783)
	$\Delta \ln L_t$	−7.081***[0]	1986	1994	−0.026* (−1.754)	0.025*** (3.254)	0.087*(1.617)	0.184* (1.814)
Venezuela	$\ln Y_t$	−4.642[4]	1983	2000	0.016 (0.315)	−0.070*** (−2.786)	0.034 (0.646)	−0.061** (−2.528)
	$\Delta \ln Y_t$	−5.421[0]	1983	2003	−0.160*** (−4.140)	0.018(1.086)	0.315***(6.672)	−0.117***(−4.353)
	$\ln H_t$	−4.819[3]	1980	1989	0.022 (0.351)	−0.155*** (−3.739)	−0.110* (−1.765)	0.084** (2.238)
	$\Delta \ln H_t$	−6.822[1]	1982	2003	0.096** (2.026)	0.048** (2.288)	0.203*** (4.142)	−0.066*** (−3.162)
	$\ln K_t$	−3.724[0]	1979	2005	−0.003 (−0.016)	−0.315*** (−3.665)	0.193 (1.094)	0.218** (2.507)
	$\Delta \ln K_t$	−6.267[1]	1978	1987	−0.219** (−2.214)	−0.024* (1.851)	0.235 (1.352)	−0.160** (−2.228)
	$\ln L_t$	−4.418[4]	1988	2001	−0.031** (−1.922)	−0.020*** (−3.405)	−0.051*** (−2.784)	0.016** (2.549)
	$\Delta \ln L_t$	−6.343[1]	1998	2007	−0.074*** (−4.647)	0.034*** (4.724)	0.017 (1.007)	−0.042*** (−4.208)

TB is the estimated break points. *, **, *** imply 10%, 5% and 1% levels of significance. Critical values are in Lee and Strazicich (2003: 2004). *TB1 and TB2* are the structural break dates. DU1 and DU2 are the dummy variables for breaks in intercept, while DT1 and DT2 are the dummy variables for trend breaks. Critical values for the other coefficients are based on the standard *t*-distribution 1.65, 1.96, 2.58 with maximum lag set at 4, the optimal lags are selected based on Akaike Information Criterion. The estimates are free of serial correlation and heteroscedasticity. The optimal lag length is reported in the brackets, while the t-statistics are reported in parenthesis.

* Denote significance at 10%, respectively.

** Denote significance at 5%, respectively.

*** Denote significance at 1%, respectively.

traditional unit root tests-including the Said and Dickey [46] or ADF, Phillips and Perron [47] or PP, Elliott et al. [48] or DF-GLS to examine the nonstationarity of the four series. The results which are reported in Table 2 reveal that we cannot reject the null of nonstationarity when the variables are in level. However, the null hypothesis cannot be rejected once the series are in the first difference. The outputs of these tests become questionable, in the presence of structural break (s). Consequently, the test statistics of the Lee and Strazicich [33,34] tests are presented in Table 3. We cannot reject the null of unit root for all the variables, when specified in level form. Expressed in their first differences, the null of unit root can be rejected in all cases. Almost 28% of the structural breaks are concentrated in the late 1970s

Table 4
Cointegration test with no structural break.

Country	Null hypothesis	Trace-statistic	Critical value		
			99%	95%	90%
Argentina	$r=0$	57.287***	54.682	47.856	44.494
	$r\leq1$	22.855	35.458	29.797	27.067
	$r\leq2$	8.602	19.937	15.495	13.429
	$r\leq3$	1.313	6.635	3.841	2.706
Brazil	$r=0$	88.369**	54.682	47.856	44.494
	$r\leq1$	33.571	35.458	29.797	27.067
	$r\leq2$	11.877	19.937	15.495	13.429
	$r\leq3$	0.022	6.635	3.841	2.706
Chile	$r=0$	42.837	54.682	47.856	44.494
	$r\leq1$	17.360	35.458	29.797	27.067
	$r\leq2$	3.169	19.937	15.495	13.429
	$r\leq3$	0.170	6.635	3.841	2.706
Colombia	$r=0$	67.610***	54.682	47.856	44.494
	$r\leq1$	35.622***	35.458	29.797	27.067
	$r\leq2$	14.774*	19.937	15.495	13.429
	$r\leq3$	0.006	6.635	3.841	2.706
Ecuador	$r=0$	64.505***	54.682	47.856	44.494
	$r\leq1$	25.813***	35.458	29.797	27.067
	$r\leq2$	9.927	19.937	15.495	13.429
	$r\leq3$	0.350	6.635	3.841	2.706
Peru	$r=0$	58.878***	54.682	47.856	44.494
	$r\leq1$	25.431	35.458	29.797	27.067
	$r\leq2$	8.097	19.937	15.495	13.429
	$r\leq3$	0.548	6.635	3.841	2.706
Venezuela	$r=0$	61.716***	54.682	47.856	44.494
	$r\leq1$	25.173	35.458	29.797	27.067
	$r\leq2$	11.645	19.937	15.495	13.429
	$r\leq3$	0.510	6.635	3.841	2.706

* Denote significance at 10%, respectively. r is the cointegration rank.
** Denote significance at 5%, respectively. r is the cointegration rank.
*** Denote significance at 1%, respectively. r is the cointegration rank.

and early 1980s and another 34% of the total breaks are located in the late 1990s and early 2000s.

Since the series are integrated of order (1), it is appropriate to employ the Johansen approach to test for possible cointegration relationship in the variables. As a benchmark, we start by reporting the trace statistics without structural breaks of Johansen [40,42]. The results in Table 4 show that we can reject the null of no cointergation in Argentina, Brazil, Chile, Colombia, Ecuador, Peru and Venezuela. However, we cannot reject the null of no cointegration in the case of Chile. The implication of the results is that with the exception of Chile, there are cointegrating vectors among the variables.

Due to the structural break issue associated with the cointegration test without break, we proceed to presenting the cointegration tests with two structural breaks in linear trend, as proposed by Johansen et al. [41]. The structural breaks are selected based on the most recurring dates observed in the stationarity test[5]. Reported in Table 5, the results provide evidence for cointegrating vector in all the countries at 1% significance level. While one cointegrating vector exists in Argentina, Chile, Colombia, Ecuador, Venezuela, two cointegrating vectors are established in Brazil and Peru. Having established that the variables are cointegrated, we proceed with the Granger causality test within the framework of VECM. We start with the results of Granger causality test without structural breaks, as reported in Table 6. The results indicate that causality flows from hydroelectricity consumption, capital and labour force to economic growth in the long run in Argentina, Brazil, Colombia, Ecuador, Peru and Venezuela. The unidirectional causality is in agreement with the growth hypothesis. However, the study fails to observe any causality in the variables for Chile.

Consistent with the cointegration tests, we turn to the causality test that provides for two structural breaks in the trend in Table 7. There is long run causality between hydroelectricity consumption and economic growth in the countries. Bidirectional causality between hydroelectricity consumption and economic growth is observed in the case of Argentina and Venezuela in the long run. There is evidence for long run unidirectional causality from hydroelectricity consumption to economic growth in Brazil, Chile, Colombia, Ecuador and Peru. However, limited evidence of causality between the two variables is found in the short run. There is short run causality from economic growth to hydroelectricity consumption in Colombia and Venezuela. Furthermore, there is short run causality running from hydroelectricity consumption to economic growth in Argentina and Chile. No short run causality between hydroelectricity consumption and economic growth is found in Brazil, Ecuador and Peru.

Considering the relationship with other variables, there is evidence for long run causality running from capital stock and labour force to economic growth with feedback from economic growth to capital stock in Argentina, Brazil, Chile and Venezuela. There is also feedback from economic growth to labour force in Argentina, Chile and Venezuela. There is evidence for long run causality running from capital stock and labour force to hydroelectricity in Argentina and Venezuela. On the other hand, there is long run causality from hydroelectricity consumption to capital stock in Argentina, Brazil, Chile and Venezuela and there is long run causality from hydroelectricity consumption to labour force in Argentina, Chile and Venezuela. Although the foregoing causal analysis is informative, they however do not capture the pattern in which explanatory variables influence the dependent variable. Positive and significant signs of hydroelectricity consumption in a growth equation suggest that expansionary energy proposition is worthwhile. The results, which are reported in Table 8, show that hydroelectricity consumption, capital stock and labour force have positive impact on economic growth in most cases.

5. Discussion

The study provides evidence for long run unidirectional causality from hydroelectricity consumption to economic growth in Brazil, Chile, Colombia, Ecuador and Peru. Previous studies with the same findings include Ben Aïssa et al. [18] for 11 African countries; Pao et al. [21] for Mexico, Indonesia, South Korea; Bildirici and Özaksoy [27] for the case of Hungary and Poland; Omri et al. [30] for Hungary, India, Japan, Netherlands, and Sweden. The results further show the existence of bidirectional causality between hydroelectricity consumption and economic growth in the case of Argentina and Venezuela in the long run, which is consistent with the analysis of Apergis and Payne [16] for six Central American countries; Apergis and Payne [17] for 80 developing and developed countries; and Salim et al. [23] in 29 OECD countries.

The findings further suggest that two structural breaks are significant in most cases. Besides, the study shows that more causality is found when structural breaks are introduced in the models. Providing for structural breaks provides the ability to reject a false null hypothesis [13,50–52]. Therefore, some of the previous studies on renewable energy consumption–economic growth nexus that have failed to observe causality might be due

[5] We differ from studies such as dos Santos and Kassouf [49] that have utilized the average of the dates from the unit root tests with structural breaks to determine the break dates for cointegration test. The rationale for choosing the most recurring (mode) in our case is that an actual break is experienced in this particular date unlike in the case of averaging whereby no break might have actually occurred in the chosen date.

Table 5
Cointegration test with two structural breaks.

Country	Null hypothesis	Trace-statistic	TB1	TB2	Critical value		
					99%	95%	90%
Argentina	$r=0$	186.319***	1990 (4.888)	2001 (0.681)	109.23	102.66	95.21
	$r\le 1$	98.282***			78.83	72.91	66.75
	$r\le 2$	52.641			52.17	45.47	42.13
	$r\le 3$	22.103			28.66	23.46	20.96
Brazil	$r=0$	141.662***	1982 (0.302)	2002 (0.767)	113.23	103.85	99.06
	$r\le 1$	58.289			82.00	73.89	69.79
	$r\le 2$	34.823			54.43	47.69	44.32
	$r\le 3$	15.144			29.93	24.70	22.17
Chile	$r=0$	120.182***	1981(0.279)	2000(0.721)	113.73	104.36	99.58
	$r\le 1$	51.422			82.37	74.28	70.19
	$r\le 2$	28.927			54.64	47.93	44.58
	$r\le 3$	13.490			29.94	24.77	22.27
Colombia	$r=0$	122.453***	1991(0.512)	1999 (0.698)	111.40	102.06	97.30
	$r\le 1$	38.638			80.47	72.39	68.30
	$r\le 2$	16.061			53.31	46.56	43.20
	$r\le 3$	0.083			29.44	24.13	21.58
Ecuador	$r=0$	128.546***	1980(0.256)	2000(0.721)	113.40	104.04	99.26
	$r\le 1$	61.150			82.10	74.01	69.92
	$r\le 2$	16.126			54.46	47.73	44.38
	$r\le 3$	13.975			29.91	24.70	22.18
Peru	$r=0$	189.557***	1986 (0.395)	1998 (0.674)	114.54	105.13	100.33
	$r\le 1$	103.388***			83.15	75.003	70.29
	$r\le 2$	36.414			55.27	48.56	45.20
	$r\le 3$	0.222			30.14	25.05	22.59
Venezuela	$r=0$	119.791***	1983(0.326)	2003 (0.791)	112.77	103.39	98.60
	$r\le 1$	62.011			81.65	73.54	69.44
	$r\le 2$	26.439			54.21	47.45	44.09
	$r\le 3$	2.856			29.83	24.58	22.05

r is the cointegration rank. The location of the structural break is reported in the brackets. *TB1* and *TB2* are the structural break dates.

*** Denote significance at 1%, respectively.

Table 6
Causality test with no break.

Country	Dependent variable	Short run causality				Long run causality
		$\Delta \ln Y_{t-i}$	$\Delta \ln H_{t-i}$	$\Delta \ln K_{t-i}$	$\Delta \ln L_{t-i}$	ECT_{t-1}
Argentina	$\Delta \ln Y_t$		2.699 [0.609]	1.660 [0.798]	4.574 [0.334]	−0.254*** (−3.045)
	$\Delta \ln H_t$	2.868 [0.580]		2.370 [0.668]	1.257 [0.867]	1.684 (1.302)
	$\Delta \ln K_t$	1.753 [0.781]	2.967 [0.563]		3.763 [0.439]	−0.299** (−2.539)
	$\Delta \ln L_t$	1.432 [0.839]	5.346 [0.254]	3.297 [0.509		−0.828* (−1.709)
Brazil	$\Delta \ln Y_t$		6.641 [0.156]	3.051 [0.549]	7.387 [0.117]	−0.447* (−1.734)
	$\Delta \ln H_t$	1.547 [0.818]		1.840 [0.765]	3.591 [0.464]]	0.192 (0.571)
	$\Delta \ln K_t$	17.879*** [0.001]	13.249**[0.01]		10.097** [0.039]	−2.340*** (−4.428)
	$\Delta \ln L_t$	4.391 [0.3556]	5.821 [0.213]	4.979 [0.290]		0.354** (1.962)
Chile	$\Delta \ln Y_t$		1.010 [0.908]	7.072 [0.132]	4.213 [0.378]	
	$\Delta \ln H_t$	3.729 [0.444]		4.426 [0.351]	1.215 [0.876]	
	$\Delta \ln K_t$	6.449 [0.168]	1.373 [0.849]		3.703 [0.448]	
	$\Delta \ln L_t$	4.024 [0.403]	3.238 [0.519]	3.808 [0.433]		
Colombia	$\Delta \ln Y_t$		6.633 [0.157]	7.709 [0.103]	9.410* [0.052]	−0.557** (−2.375)
	$\Delta \ln H_t$	16.761*** [0.002]		15.553*** [0.004]	7.110 [0.130]	−0.682 (−0.832)
	$\Delta \ln K_t$	8.661* [0.070]	4.168 [0.384]		15.073*** [0.005]	3.250 (2.903)
	$\Delta \ln L_t$	2.651 [0.618]	2.259 [0.688]	1.616 [0.806]		0.108 (0.459)
Ecuador	$\Delta \ln Y_t$		4.843 [0.304]	9.471* [0.050]	2.907 [0.574]	−0.253** (−2.201)
	$\Delta \ln H_t$	14.011*** [0.007]		8.750* [0.068]	12.544** [0.014]	2.394*** (4.441)
	$\Delta \ln K_t$	5.476 [0.242]	3.080 [0.545]		4.069 [0.397]	− 1.113** (−2.374)
	$\Delta \ln L_t$	3.663 [0.454]	2.896 [0.575]	3.692 [0.449]		−0.018 (−0.198)
Peru	$\Delta \ln Y_t$		1.460 [0.834]	4.128 [0.389]	5.731 [0.220]	−0.099* (−1.732)
	$\Delta \ln H_t$	1.766 [0.779]		2.727 [0.604]	1.554 [0.817]	0.177 (0.940)
	$\Delta \ln K_t$	4.740 [0.315]	3.427 [0.489]		4.911 [0.297]	0.029 (0.081)
	$\Delta \ln L_t$	3.387 [0.495]	5.104 [0.277]	1.886 [0.757]		−0.116*** (−2.701)
Venezuela	$\Delta \ln Y_t$		1.606 [0.808]	0.588 [0.964]	0.497 [0.974]	−0.403** (−2.197)
	$\Delta \ln H_t$	12.400** [0.015]		10.168** [0.038]	5.322 [0.256]	1.015*** (4.295)
	$\Delta \ln K_t$	1.307 [0.860]	2.664 [0.616]		0.849 [0.932]	1.436 (1.473)
	$\Delta \ln L_t$	0.757 [0.944]	3.526 [0.474]	0.451 [0.978]		0.002 (0.015)

The estimates are free of serial correlation and heteroscedasticity. The probability values are reported in the brackets, while the t-statistics are reported in parenthesis.

* Denote significance at 10%, respectively.

** Denote significance at 5%, respectively.

*** Denote significance at 1%, respectively.

Table 7
Causality test with two structural breaks.

Country	Dependent variable	Short run causality				Long run causality
		$\Delta \ln Y_{t-i}$	$\Delta \ln H_{t-i}$	$\Delta \ln K_{t-i}$	$\Delta \ln L_{t-i}$	ECT_{t-1}
Argentina	$\Delta \ln Y_t$		16.193*** [0.006]	12.050 [0.034]	13.153** [0.022]	−0.373*** (-3.739)
	$\Delta \ln H_t$	7.594 [0.180]		8.294 [0.141]	5.144 [0.399]	−0.738* (−1.915)
	$\Delta \ln K_t$	8.936 [0.112]	11.417** [0.044]		8.340 [0.138]	−12.050*** (−3.031)
	$\Delta \ln L_t$	15.248*** [0.009]	33.928*** [0.000]	19.108*** [0.002]		−0.995*** (−4.047)
Brazil	$\Delta \ln Y_t$		5.865 [0.210]	9.142** [0.058]	10.458* [0.033]	−0.332*** (−2.804)
	$\Delta \ln H_t$	1.797 [0.773]		2.405 [0.662]	3.248 [0.517]	−0.277 (−0.374)
	$\Delta \ln K_t$	78.506*** [0.000]	25.664*** [0.000]		46.512*** [0.000]	−5.744*** (−8.588)
	$\Delta \ln L_t$	2.083 [0.721]	3.540 [0.472]	3.120 [0.538]		0.386 (0.899)
Chile	$\Delta \ln Y_t$		20.886*** [0.000]	23.711*** [0.000]	18.058*** [0.000]	−0.687*** (−4.539)
	$\Delta \ln H_t$	2.297 [0.807]		2.476 [0.780]	2.742 [0.740]	−0.407 (−0.593)
	$\Delta \ln K_t$	43.238*** [0.000]	23.708*** [0.000]		9.463* [0.092]]	−2.215*** (−5.084)
	$\Delta \ln L_t$	22.723*** [0.000]	18.038*** [0.000]	22.158*** [0.000		−0.290*** (−4.138)
Colombia	$\Delta \ln Y_t$		6.517 [0.259]	6.261 [0.282]	2.246 [0.814]	−0.900* (−1.890)
	$\Delta \ln H_t$	27.931*** [0.000]		13.898** [0.016]	9.655* [0.086]	−2.163 (−1.561)
	$\Delta \ln K_t$	3.650 [0.601]	6.815 [0.235]		1.407 [0.923]	−3.556 (−1.476)
	$\Delta \ln L_t$	5.492 [0.359]	6.632 [0.250]	1.713 [0.887]		0.757* (1.713)
Ecuador	$\Delta \ln Y_t$		3.481 [0.481]	9.792** [0.044]	3.859 [0.425]	−0.683** (−1.984)
	$\Delta \ln H_t$	6.572 [0.160]		6.119 [0.190]	4.709 [0.319]	−1.620 (−0.898)
	$\Delta \ln K_t$	2.680 [0.612]	4.605 [0.330]		3.532 [0.473]	4.729 (4.256)
	$\Delta \ln L_t$	5.714 [0.222]	2.654 [0.617]	9.254* [0.055]		0.574 (2.429)
Peru	$\Delta \ln Y_t$		2.428 [0.787]	7.754 [0.170]	1.689 [0.890]	−0.897* (−1.872)
	$\Delta \ln H_t$	5.596 [0.348]		6.746 [0.240]	1.936 [0.858]	−2.069 (−1.584)
	$\Delta \ln K_t$	3.771 [0.583]	0.709 [0.983]		5.833 [0.323]	2.402 (0.887)
	$\Delta \ln L_t$	0.816 [0.976]	0.739 [0.981]	0.329 [0.997]		−0.088 (−0.219)
Venezuela	$\Delta \ln Y_t$		1.087 [0.581]	2.664 [0.264]	3.906 [0.142]	−0.741*** (−3.473)
	$\Delta \ln H_t$	10.554*** [0.005]		4.867* [0.088]	1.570 [0.456]	−1.799*** (−3.101)
	$\Delta \ln K_t$	2.501 [0.286]	0.391 [0.823]		2.513 [0.285]	−2.443*** (−1.426)
	$\Delta \ln L_t$	3.063 [0.216]	0.063 [0.969]	1.789 [0.409]		−0.511*** (−3.367)

The estimates are free of serial correlation and heteroscedasticity. The probability values are reported in the brackets, while the *t*-statistics are reported in parenthesis

* Denote significance at 10%, respectively.
** Denote significance at 5%, respectively.
*** Denote significance at 1%, respectively.

Table 8
Long run estimates.

Country	Independent variable	Dependent variable: $\ln Y_t$	
		Johansen test with no breaks	Johansen test with two breaks
Argentina	$\ln H_t$	−0.017 (−0.960)	0.051*** (12.159)
	$\ln K_t$	−0.245 (−6.716)	0.300*** (37.317)
	$\ln L_t$	−1.059 (−10.563)	0.259*** (4.098)
Brazil	$\ln H_t$	0.373*** (−8.649)	0.423*** (24.653)
	$\ln K_t$	−0.082** (−2.393)	0.099*** (5.955)
	$\ln L_t$	−0.726*** (−9.050)	0.362*** (9.663)
Chile	$\ln H_t$	–	0.661*** (3.479)
	$\ln K_t$	–	0.178*** (3.579)
	$\ln L_t$	–	2.094*** (4.556)
Colombia	$\ln H_t$	−0.261*** (−7.406)	0.420*** (6.268)
	$\ln K_t$	−0.118*** (−3.601)	0.080*** (4.889)
	$\ln L_t$	0.6010*** (14.051)	2.237*** (11.941)
Ecuador	$\ln H_t$	−0.073 (−1.304)	0.045** (1.962)
	$\ln K_t$	−0.073 (−1.062)	−0.528*** (11.005)
	$\ln L_t$	−0.173 (−0.788)	0.810*** (5.579)
Peru	$\ln H_t$	2.046*** (7.044)	0.321*** (2.752)
	$\ln K_t$	−0.390*** (−5.199)	0.288*** (17.481)
	$\ln L_t$	2.165*** (6.057)	0.071 (0.284)
Venezuela	$\ln H_t$	0.300*** (4.751)	0.424 (1.199)
	$\ln K_t$	0.438*** (12.451)	0.714*** (9.507)
	$\ln L_t$	0.040*** (0.274)	1.863** (1.990)

The estimates are free of serial correlation and heteroscedasticity. The *t*-statistics are reported in parenthesis.

** Denote significance at 5%, respectively.
*** Denote significance at 1%, respectively.

in part to the non-inclusion of structural breaks. About 31 breaks or 28% of the total breaks are located in the late 1970s and early 1980s. This period coincides with the period when Latin American countries experienced a situation whereby their earning powers were lower than their foreign debt. As such they were unable to pay their debt obligations as their interest rates continued to increase. From the mid-1970s, the governments in the region heavily borrowed from industrialized countries [53].

Another 37 breaks or 34% of the total breaks are located in the late 1990s and early 2000s. These periods coincide with the phase of shocks in several Latin American countries. During that era, a long lasting recession started in Argentina and Uruguay, which finally led to a destabilizing economic crisis. Moreover, Brazil underwent severe stress tests in the early 2000s, mainly as a result of uncertainties related to increasing spreads worldwide, particularly on U.S. corporate bonds. The economy registered unexpected decline in capital flows of almost 6% of GDP, depreciation of the exchange rate of almost 50%, and a considerable rise in the Brazilian bond spread over U.S. Treasury bonds and the rising public debt profile of the country [51]. Subsequent to the end of the asset price bubble in the United States and the September 11 attacks, Chilean economy was also affected during that time which is manifested in the contraction of its output [54]. Similarly, Ecuador also suffered its worst economic crisis during the same period [55]. The 1999 economic downturn in the country was among the most serious and inflation hit record highs in the following year. The most important cause of the crisis was due to a combination of policy-induced and exogenous shocks which caused the market to lose confidence in both the domestic currency and the banking system, while government debts rose drastically. The destruction of wealth during the crisis and the associated increased in unemployment rate caused the biggest emigration flow in the country's history [55].

There is limited causality between hydroelectricity consumption and economic growth in the short run. This is not surprising as most hydroelectric stations have long lives. The projects entail large-scale investments and the gestation period is long term in nature. Although hydroelectricity power plants require low operating and maintenance cost (and also lower even in comparison with other types of power generators) they require long term investments that require long term planning agreements. Consequently, the effect of hydroelectricity projects may not be felt in the short run but rather in the long run.

6. Conclusion and policy implications

The purpose of this study is to examine the relationship between hydroelectricity consumption and economic growth in Latin America countries for the period of 1970–2012. We analyze the causal relationship between hydroelectricity consumption and economic growth within a neoclassical model that includes capital and labour force. Having established that the variables are cointegrated, the results of Granger causality test without structural breaks indicate that causality runs from hydroelectricity consumption, capital and labour force to economic growth in the long run in Argentina, Brazil, Colombia, Ecuador, Peru and Venezuela. The unidirectional causality is in agreement with the growth hypothesis. However, we did not find any causality between the variables for Chile. Using the causality tests with two structural breaks, we observe bidirectional causality between hydroelectricity consumption and economic growth in the case of Argentina and Venezuela in the long run. There is evidence for long run unidirectional causality from hydroelectricity consumption to economic growth in Brazil, Chile, Colombia, Ecuador and Peru. However, limited evidence of causality between the two variables is found in the short run.

The implication of these results is that expansion in hydroelectricity activities will boost economic growth, especially in the long run. Therefore, the adoption of polices to conserve the use of hydroelectricity will adversely affect the real GDP in the region. Any shortage of the hydropower resources will also inhibit economic growth and any shocks to hydroelectricity will be passed to the output. In Argentina and Venezuela, where the bidirectional causality is confirmed, any shocks to economic growth will also be passed to hydroelectricity and the chain will persist via the feedback flow.

Since expansionary hydroelectricity policies are beneficial to the countries in the region, substituting fossil fuels with hydropower should be considered as a feasible policy as this will reduce emission problems in the country. With abundant hydroelectricity resources, pursuing such policies is not insurmountable. For instance, Brazil is yet to tap into 70% of its hydropower potential, while Argentina, Ecuador and Venezuela have estimated potentials of 40 GW, 18.9 GW and 83.4 GW, respectively [9–11,56].

In this direction, there are some on-going activities meant to improve the availability of hydroelectricity. Brazil is planning new hydroelectric power projects that include the Belo Monte plant. It is projected that upon the completion of the project, it will be the third biggest hydroelectric power plant in the world [5]. Argentina has completed its Energy Policy 2030 Plan, which shows that the country aims to raise the contribution of hydropower in the total electricity generated and also decrease the share of gas in the electricity mix from 52% to 30%. In the same vein, Chile has recently published its National Energy Strategy 2012–2030 roadmap, which shows that the country aims to raise the contribution of hydropower in the total electricity generated from the current 34% to 48%. Peru, which now source 65% of its electricity from hydropower, is currently undertaking the construction of the 406 MW Chaglla hydropower plant, planned to become operational by 2016, and has also contracted the development of the 98 MW Santa Teresa and the 510 MW Cerro del Aguila hydropower plants [10]. In Colombia, the two major hydropower projects that are under construction are the 820 MW Sogamoso and the 2400 MW Ituango projects. The Ituango station is expected to become the biggest hydropower plant in Colombia, once it commences operation in 2018, setting the scene for the country's future plans [10]. To address capacity shortages, the government of Ecuador has roll out blueprints to construct six additional hydroelectric power plants with a combined capacity of 2.8 GW in the near future [7]. Venezuela plans to expand hydroelectric production in the future [8].

Since the study provide evidence for nonstationarity of the series at level, any action triggered by the authorities to affect these variables are likely to be effective. Therefore, the blueprints aimed at increasing long term deployment of hydroelectricity will be effective [57,58].

The evidence provides that causality runs from labour and capital stock to output in all cases with feedback from output. Therefore, the developments in hydropower sub-sector alone are insufficient to promote economic growth; it must be complemented with other factors that may improve labour and capital productivity. These include better machinery and equipment, boosting research and development; promoting savings and investment, improving health care facilities, and enhancing property rights and rule of law. These policies may not only directly influence economic growth, but also instigate economic growth through the development of the energy sector.

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