In the Short Run, Energy Efficiency Concerns and Trade Protection Hurt Each Other and Growth, but in the Long Run, not Necessarily so: 1980-2010 Latin American Evidence

Alexandra Tsiotras
Université Libre de Bruxelles

Antonio Estache
SBS-EM, ECARES, Université Libre de Bruxelles and CEPR

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ECARES
ULB - CP 114/04
50, F.D. Roosevelt Ave., B-1050 Brussels BELGIUM
www.ecares.org

In the short run, energy efficiency concerns and trade protection hurt each other and growth, but in the long run, not necessarily so: 1980-2010 Latin American Evidence [1]

Alexandra Tsiotras (Université libre de Bruxelles)

Antonio Estache (Université Libre de Bruxelles, ECARES, CEPR)

Abstract:

The paper studies the 3-way causal relationships between energy consumption, output and trade for a sample of 15 Latin American countries over the period 1980 to 2010. The results of our panel cointegration and error-correction model based on GMM estimators highlight a unidirectional relationship running from energy use to real GDP (in the short and long run) and from energy use to exports (in the short run). This confirms earlier results for a smaller sample of countries in the region and shows that energy consumption cuts can have significant economic costs. In contrast to earlier results, we find that these conclusions are more robust in the short than in the long run, suggesting that if technological change (in particular energy efficiency improvements) is accounted for, the growth and trade costs of energy consumption cuts should be lower than often feared. Energy efficiency improvements appear to be happening. The case for energy efficiency improvements further increases with the finding that under current technologies, cutting energy consumption would hurt growth more than an import substitution policy.

[1] We are grateful to Paolo Bolsi, Catherine Dehon, Domenico Giannone, Lorenzo Ricci and Maurizio Zanardi for suggestions and comments. Any mistake is ours and ours only.

1. Introduction

Lowering energy consumption continues to be central to fight undesirable emissions and pollution around the world. A common source of reluctance to do so is the concern for its impact on growth and trade. The extent to which energy and growth policies influence each other has already been widely analysed (e.g. Kalimeris et al. (2014) for a recent meta-analysis of 158 studies of the causality between energy and GDP covering the period 1978-2011). One of the lessons of this research so far is that there is a wide diversity of results. These depend on the sample size (countries and periods), the sample types (panel vs. time series vs. cross-sections), the specificity of the model (production vs. others), the techniques (cointegration, error-correction, diff-in diff…) and the horizon (short vs. long term)[2].

For Latin America, the literature on the relationship between energy use and economic growth has identified more coherent general results than for other regions[3]. More precisely, empirical studies dealing only with Latin America provide evidence of an adverse effect of energy conservation policies on growth (i.e. the so-called growth hypothesis). For instance, Apergis and Payne (2009 and 2010) relied on a panel cointegration and error-correction model for six Central American and nine South American countries to generate their evidence.

The empirical evidence on the specific importance of trade for growth and vice versa has also enjoyed a large volume of research supported by a wide range of techniques as well. Although the results tend to suggest a two-way causality, once again, some of the conclusions are sample, model, horizon and techniques specific. Moreover, causality studies often focus on trade openness rather than on the distinction between the interactions between growth on the one hand and exports and imports on the other. For instance, Awokuse (2008) shows not only that policies promoting imports foster economic growth in Argentina, Colombia and Peru, but also found some indications of a unidirectional causation running from exports to economic growth in Peru.

The extent to which energy policies may interact with trade has enjoyed a lot less research. Cole (2006) is widely quoted as one of the precursors with an assessment of the interactions between trade liberalization and energy consumption for a sample of 32 countries[4]. The topic continues to be analysed, but there is still room to improve our collective knowledge. For instance, Ghani (2012) enquired the same question by opting for a different empirical method, namely a difference-in-difference regression in income. He advocated that liberalization alone does not trigger a change in the growth rate of energy consumption. More importantly, Ghani's findings suggested that discrepant effects of trade liberalization exist and are linked to a country's economic development level.

The literature on the 3-ways interactions is even rarer[5]. To the best of our knowledge, there are about 15 published papers on the topic. These take three different angles, covering either a single country (e.g. Sami, 2011; Sultan, 2012; and Shahbaz et al., 2013), a homogeneous region (e.g. Hossain, 2012; Sadorsky, 2012; and Shakeel et al., 2013.) or a heterogeneous group of countries (e.g. Nasreen and Anwar, 2014). Their results with regard to the existence and direction of causation were

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[2] In his earlier survey, Ozturk's (2010) explained the divergence of results on the energy-growth nexus, with four factors: (i) the applied econometric approach combined with the model specification, (ii) the time span coverage of the analysis, (iii) the employed proxies for the variables of interest, and (iv) heterogeneous patterns of energy use and economic development.

[3] Ozturk (2010) and Payne (2010a and 2010b) carried out a literature survey on the relationship between energy use and economic output. They maintained that previous empirical works came to ambivalent results, highlighting the existence of four possible relationships between these two variables: conservation (Y->E), growth (E->Y), neutrality (E has no impact on E, nor Y on E) and feedback (E and Y interact) hypotheses.

[4] No exact information on the developed and developing countries included in Cole's (2006) sample is given.

[5] Our main variables of interest are energy consumption, economic output and trade. The latter is divided into goods exports and goods imports, which leads us to have four indicators in total and two distinct specifications of our model (export and import specification).

mostly inconclusive, even if multiple studies had focused on the same country (e.g. Lean and Smyth, 2010a and 2010b). For instance, while studies covering South Asian economies were not able to reach a consensus (e.g. Hossain, 2012; and Shakeel et al., 2013), more robust results were obtained for the Middle East (Narayan and Smyth, 2009; and Sadorsky, 2011). In particular, it appears that, for Middle Eastern countries, an alternation of trade pattern engenders a change in energy use. Moreover, a bidirectional relation exists between energy and economic output. For OECD countries between 1980 and 2010, Dedeoğlu and Kaya (2013), provide evidence of cointegration and feedback causations between energy use and exports; energy use and imports; and energy use and GDP.

Only one paper has done this for a sample of seven Latin American countries, namely Sadorsky (2012). He did so for a period spanning from 1980 to 2007, just before the recent crisis that initially hit OECD countries before tricking down to emerging and poorer economies. He relied on a panel cointegration and vector error-correction approach based on a production function framework[6]. He concluded that, in his small sample of countries, policies aiming at modifying energy consumption affected the economic growth and trade pattern of these economies, but also that trade restrictions and cutbacks in economic growth reduced energy use during that period.

Our paper contributes to the literature on energy use, economic growth and trade in three distinct ways. First, we more than double the number of Latin American countries as compared to Sadorsky (2012)[7]. This allows to better account for the large diversity in the access to energy resources in the region. Secondly, we expand the period to include 2008 to 2010, 3 years during which energy consumption dropped significantly in the region. Trade in Latin America was indeed impacted by the growth fall in the OECD. Finally, we rely on a cointegration and error-correction method as many authors have done, but we base it on GMM estimators and a panel dataset (instead of a time series approach). This approach is not only more powerful, but also offers more degrees of freedom. Furthermore, in settings where the lagged dependent variable is included in the regression and right-hand side variables are not perfectly uncorrelated with the error term, GMM estimators are efficient and consistent (Roodman, 2009).

The paper is organized as follows. Section 2 provides a brief overview of the Latin American context in terms of energy consumption, economic growth and trade. Section 3 describes the empirical method used to answer our research question. Section 4 is a description of our data. Section 5 discusses the results. Section 6 draws policy implications.

2. The Latin American context

Latin America continues to be quite an economically heterogeneous region in various ways of relevance to the assessment conducted here. First, they differ in terms of the speed at which they are growing. Some countries have been slow to pull out of poverty (as in Central America or some of the island economies), whereas some are already member of the OECD (Chile and Mexico) and Chile is no longer considered a developing country by international donor agencies. Second, they differ in terms of their degree of energy autonomy. Some are established producers and exporters (Venezuela, Mexico or Argentina), whereas many of the poorest ones are importers.

In spite of its strong heterogeneity, the region's contribution to the global economy has considerably increased over the last three decades, much of this by playing its part in the global growth of trade following the trade liberalization policies started mostly during the 1990s. In the process, the

[6]Note: We partially follow Sadorsky's (2012) approach by applying a production function to model the relationship between output, energy consumption and trade. However, in contrast to his model, we use an error-correction model based on GMM estimators. To ensure comparability, we applied our methodology on Sadorsky's (2012) sample and our results are partially in line with his findings.
[7] A small number of countries was omitted due to inexistence of complete data.

standard of living of a steadily fast-growing population has progressively improved (despite the continuous high distortions in the distribution of income).

This evolution engendered a surge in energy demand. With the growing concerns for pollution and climate change, this surge has now become a central policy concern (Vergara et al., 2013). The concern is somewhat mellower than in other regions thanks to the domination of renewables in the region (around 60% thanks to the huge hydro resources, which represent over 90% of renewables (IEA, 2013)).

The opportunities to increase the share of renewables are significant in the region. The Inter-American Development Bank expects the region to be able to cover all of its energy needs from renewable sources in the foreseeable future (Inter-American Development Bank, 2013; and Vergara et al., 2013). But as long as 40% of the sources of a growing energy demand continue to generate CO2, particles and other undesirable emissions, there will be a case for a push to improve energy efficiency in the long run and possibly slow consumption in the short run. This is why a better assessment of the causal tie between energy consumption, GDP and trade in this region is so important. It is essential to identify the effective importance of trade-offs during the transition to cleaner energy sources.

3. Empirical method

To test the causal relation between energy use, economic output and trade, we estimate a production function defining the interactions between the various variables we want to analyse. To do so, we rely on a panel cointegration and error-correction model based on GMM estimators, thereby combing the methodological approaches used by Lean and Smyth (2010b), Gries and Redlin (2012), Sadorsky (2012) and Yasar et al. (2006). This approach enables us to see whether coherent and robust results in terms of the dynamics at play are found in this region.

More specifically, our empirical methodology can be summarized in three main steps:

First, we start from a production function, following Lean and Smyth (2010b) and Sadorsky (2012). Output is a function of the three usual factors of production (capital (K), labour (L) and energy (E)), a trade component (T), reflecting either exports or imports, and a variable (F) that captures country specific properties:

$$Y_{it} = f(K_{it}, L_{it}, E_{it}, T_{it}, F_i) \quad (1)$$

The specific form estimated is based on a transformation of all variables in natural logarithms, which facilitates the interpretation of the results:

$$y_{it} = \beta_{1i}k_{it} + \beta_{2i}l_{it} + \beta_{3i}e_{it} + \beta_{4i}t_{it} + f_i + \varepsilon_{it} \quad (2)$$

where i = 1,..., N reflects the countries in our sample, t = 1,..., T reflects the time periods included and f_i represents the entity fixed effect.

The estimation of equation (2) yields the long-run elasticities of the growth regression. However, Pedroni (2001a and 2001b) argues that ordinary least squares (OLS) estimators are, in the panel cointegration context, asymptotically biased and provide inconsistent results. Therefore, we follow Sadorsky (2012) by applying a fully modified OLS (FMOLS) estimation technique to overcome this issue. More precisely, we apply a between-dimension FMOLS. This method does not only allow for more heterogeneity across cross-sections, but also establishes estimates that are not biased by correcting for endogeneity and serial correlation (Pedroni, 2001a and 2001b).

The second step arises from our interest in the short-run dynamics and Engle and Granger's (1987) suggestion that a standard Granger causality[8] test provides inadequate results in the presence of cointegration between the dependent and independent variables. Thus, we are carrying out five unit root tests and Kao's residual panel cointegration test.

The stationarity of our variables is first verified by using four panel unit root tests that assume cross-sectional independence. It is worth highlighting that, while three of the four assume the existence of an individual non-stationarity (Im et al., 2003; Maddala and Wu, 1999; and Choi, 2001), Levin et al.'s (2002) null hypothesis deviates by supposing common unit roots. Nevertheless, Table A.1 in the appendix shows that none of the four tests was able to reject the null hypothesis for the variables in levels at any commonly used significance level. Therefore, it can be maintained that all variables have a unit root in level, but, as illustrated in Table A.1, not in first differences. This means that a stochastic trend is inherent in levels, implying that persistent consequences arise from random shocks affecting not only the econometric method we rely on, but also the economic recommendations of measures and strategies to pursue (Stock and Watson, 2012; and Libanio, 2005).

As Sadorsky (2012), we apply Pesaran's (2004) test for cross-sectional dependence, since it is likely that our sample countries might be interconnected, for instance, due to the presence of regional or local spillovers (Boston College, 2011).Table 1 shows that cross-sectional dependence among all variables exists, implying that the previous panel unit root tests may provide inaccurate results. Accordingly, we rely on Pesaran's (2007) panel unit root test, which allows for cross-section dependence. It points to the non-stationarity in levels, but not in first differences.

Table 1: Pesaran's test for cross-sectional dependence

Variables	y	k	l	e	x	m
Pesaran's CD-test	54.52***	45.44***	56.41***	49.93***	46.39***	50.65***

Notes: Null hypothesis: cross-section independence; *** denotes significance at 1% level; ** 5% level; * 10% level; number of observation: 527

Given that all our variables are integrated of order one, we examine whether a cointegration in equation (2) exists. To do so, we apply Kao's (1999) residual cointegration test, which assumes homogeneous slope coefficients across all cross-sections. In particular, it tests the residual of equation (2) for its stationarity and provides two different statistics to determine whether to reject the null hypothesis of no cointegration[9]. In contrast to Pedroni's (1999, 2004) test for cointegration, it allows for less heterogeneity in the panel. Table 2 below illustrates that for both specifications of equation (2) we can reject the null hypothesis at 1% indicating that our variables are cointegrated. This suggests that our variables of interest pursue a common trend and therefore, a long-run relationship between them may exist. This has obvious long-term policy implications[10].

[8] "Granger causality means that if X Granger-causes Y, then X is a useful predictor of Y, given the other variables in the regression" (Stock and Watson (2012), p.580)

[9] We are aware that basing our cointegration analysis on the test developed by Pedroni (1999, 2004), which is widely used in the literature, would be more suitable. However, this test does not reject the null hypothesis of no cointegration in our sample, although both Kao's and Johansen's tests confirm the existence of cointegration. Thus, since there is a chance that cointegration may exists in our sample, we apply an ECM in order to avoid misspecification.

[10] However, Elliott (1998) argued that cointegration approaches often assume an exact non-stationarity, which is not always the case in our sample. Therefore, our inferences about the long run are highly fragile and need to be interpreted and treated with caution.

Table 2: Kao's residual cointegration test for the export and import specification of Equation (2)

	Export specification			Import specification		
	Coeff.	t-stat	Prob.	Coeff.	t-stat	Prob.
Kao Residual Cointegration Test – ADF		-4.1348	0.0000		-4.3365	0.0000
Number of observations			465			465
Augmented Dickey-Fuller Test equation	-0.1494	-5.4601	0.0000	-0.1542	-5.6864	0.0000
Number of observations			450			450

Notes: Null hypothesis: no cointegration; lag length selected based on the Schwarz Information Criterion, Newey-West automatic bandwidth selection and Bartlett kernel and no deterministic trend is assumed

The third step of our empirical strategy stems from the evidence of non-stationarity of the variables in levels and of cointegration between the variables used in equation (2). This evidence suggests that "*there must be some force that pulls the equilibrium error back towards zero*[11]". This is why we apply an error-correction model (ECM) to determine whether causation between energy use, economic output and trade exists.

In this respect, we are following the procedure of Yasar et al. (2006) and Gries and Redlin (2012), with the only difference that we decided to apply a two-step ECM based on GMM estimators instead of a one-step ECM. Using a two-step ECM implies that we first estimate equation (2) to determine the error correction term, which are the equation's residuals. Then, in a second step, we run the ECM by including the previously obtained error correction term. For this, we employ the Arellano and Bond (1991) and Blundell and Bond (1998) difference GMM and system GMM estimators and our ECM can be written as follows:

$$\Delta y_{it} = \propto_{1i} + \sum_{j=1}^{q} \gamma_{11ij}\Delta y_{it-j} + \sum_{j=1}^{q} \gamma_{12ij}\Delta k_{it-j} + \sum_{j=1}^{q} \gamma_{13ij}\Delta l_{it-j} + \sum_{j=1}^{q} \gamma_{14ij}\Delta e_{it-j} + \sum_{j=1}^{q} \gamma_{15ij}\Delta t_{it-j} + \gamma_{16i} ect_{it-1} + u_{1it} \quad (3a)$$

$$\Delta k_{it} = \propto_{2i} + \sum_{j=1}^{q} \gamma_{21ij}\Delta y_{it-j} + \sum_{j=1}^{q} \gamma_{22ij}\Delta k_{it-j} + \sum_{j=1}^{q} \gamma_{23ij}\Delta l_{it-j} + \sum_{j=1}^{q} \gamma_{24ij}\Delta e_{it-j} + \sum_{j=1}^{q} \gamma_{25ij}\Delta t_{it-j} + \gamma_{26i} ect_{it-1} + u_{2it} \quad (3b)$$

$$\Delta l_{it} = \propto_{3i} + \sum_{j=1}^{q} \gamma_{31ij}\Delta y_{it-j} + \sum_{j=1}^{q} \gamma_{32ij}\Delta k_{it-j} + \sum_{j=1}^{q} \gamma_{33ij}\Delta l_{it-j} + \sum_{j=1}^{q} \gamma_{34ij}\Delta e_{it-j} + \sum_{j=1}^{q} \gamma_{35ij}\Delta t_{it-j} + \gamma_{36i} ect_{it-1} + u_{3it} \quad (3c)$$

$$\Delta e_{it} = \propto_{4i} + \sum_{j=1}^{q} \gamma_{41ij}\Delta y_{it-j} + \sum_{j=1}^{q} \gamma_{42ij}\Delta k_{it-j} + \sum_{j=1}^{q} \gamma_{43ij}\Delta l_{it-j} + \sum_{j=1}^{q} \gamma_{44ij}\Delta e_{it-j} + \sum_{j=1}^{q} \gamma_{45ij}\Delta t_{it-j} + \gamma_{46i} ect_{it-1} + u_{4it} \quad (3d)$$

$$\Delta t_{it} = \propto_{5i} + \sum_{j=1}^{q} \gamma_{51ij}\Delta y_{it-j} + \sum_{j=1}^{q} \gamma_{52ij}\Delta k_{it-j} + \sum_{j=1}^{q} \gamma_{53ij}\Delta l_{it-j} + \sum_{j=1}^{q} \gamma_{54ij}\Delta e_{it-j} + \sum_{j=1}^{q} \gamma_{55ij}\Delta t_{it-j} + \gamma_{56i} ect_{it-1} + u_{5it} \quad (3e)$$

[11] Verbeek (2012), p. 347

where Δ denotes the first difference, q is the determined lag length, ect represents the residuals from equation (2) and thus, reflects the error correction term and u is a random error term.

From these five equations, it is possible to identify the short- and long-run causalities among variables by applying an F-test. More precisely, the existence of a short-run causation, for instance, running from real GDP to energy consumption (Equation (3d)) can be determined by checking whether $H_0 = \gamma_{41ij} = 0\ \forall_{ij}$ is rejected (Stock and Watson, 2012). By the same token, the long-run causal tie exists, when the t-test of the respective error correction term's coefficient indicates that the null hypothesis assuming no long-run causation ($H_0 = \gamma_{k6i} = 0\ \forall_i$, where k= 1,..., 5) can be rejected. In addition, the error correction terms "*give the adjustment rates at which short-run dynamics converge to the long-run equilibrium relationship*"[12].

As we are applying difference GMM and system GMM estimators, which use instruments to overcome the issue of autocorrelation and endogeneity, we need to check whether our used instruments are exogenous and valid. In this regard, we are computing the Sargan test for over-identification of restrictions and the Arellano-Bond test for serial correlation (Roodman, 2009).

Finally, we intend to explain the underlying factors that have led us to opt for this dynamic panel methodology. First, the literature argued that having a time and country dimension provides more accurate estimates compared to time series and that it is more powerful and possesses more degrees of freedom (Costantini, 2010; Osbat, 2004 and Pedroni, 2001a). In addition, GMM estimators are suitable and consistent in settings where the lagged dependent variable is included in the regression and right-hand side variables are imperfectly correlated with the error term (Roodman, 2009)[13].

4. Data

Our balanced dataset consists of annual observations from 1980 to 2010 for 15 Latin American countries[14]. The data was retrieved from the World Development Indicators (World Bank, 2010 and 2013), the UN Statistics Division (2013) and the Penn World Table Version 7.1 (Hestion et al., 2012). The main variables included are energy use (kt of oil equivalent), real GDP and real goods exports and imports (base year 2005). We assume, as other researchers (e.g. Apergis and Payne (2009, 2010); Sadorsky (2012); Sari and Soytas (2007)), that capital can be approximated by the real gross fixed capital formation[15] and labour is represented by a country's total labour force.

In terms of the shape parameters, all six variables appear to show a leptokurtic distribution (i.e. kurtosis is bigger than 3) indicating the existence of fat tails and are skewed (cf. Table A.2 in the appendix). This observation is particularly true for exports (skewness of 2.96 and kurtosis of 21.95), which seems to be highly asymmetric and peaked. These results imply that the distributions of energy consumption, output, capital, labour, exports and imports do not follow a normal distribution pattern, which may cause problems in the empirical method. However, when we use the variables in

[12] Gries and Redlin (2012), p. 8

[13] However, it is worth emphasising that our panel consists of a large number of time periods, whereas the number of countries is relatively small. Therefore, Roodman (2009) claimed that:
"*If T is large, dynamic panel bias becomes insignificant, and a more straightforward fixed-effects estimator works. Meanwhile, the number of instruments in difference and system GMM tends to explode with T. If N is small, the cluster-robust standard errors and the Arellano-Bond autocorrelation test may be unreliable*." (Roodman (2009), p. 128)

[14] Bolivia, Brazil, Chile, Colombia, Costa Rica, Dominican Republic, Ecuador, El Salvador, Guatemala, Honduras, Mexico, Paraguay, Peru, Uruguay; we had to drop Argentina and Panama because of data inconsistencies

[15] As explained by Sari and Soytas (2007) in footnote 2: "The aggregate capital stock at the end of time t is given by $K_t = (1-\delta)K_{t-1} + I_t$ which is called perpetual inventory method (Jacob et al., 1997), where I_t is the investment at time t and δ is depreciation rate. Assuming that the depreciation rate is constant, the variance in capital is mostly related to the change in investment. Therefore, following Jin and Yu (1996) and Shan and Sun (1998) among others, we use growth of "gross capital formation" as a reliable proxy for growth of capital stock in our analyses."

levels and transform them in logarithmic terms, the shape parameters change and it appears that all variables follow a distribution that is pretty much alike a normal distribution.

Finally, Table 3 shows the correlation between pairs of our variables (expressed in growth rates). The vast majority of correlations are positive, apart from output-labour, energy-labour, exports-capital and exports-labour, and that energy consumption is particularly correlated with real GDP and real gross fixed capital formation.

Table 3: Correlations between the variables (expressed in growth rates)

	GDP	Energy	Capital	Labour	Imports	Exports
GDP	1.0000					
Energy	0.3941	1.0000				
Capital	0.7010	0.3406	1.0000			
Labour	-0.0116	-0.0307	0.0427	1.0000		
Imports	0.4071	0.1547	0.4694	0.0251	1.0000	
Exports	0.0113	0.0601	-0.0060	-0.0211	0.5540	1.0000

5. Empirical results[16]

Our empirical results, illustrated in Tables 4, 5 and 6, are divided into three sub-sections. First, we shed light on our results for the short-run dynamics in the export specification. This is followed by an analysis of the short-run dynamics for the equations with imports. Finally, this section concludes with the long-run effects combined with output elasticities[17].

5.1. Short-run dynamics exports

From Table 4, we can draw several conclusions concerning the presence of causality between our three main variables of interest. First, a bidirectional relationship between energy use and economic growth is observed in the short term, whereas in the long run only a unidirectional causality from energy consumption to real GDP exists.

Secondly, our short-run Granger causality test provides evidence of a one-way causation running from energy use to exports and that a causal tie between exports and output exists. However, the direction and extent of the latter depend on the applied GMM estimator. Put differently, while the system GMM estimator suggests that the causality runs from exports to output, the difference GMM estimator finds a feedback relation between these two variables. Nevertheless, both estimators indicate that some indirect causation runs from energy to output through exports in the short run.

With regard to the error correction term, Table 4 shows that this term is only significant in the output model, suggesting that short-run movements in this specification can be explained by the necessity to return to the long-run output equilibrium.

[16] It should be emphasised that according to Elliott (1998) our long-run inferences should be treated with caution due to the econometrical approach followed and its assumptions.
[17] To ensure comparability, we applied our methodology on Sadorsky's (2012) sample and our results are partially in line with his findings. However, differences exist with regard to the long-run causations running from output to energy consumption, exports and imports.

Table 4: Panel causality results for Latin America (export specification)

Origin of causation (independent variables)	Effect (dependent variable)									
	Δy		Δk		Δl		Δe		Δx	
	Diff. GMM	Sys. GMM	Diff. GMM	Sys. GMM	Diff. GMM	Sys. GMM	Diff. GMM	Sys. GMM	Diff. GMM	Sys. GMM
Short-run dynamics										
Δy			2.5650***	2.8370***	-0.0555*	-0.0781***	0.7467***	0.7717***	-0.1364**	0.0363
			(0.0000)	(0.0000)	(0.0656)	(0.0022)	(0.0000)	(0.0000)	(0.0474)	(0.1431)
Δk	0.1048***	0.1076***			0.0167	0.0249**	0.0520	0.0607	-0.1177	-0.1418
	(0.0000)	(0.0000)			(0.1090)	(0.0128)	(0.2421)	(0.3009)	(0.8599)	(0.7968)
Δl	-0.0505	-0.0988	0.5599	0.7101*			-0.2315	-0.2160	-1.2132	-1.3559
	(0.7955)	(0.6260)	(0.2578)	(0.0869)			(0.1079)	(0.2235)	(0.1846)	(0.1123)
Δe	0.1263**	0.1443	0.1378	0.1488*	-0.0180	-0.0249			0.4297**	0.5075**
	(0.0240)	(0.1174)	(0.1346)	(0.0587)	(0.4103)	(0.3600)			(0.0440)	(0.0112)
Δx	0.0254***	0.0287***	-0.0431	-0.0435	0.0022	0.0033*	0.0067	0.0067		
	(0.0001)	(0.0034)	(0.5112)	(0.3593)	(0.2115)	(0.0661)	(0.5681)	(0.5416)		
Long-run dynamics										
lag ect	-0.0833*	-0.0129**	0.3058	0.0170	0.0057	0.0024	0.1361	0.0277	0.2443	0.0497
	(0.0650)	(0.0290)	(0.1330)	(0.5660)	(0.7610)	(0.7850)	(0.1490)	(0.1050)	(0.1850)	(0.4760)
Sargan test (p-value)	0.4208	0.1977	0.4489	0.0614	0.2288	0.0837	0.1671	0.0390	0.5716	0.4007
AR1 (p-value)	0.0015	0.0016	0.0018	0.0017	0.0020	0.0020	0.0080	0.0077	0.0037	0.0040
AR2 (p-value)	0.2716	0.3279	0.0393	0.0989	0.7068	0.5391	0.3128	0.3566	0.9025	0.8458
Observations	405	420	405	420	405	420	405	420	405	420

Notes: These results were obtained using difference and system GMM estimators with robust standard errors and a determined lag length of 2. In the short-run dynamics section of this table, the upper case cell reflects the sum of the lagged coefficients for the corresponding short-term changes, while the lower case cell denotes the probability values (Chi-square) of the F-test. In the long-run dynamics sections, the coefficient of the error correction term is represented in the upper case cell and the lower case cell denotes the obtained probability values of the respective t-test. The Sargan test is calculated based on the estimation with GMM standard errors. Significance at the 1%, 5% and 10% levels are denoted by ***, ** and * respectively.

Since we use difference and system GMM estimators, we also need to test the validity and exogeneity of the instruments used. To do so, we relied on the Sargan test for over-identification of restrictions and the Arellano-Bond serial correlation test. The results are on the bottom part of Table 4.

The Arellano-Bond test indicates that the null-hypothesis of no autocorrelation at order 2 cannot be rejected, implying that the instruments used are valid. However, the capital equation forms an exception by highlighting the presence of autocorrelation. In addition, Sargan's null hypothesis of over-identification of restrictions cannot be rejected at any generally used level of significance, with the exception of the capital, labour and energy equations estimated with the system GMM estimator. However, as the main obtained result (i.e. the causality between energy use and GDP) coincide regardless of the estimator used, it can be argued that the found causality is valid, whereas the others (e.g. the link between capital and energy) need to be interpreted with more caution.

In this respect, it is worth stressing that some discrepancies between the two estimators used exist (e.g. the extent of causation between exports and output) and as the system GMM estimators' instruments are not always exogenous and valid, we decided that our policy implications should only be based on the difference GMM estimators' results.

Finally, we have carried out a robustness check consisting of adding Argentina and Panama to our database (cf. appendix). The robustness check confirms the previous findings with only some minor deviations (e.g. the existence of a long-run unidirectional link from energy use to exports)[18].

5.2. Short-run dynamics imports

Our empirical findings in Table 5 illustrate that a feedback relation between energy and output and imports and output[19] exists in the short term. In the long term, however, these two-way causalities must give way to unidirectional links running from energy use to GDP and imports to GDP, given that the lagged error correction term is only significant in the output equation. The presence of long-run dynamics in the output equation is in line with the previously obtained results in the equation with exports and the magnitude of the adjustment rates (i.e. absolute value of the error correction term) resembles between these two specifications.

In addition, it can be maintained that energy consumption also exerts an indirect influence on economic output via real gross fixed capital formation. This can be explained by the one-way causation running from energy use to capital combined with a feedback relationship between real GDP and capital. However, according to the short-run Granger causality test, no proof of causation can be found between energy use and imports at any generally used level of significance.

With regard to the two computed tests to see whether the applied instruments are exogenous and valid, the bottom part of Table 5 reveals that no autocorrelation at order 2 exists, implying that the used instruments are valid. Furthermore, the Sargan test for over-identification of restrictions shows that, with the exception of the capital, labour and energy equations estimated with the system GMM estimator, our model and instruments are valid. However, as both estimators do not provide the same results, we decided not to consider the system GMM estimators' evidence of causations when discussing the policy implications of our results.

Finally, as for the export specification, our robustness check indicates only slight discrepancies from our base model (e.g. the existence of a feedback relation between imports and real GDP, in both the short and long run)[20].

[18] Refer to the appendix for more information on our robustness check
[19] The bidirectional relation between imports and output is only proven by the difference GMM estimator.
[20] Refer to the appendix for more information on our robustness check

Table 5: Panel causality results for Latin America (import specification)

Origin of causation (independent variables)	Effect (dependent variable)									
	Δy		Δk		Δl		Δe		Δm	
	Diff. GMM	Sys. GMM	Diff. GMM	Sys. GMM	Diff. GMM	Sys. GMM	Diff. GMM	Sys. GMM	Diff. GMM	Sys. GMM
Short-run dynamics										
Δy			1.9947***	2.2200***	-0.0482*	-0.0718***	0.7577***	0.7981***	0.7800*	0.7614
			(0.0000)	(0.0000)	(0.0886)	(0.0079)	(0.0000)	(0.0000)	(0.0766)	(0.1762)
Δk	0.0851***	0.0875***			0.0145	0.0231**	0.0768	0.0804	0.6248***	0.6087***
	(0.0000)	(0.0000)			(0.1231)	(0.0113)	(0.1690)	(0.1624)	(0.0000)	(0.0000)
Δl	-0.0459	-0.0833	0.6638	0.8532*			-0.2683*	-0.2556	-0.6525	-0.7486
	(0.8042)	(0.6486)	(0.2151)	(0.0999)			(0.0728)	(0.1161)	(0.2873)	(0.1384)
Δe	0.1379**	0.1566*	0.1308*	0.1358***	-0.0198	-0.0250			-0.1292	-0.0627
	(0.0169)	(0.0760)	(0.0681)	(0.0075)	(0.3703)	(0.3100)			(0.4916)	(0.7428)
Δm	0.0337**	0.0377**	0.2875***	0.2846***	0.0035	0.0040	-0.0447	-0.0404		
	(0.0391)	(0.0254)	(0.0000)	(0.0000)	(0.7882)	(0.8814)	(0.4060)	(0.4451)		
Long-run dynamics										
lag ect	-0.0833*	-0.0137***	0.2429	0.0177	0.0017	0.0004	0.1434	0.0271	0.0745	-0.0788
	(0.0550)	(0.0090)	(0.1120)	(0.5270)	(0.9270)	(0.9670)	(0.1550)	(0.1580)	(0.4830)	(0.2270)
Sargan test (p-value)	0.4468	0.2193	0.3504	0.0423	0.2053	0.0474	0.1493	0.0343	0.4453	0.3540
AR1 (p-value)	0.0015	0.0006	0.0015	0.0010	0.0018	0.0017	0.0081	0.0077	0.0007	0.0008
AR2 (p-value)	0.2834	0.8976	0.3657	0.4857	0.9263	0.9321	0.4129	0.4745	0.8333	0.6448
Observations	405	420	405	420	405	420	405	420	405	420

Notes: These results were obtained using difference and system GMM estimators with robust standard errors and a determined lag length of 2. In the short-run dynamics section of this table, the upper case cell reflects the sum of the lagged coefficients for the corresponding short-term changes, while the lower case cell denotes the probability values (Chi-square) of the F-test. In the long-run dynamics sections, the coefficient of the error correction term is represented in the upper case cell and the lower case cell denotes the obtained probability values of the respective t-test. The Sargan test is calculated based on the estimation with GMM standard errors. Significance at the 1%, 5% and 10% levels are denoted by ***, ** and * respectively.

5.3. Long-run elasticities

The long-run elasticities of the growth regression are obtained by estimating equation (2) with a between-dimension FMOLS estimator (Pedroni, 2001a). Table 6 shows that all variables are significant at 1% and that real GDP is positively affected by the other four variables in both specifications.

In the export model, real GDP raises by around 0.17%, 0.57%, 0.33% and 0.04% when capital, labour, energy use and exports increases by 1% respectively. With regard to the output equation with imports, Table 6 illustrates that, in the long term, economic output alters after a 1% change in capital, labour, energy consumption and imports by around 0.14%, 0.57%, 0.35% and 0.04% respectively.

Table 6: Long-run output elasticities for the equation with exports and imports (FMOLS approach)

Dependent variable: log output	Exports	Imports
k	0.1661***	0.1441***
	(0.0118)	(0.0153)
l	0.5662***	0.5712***
	(0.0288)	(0.0367)
e	0.3267***	0.3512***
	(0.0285)	(0.0299)
t	0.0414***	0.0413***
	(0.0072)	(0.0112)
Observations	450	450

Standard errors are given in parenthesis; *** denotes significance at 1% level; ** 5% level; * 10% level.
Since the software does not generate reliable R2 tests, they should not be used for extrapolation or other forecasting exercises.

The sign and magnitude of our results, in particular in terms of the elasticity of energy consumption and trade with respect to real GDP, are similar to those found by Sadorsky (2012) for a smaller sample of South American countries. The main differences are that the role played by labour differs and that in our equation with imports the impact of energy use on economic output is smaller by 0.05%. In comparison with Apergis and Payne's (2009 and 2010) FMOLS results, our estimates of energy consumption are somewhat different. They are lower than theirs for South America (0.33% in the export specification and 0.35% in the import specification compared to 0.42%. However, they are higher than what they find for Central American economies (0.33% (in the export specification) or 0.35% (in the import specification) instead of 0.28%.[21] All this emphasizes the importance of the specific characteristics of the sample analysed.

6. Policy implications

The main policy implication resulting from our analysis is that Latin America's economic growth and export behaviour is sensitive to its energy and environmental policy developments. In other words, energy conservation policies are potentially at odds with measures aiming at fostering economic growth and exports in Latin America. For instance, our results establish the short-term causations running from energy consumption to economic output and exports. The former is also valid in the

[21] This result might be explained by the fact that our sample consists of Latin American countries and that discrepancies in energy demand and supply among Central and South American countries exist. For instance, it can be stressed that South American countries accounted for nearly 70% of the region's energy production in 2010 and that the main energy producers were Brazil and Venezuela (246,629.61 and 198,568.41 ktoe in 2010 respectively) (World Bank, 2013).

long run. This strengthens the case to emphasize the adoption of new technologies rather than changes in demand in the design of energy policies in the short run. If demand management is needed, it is likely to have a cost in terms of growth and trade. Policies aiming at expanding renewable energy sources, improving energy efficiency and fostering R&D in energy-saving applications are the obvious avenues as already pointed out by many of the other researchers working on Latin America.

With regard to the energy-import nexus, we cannot find any proof of causation between these two variables, implying that a policy modification in one area does not influence the other. Nevertheless, our results indicate the existence of a unidirectional link from imports to economic output in the long term, whereas a feedback relationship between these two variables is observed in the short run. The latter does not only advocate that import restrictions impede economic performance, but also that cutbacks in economic growth negatively affect imports.

Moreover, the long-run relationship is in line with the import-led growth hypothesis, which suggests that programs seeking to foster imports exert a positive influence on economic growth. The reason for this causal tie can be explained by the fact that imports do not only enable a transfer of knowledge and technology, but also help a country to obtain required intermediate production factors. The former is of particular importance for developing countries, as it allows them to take advantage of previously invented technologies by leapfrogging them, which may enhance productivity and thus economic growth (Grossman and Helpman, 1991).

Finally, while we have found evidence in support of the one-way causality from exports to real GDP in the long run, a two-way causality between these two variables is observed in the short term. This implies that export restrictions hinder the economic growth expansion of a country. Therefore, policies promoting trade, both exports and imports, elevate economic growth. However, our results show that the role of imports in the economic growth path of Latin American countries is slightly larger than of exports, suggesting that export promotions combined with retained import restrictions are less efficient.

7. Conclusion

The main purpose of the paper was to get a broader sense of the robustness of some of the 3-way causality assessments of the dynamics of growth, trade and energy consumption in Latin America. Despite the usual data challenges, and thanks to an approach that had not been used so far in this context for Latin America and thanks also to a broader dataset, we have been able to validate three of the results generated by Sadorsky (2012) for the region, but we ended up rejecting one. More specifically, we reached the following four insights.

First, the key policy relevant result on the one-way causality from energy consumption to output (in the long run) and exports (in the short run) reached by Sadorsky (2012) is validated both by the alternative approach and the larger sample. This implies that policies tackling rising energy consumption could interfere with the economic growth and export paths of Latin American economies.

Second, also in line with Sadorsky (2012), we find a one-way causation from our trade variables (exports and imports) to real GDP, suggesting that trade restrictions slightly impede economic performance. Therefore, an increase in protectionism in Latin America can lessen economic growth albeit to a lesser degree than a reduction in energy consumption.

Third, the magnitude of our estimated long-run elasticities of energy consumption and trade with respect to real GDP resemble those found by Sadorsky (2012) for South American economies. We find differences (e.g. the smaller impact of energy use on economic output in the equation with imports) but these do not really influence the main results, not, most importantly, their policy implications.

Finally, in contrast to Sadorsky (2012), who argued that a long-run feedback relation between energy use, economic output and trade (reflecting either exports or imports) exists, we only find a one-way causation from energy consumption, exports and imports to economic growth in the long run.

Overall, the main point may be that policies aiming at reducing energy consumption under current production technologies will influence growth and trade. Moreover, policies restricting energy consumptions have had, in the last 30 years, an even stronger impact on economic performance than policies aiming at restricting trade. However, the long-term effects of an economic turmoil or trade restrictions on energy consumption are still ambiguous since different samples and econometric techniques lead to different conclusions.

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Appendix

A1: Tables

Table A.1: Panel unit root tests

Method	y		Δy		e		Δe	
	Statistic	Prob.	Statistic	Prob.	Statistic	Prob.	Statistic	Prob.
Null hypothesis: Unit root (assumes common unit root process)								
Levin, Lin & Chu t*	2.7037	0.9966	-8.8686	0.0000	1.3493	0.9114	-5.2461	0.0000
Null hypothesis: Unit root (assumes individual unit root process)								
Im, Pesaran and Shin W-stat	7.4189	1.0000	-9.8276	0.0000	5.7202	1.0000	-9.1051	0.0000
ADF - Fisher Chi-square	1.7735	1.0000	152.6510	0.0000	4.5058	1.0000	141.7700	0.0000
PP - Fisher Chi-square	0.9245	1.0000	167.8620	0.0000	3.9858	1.0000	276.6420	0.0000
Pesaran's CADF test	0.4940	0.6890	-4.2280	0.0000	0.2090	0.5830	-3.5910	0.0000

Method	k		Δk		l		Δl	
	Statistic	Prob.	Statistic	Prob.	Statistic	Prob.	Statistic	Prob.
Null hypothesis: Unit root (assumes common unit root process)								
Levin, Lin & Chu t*	1.0466	0.8523	-11.5164	0.0000	-3.2881	0.0005	-6.1148	0.0000
Null hypothesis: Unit root (assumes individual unit root process)								
Im, Pesaran and Shin W-stat	3.4245	0.9997	-11.2277	0.0000	2.2815	0.9887	-6.5352	0.0000
ADF - Fisher Chi-square	10.5266	0.9996	175.3270	0.0000	19.4266	0.9306	99.1977	0.0000
PP - Fisher Chi-square	5.5472	1.0000	224.3730	0.0000	32.8213	0.3304	185.7300	0.0000
Pesaran's CADF test	-0.9190	0.1790	-5.1420	0.0000	0.7930	0.7860	-3.4980	0.0000

Method	m		Δm		x		Δx	
	Statistic	Prob.	Statistic	Prob.	Statistic	Prob.	Statistic	Prob.
Null hypothesis: Unit root (assumes common unit root process)								
Levin, Lin & Chu t*	1.9384	0.9737	-7.7103	0.0000	0.2169	0.5859	-8.1712	0.0000
Null hypothesis: Unit root (assumes individual unit root process)								
Im, Pesaran and Shin W-stat	5.7697	1.0000	-9.8761	0.0000	3.8462	0.9999	-10.1390	0.0000
ADF - Fisher Chi-square	4.1158	1.0000	151.9990	0.0000	7.1254	1.0000	157.7310	0.0000
PP - Fisher Chi-square	4.4200	1.0000	310.7570	0.0000	9.4758	0.9999	300.3350	0.0000
Pesaran's CADF test	-0.9380	0.1740	-5.1600	0.0000	-0.2990	0.3830	-4.3290	0.0000

Notes: We conducted the panel unit root tests with constant and all tests used a lag length determined by the Schwarz Information Criterion, with the exception of Pesaran's CADF were the estimated lag is 2. Probabilities for the Fisher tests are computed using an asymptotic Chi-square distribution. All other tests assume asymptotic normality (EViews 8, 2013).

Table A.2: Descriptive statistics (in growth rates)

	e	y	k	l	x	m
Mean	2.9539	3.1066	4.5168	2.8015	9.0929	8.6873
Maximum	43.0798	18.2866	100.6472	14.2110	215.6925	151.6888
Minimum	-22.3314	-11.8000	-46.4312	-4.4384	-52.8458	-53.2083
Std. Dev.	6.1920	3.9748	16.0363	1.7629	24.9626	19.9872
Skewness	0.6993	-0.7578	0.8691	0.9224	2.9567	1.3684
Kurtosis	9.0174	4.9390	8.4363	8.7659	21.9530	10.5826
Observations	450	450	450	450	450	450

Table A.3: Descriptive statistics (in natural logarithms)

	e	y	k	l	x	m
Mean	9.1931	24.3222	22.5530	15.4429	22.3625	22.4780
Maximum	12.4908	27.7234	26.1432	18.4387	26.4317	26.4708
Minimum	7.0305	22.1569	19.9052	13.6122	18.6129	19.2221
Std. Dev.	1.4051	1.5055	1.5691	1.1524	1.5520	1.4207
Skewness	0.7940	0.6947	0.6654	0.8925	0.3728	0.3656
Kurtosis	2.4661	2.4807	2.5485	2.9935	2.7795	3.0800
Observations	465	465	465	465	465	465

A2. Robustness check

Our data section highlighted that two countries in our sample, namely Argentina and Panama, might influence the robustness of our findings due to data misspecification or colossally high trade growth rates and they were therefore omitted in our base model. However, for the sake of completeness and comparison, we decided to carry out our analysis again by adding Argentina and Panama to our database.

For our robustness check, the results of the unit root tests and Kao's panel cointegration test do not alter from our base specifications, indicating that non-stationarity exists in levels, but not in first differences and that our variables are cointegrated. However, as regards the direction and existence of causality among the variables some discrepancies between our base model and the robustness check exist.

First, our export and import specifications show that, in contrast to our base model, an indirect causality runs from real GDP to energy consumption through capital and that the error correction term is not only significant in the output equation, but also in the capital and trade equations[22] of our model. Therefore, the import specification also suggests that a feedback relation between imports and real GDP exists, in both the short and long run.

Secondly, while a unidirectional relation from energy use to exports is also confirmed in the long run, both estimators indicate the absence of a causal link between exports and output when including Argentina and Panama in our database.

Thirdly, it is worth stressing that the Sargan test for over-identification of restrictions and the Arellano-Bond test for serial correlation indicate that the used instruments in our robustness check are valid and exogenous, even for some of the equations that caused troubles in the base model.

Overall, our robustness check reveals only minor deviations from our base model when adding Argentina and Panama. We can therefore argue that the major inference drawn with regard to the policy implications of our findings (i.e. energy conservation policies are potentially at odds with policies aiming at fostering economic growth and exports in Latin America) remains accurate.

[22] The error correction term in the export and import equations is only significant in the difference GMM estimator approach.