# Do urbanization, income, and trade affect electricity consumption across Chinese provinces?

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ABSTRACT

The aim of this paper is to investigate the short- and long-run links among urbanization, output (Gross Domestic Product, GDP), trade openness, and electricity consumption in China, using a rich dataset at the provincial level. The short-run Granger causality analysis discloses a unidirectional causal relationship running from electricity to output and weak feedback effects between trade and urbanization. The long-run Granger causality analysis shows output, urbanization, and trade trigger electricity consumption whereas trade, urbanization, and electricity cause output. The Group Mean and Lambda-Pearson causality tests reveal a large heterogeneity in the long-run effects which suggests there is no "one-size fits all" policy and each region should formulate a differentiated urbanization/growth strategy based on its own characteristics to control electricity utilization.

## 1. Introduction

Urbanization and industrialization have always been two of the main engines of economic growth and the past three decades have witnessed a tremendous development of the Chinese economy with increasing industrial concentration and migration from rural areas into cities. With the implementation of policy reforms, after the famous speech by Deng Xiaoping at the plenum of the Communist Party on 22 December 1978, China entered a period of rapid economic and social transformations. Within the new political environment, the utilization of foreign capital and external technological cooperation in all forms became acceptable. Provincial and local governments received increasing authority and the Communist Party allowed local experiments with new forms of organization of production such as the introduction of the household responsibility system which soon replaced the people's communes (Yao, 2018). China gradually relaxed from central planning, opened up for trade, and rapidly integrated into the world economy (Chow, 2018; Garnaut, 2018). Price liberalization, diversified ownership and individual property rights created a booming market-based system (Fang et al., 2018). According to IMF, China is currently the largest economy in the world in terms of GDP in PPP as well as the most important exporter in current US$. However, discrepancies across and within regions are still significant with wealthy industrial areas along the east coast and provinces which are still developing in the western part of the nation.

Double-digit economic growth was supported by large-scale migration from the countryside and the urbanization rate rose from 17.9% in 1978 to 58.8% in 2017 (National Bureau of Statistics of China, 1999–2019). This figure is currently close to the world average, but still below that of developed countries, where the proportion of the urban population is between 70 and 90%. Nonetheless, China now has the largest urban population as it accounts for about 20% of the world's total and it is projected to exceed 1 billion people in the next two decades (China Urban Research Committee, 2008).

Such a dramatic expansion in output and urbanization came at several costs yet. First, about half of urban growth has been at the expense of arable land and often rural areas have been expropriated forcing people into newly developed zones. Second, urbanization is linked to living conditions. For instance, the household registration system prevents migrant peasant workers from integrating into urban communities and forbids access to urban medical insurance and other services. However, wage differentials push people to move into cities since urban residential income has been consistently three times larger than rural income. Income disparity is the lowest in the prosperous eastern regions, whereas it is the largest in the relatively less developed western area (Dong and Hao, 2018). Third, a large share of Chinese cities suffers some energy and water shortages, as these basic resources are in high demand for the promotion of urbanization (Wang et al., 2018), exacerbated by severe surface and ground pollution (Bai et al., 2014). Actually, China has become the greatest emitter of $CO_2$ in the world and the

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largest energy consumer. In 2016, China accounted for 21.5% of the world energy consumption, while the United States for 15.7% (IEA, 2018).

Electricity utilization also dramatically escalated as, according to the Chinese Energy Statistical Yearbook, per capita consumption increased by over four-and-a-half times in the last twenty years. In 2017 the IEA estimates electricity consumption of 5899 TW in China vs 4077 TW in USA, but Chinese per capita electricity consumption is only one third of the US one (IEA, 2018). With Chinese lifestyles closing the gap with Western ones, there has been concern over the effects on the environment in the long-run as Chinese electricity supply is still dominated by coal. In 2016 thermal power accounted for about 66% of the electricity generation mix. Hence, it is crucial to gather a better understanding about the relationship among urbanization, economic growth and electricity consumption for leading a sustainable development. Researchers often focused on the electricity-GDP nexus neglecting urbanization yet (Herrerias et al., 2013; Zhong et al., 2019). To the best of our knowledge, no previous study addresses the role of economic growth, urbanization, and international trade on electricity consumption in an integrated framework at the regional level. To assess the relationships among these variables we make use of a balanced panel of 28 provinces during the period 1995–2016.

Our work makes several contributions to the existing literature. First, we pioneer the analysis of the growth-urbanization-electricity-trade nexus in Chinese provinces. Second, we employ a robust econometric approach to tackle the presence of cross-sectional dependence (CSD). Third, we address both the short- and long-causal relationships between the above-mentioned variables. The former is achieved using the Granger causality test devised by Dumitrescu and Hurlin (2012). The latter by a two-step approach proposed by Canning and Pedroni (2008) which takes into account CSD (Eberhardt and Teal, 2013).

The main findings of the study are as follows. First, Pesaran's CD test and the exponent of CSD introduced by Bailey et al. (2016) confirm the presence of strong cross-sectional dependence. The PANIC and PANICCA approaches suggest the variables under scrutiny are integrated of order one. The short-run causality tests show electricity does cause output, while urbanization appears to be caused by trade and GDP. Notably, only urbanization and trade seem to display network effects, but only at a 10% significance level. In the long-run, causality runs from GDP, urbanization and trade openness to electricity consumption as well as from trade, urbanization, and electricity consumption to GDP.

The rest of the paper is organized as follows. The following section summarizes the empirical literature on the electricity-growth and electricity-urbanization nexuses. Section 3 describes the methodology, while section 3 discusses the results. The last section concludes and addresses policy implications.

## 2. Literature review

An extensive body of applied studies has explored the causal relationships between energy and GDP. A comprehensive review is provided by Menegaki and Tsani (2018), while Tsani and Menegaki (2018) discuss issues and advancements associated with new methodologies employed with the aim to identify the missing links in the energy-income puzzle. The investigation about the direction of causality has obvious implications for environmental conservation policies and development strategies. The growth hypothesis assumes that unidirectional causality runs from energy to income, whereas the opposite direction holds for the conservation hypothesis. The former asserts energy is an important intermediate input and disruptions in energy supply or policies aimed at protecting the environment may have an adverse impact on economic growth. On the contrary, if the conservation hypothesis holds, energy conservation policies can be implemented without deteriorating the economic performance. The neutrality hypothesis states there is no relationship between output and electricity consumption, whilst the feedback hypothesis posits a bi-directional link amid them. Under the former, electricity conservation efforts are not likely to thwart economic growth, whereas the latter hints to complex interdependent relationships that must be investigated with more complete models.

It is hard to convince that studies have reached a consensus. For instance, according to Narayan and Prasad (2008), energy consumption and economic growth are unrelated in most of selected developed countries. Only in 8 out of 30 OECD nations causality runs from energy consumption to economic growth. This causal relationship is also investigated by Chontanawat et al. (2008) who add 78 non-OECD countries. In contrast to the previous finding, they discover that the growth hypothesis holds for 70% of the selected OECD countries, while only 46% of the selected non-OECD countries supports this evidence. Hence, they claim that conservation policies are valid in most of the high-developed countries and only in a third of the low-income countries.

Table 1 presents a summary of recent empirical studies on the electricity-growth nexus where the four alternative cases are investigated. Even if we focus on developing countries, we can hardly find an unambiguous picture. Several single country studies favor the growth hypothesis, mostly in the short-run, but the feedback view often holds in the long-run. Some authors address similar economies with very different results. For instance, Squalli (2007) investigates 11 OPEC member countries from 1980 to 2003 and finds GDP depends on electricity in Indonesia, Iran, Nigeria, Qatar, and Venezuela, with evidence of feedback effects in Iran and Qatar. Surprisingly, Algeria, Iraq and Libya appear to be unrelated with electricity consumption, while energy conservation can have little to no impact on economic growth in the most advanced economies of the Gulf, that is Kuwait, Saudi Arabia, and the UAE. Ozturk and Acaravci (2011) consider 11 MENA countries over the period 1971–2006, but they are obliged to drop seven of them as unit root tests indicate that some of the variables do not satisfy the underlying assumptions of the ARDL bounds testing approach or there is no cointegration between electricity consumption and growth. Finally, they can establish a short-run Granger causality from real GDP to electricity in Israel and Oman, whereas in the long-run the growth hypothesis holds in Oman and the conservation hypothesis in Egypt and Saudi Arabia. Apergis and Payne (2011) address 88 countries with a smaller time span (1990–2006). They split results according to the level of development and find bidirectional causality in both the short- and long-run for the high-income and upper-middle income countries. On the contrary, the growth hypothesis holds for the low-income country panel, while the lower-middle income nations exhibit unidirectional causality from electricity consumption to economic growth in the short-run, but bidirectional causality in the long-run.

In a recent study, Karanfil and Li (2015) use per capita electricity consumption and per capita GDP in a panel of 160 countries, which is divided into various subsamples based on countries' income levels, regional locations and OECD memberships. These authors claim the feedback hypothesis holds in the full sample and the majority of the subsamples. Yet, in the short-run, causality is running from economic growth to electricity consumption in lower- and middle-income countries as well as in East Asia, Middle East, North Africa and in the Pacific area. On the contrary, North America, Sub-Saharan Africa and upper-middle income countries exhibit evidence of neutrality, whilst the growth hypothesis is strongly rejected. Karanfil and Li (2015) also go beyond a simple GDP-electricity model and introduce urbanization as an additional explanatory variable.

Other papers explore the channels which influence the energy-urbanization nexus (Holtedahl and Joutz, 2004; Halicioglu, 2007; Liu, 2009; Mishra et al., 2009; Apergis and Tang, 2013; Liddle, 2013, Sbia et al., 2017). First, urbanization modifies energy use as it allows for economies of scale while promoting modernization and expansion of energy-intensive industries. Second, it may change consumer behavior with a likely shift towards energy-intensive goods and services even if

**Table 1**
Summary of findings in selected causality studies.

| Author | Country/Region | Period | Causality |
|---|---|---|---|
| Shiu and Lam (2004) | China | 1971–2000 | ELC → Y (SR&LR) |
| Narayan and Smyth, 2005 | Australia | 1966–1999 | Y → ELC (SR&LR) |
| Squalli, 2007 | 11 OPEC countries | 1980–2003 | LR: ELC → Y (Indonesia, Nigeria, Venezuela); Y → E (Kuwait, Saudi Arabia); E↔Y (Iran, Qatar) |
| Chen et al. (2007) | 10 Asian countries | 1971–2001 | LR: ELC–Y (India, Singapore, Taiwan, Thailand); ELC → Y (Indonesia);Y → ELC (Hong Kong, Korea) |
| Yuan et al. (2007) | China | 1978–2004 | ELC → Y (SR); ELC ↔ Y (LR) |
| Halicioglu (2007) | Turkey | 1968–2005 | Y,P,UR → ELC (LR); Y,P → ELC (SR) |
| Yuan et al. (2008) | China | 1963–2005 | ELC → Y (SR); ELC ↔ Y (LR) |
| Xie et al. (2009) | Shangai | 1978–2010 | SR: EC ↔ UR / LR: EC,UR → Y; Y → UR; UR → EC |
| Ghosh, 2009 | India | 1970–2006 | Y → ELP (SR) |
| Liu (2009) | China | 1978–2008 | EC → UR (SR&LR); Y,LF,UR → EC (LR) |
| Mishra et al. (2009) | Pacific Island countries | 1980–2005 | UR → EC (SR); ELC,UR → Y;Y,UR → ELC (LR) |
| Akinlo, 2009 | Nigeria | 1980–2006 | ELC → Y (SR&LR) |
| Yoo and Kwak, 2010 | 7 American countries | 1975–2006 | SR: ELC → Y (Argentina, Brazil, Chile, Colombia and Ecuador); ELC↔Y (Venezuela); ELC–Y (Peru) |
| Chandran et al. (2010) | Malaysia | 1971–2003 | ELC → Y (SR&LR) |
| Ozturk and Acaravci, 2011 | 11 MENA countries | 1971–2006 | SR: Y → ELC (Israel, Oman); LR: Y → ELC (Oman); ELC → Y (Egypt, Saudi Arabia) |
| Kouakou, 2011 | Cote d'Ivoire | 1971–2008 | ELC ↔ Y (SR); ELC → Y (LR) |
| Lai et al. (2011) | Macao | 1999–2008 | Y – ELC (SR); Y → ELC (LR) |
| Apergis and Payne, 2011 | 88 countries | 1990–2006 | ELC ↔ Y (high and upper-middle income); ELC → Y (low income) |
| Alam et al. (2012) | Bangladesh | 1972–2006 | ELC ↔ Y (LR); ELC–Y(SR&LR) |
| Apergis and Tang, 2013 | 88 countries | 1990–2007 | EC, LF, UR → Y in 16,4% of the sample |
| Herrerias et al. (2013) | China | 1995–2009 | ELC ↔ Y (LR full sample); ELC ↔ Y (SR 1999–2009) Y → ELC (LR 1999–2009) |
| Cheng et al. (2013) | China | 1953–2010 | ELP → Y (SR) |
| Polemis and Dagoumas (2013) | Greece | 1970–2011 | ELC ↔ Y |
| Solarin and Shahbaz, 2013 | Angola | 1971–2009 | LR: ELC ↔ Y; ELC ↔ UR |
| Saunoris and Sheridan, 2013 | United States | 1970–2009 | ELC → Y (SR); Y → ELC (LR) |
| Shahbaz et al., 2014 | United Arab Emirates | 1971–2008 | ELC ↔ Y (SR); Y → ELC (LR) |
| Cowan et al. (2014) | BRICS | 1990–2010 | ELC ↔ Y (Russia) Y → ELC (South Africa); ELC – Y (Brazil, India and China) |
| Hamdi et al. (2014) | Bahrain | 1980–2010 | ELC ↔ Y (SR&LR) |
| Liddle and Lung (2014) | 105 countries | 1971–2009 | ELC → UR (LR) |
| Karanfil and Li (2015) | 160 countries | 1980–2010 | SR: mostly ELC ↔ Y; LR: |
| Iyke, 2015 | Nigeria | 1971–2011 | ELC → Y (SR&LR) |
| Acaravcı et al. (2015) | Turkey | 1974–2013 | ELC → Y (SR&LR) |
| Zhao and Wang (2015) | China | 1980–2012 | SR: EC ↔ Y; UR → EC; Y → UR |
| Faisal et al. (2016) | Russia | 1990–2011 | ELC ↔ Y |
| Marques et al. (2016) | Greece | 2004–2014 | REC–I |
| Raza et al. (2016) | 4 Asian countries | 1980–2010 | ELC → Y |
| Khobai and Le Roux, 2017 | South Africa | 1971–2013 | UR → ELC (LR) |
| Sbia et al. (2017) | United Arab Emirates | 1975–2011 | ELC → Y (SR); Y ↔ ELC; UR ↔ ELC (LR) |
| Faisal et al. (2018) | Iceland | 1965–2013 | ELC –Y (SR&LR); UR ↔ ELC (LR) |
| Ahmad and Zhao (2018) | China | 2000–2016 | UR → EC; EC ↔ Y (Country and 3 regional panels) |
| Zhong et al. (2019) | China | 1971–2009 | ELC → Y |

*Notes*: ELC stands for electricity consumption and EC stands for energy consumption, ELP for Electricity production or supply, Y for GDP or income, UR for urbanization, LF for labor force or population, P for energy/electricity price, SR for short-run, LR for long-run.
→ indicates the direction of one-way causality; ↔ indicates two-way causality; – indicates no causality.

concerned customers may prefer green products. Third, urbanization requires larger public infrastructures, driving up demand for energy-intensive materials. Liddle and Lung (2014) examine the electricity-urbanization nexus using data from 105 countries spanning 1971–2009. They state electricity consumption Granger causes urbanization, but they miss to analyze the links with economic growth. On the contrary, the multivariate model set up by Karanfil and Li (2015) allows these authors to claim that urbanization plays a significant role in the long-run. Higher per capita electricity consumption induces urbanization in some areas such as in Sub-Saharan Africa, whilst in other developing or emerging economies, the causality runs in the opposite direction. They conclude that the electricity-growth nexus is highly sensitive to regional differences, countries' income levels, and urbanization rates.

The impact of urbanization on energy use in China and its provinces has been extensively studied, but a limited amount of literature investigates the link with electricity consumption. (Xiaoping et al., 2009; Xie et al., 2009; Ma, 2015; Yan, 2015; Yang et al., 2019). Studies have provided conflicting results. On one side, urbanization is energy deepening at the national level (Jiang and Lin, 2012; Lin and Ouyang, 2014). On the other side, it doesn't affect energy consumption (Sheng et al., 2014; Lin and Du, 2015). Regional differences matter, but just a few papers have addressed this issue. Song and Zheng (2012) highlight the positive influence of urbanization upon the aggregate energy intensity at the provincial level, but this link is not statistically significant for Ma (2015), while Liu and Xie (2013) unveil a non-linear relationship between urbanization and energy intensity in a two-regime threshold vector error correction model. Yan (2015) finds urbanization significantly increases aggregate energy intensity as well as electricity and coal intensities. Moreover, Elliott et al. (2017) show the direct impact of urbanization on energy intensity is generally positive despite the indirect impact, measured through four different channels (construction, industrial upgrading, transportation and changing lifestyles), tends to be negative with a substantial heterogeneity across Chinese provinces. Finally, Ahmad and Zhao (2018) uncover network effects between energy consumption and economic growth for the country as well as for subnational panels (Eastern economic zone, Intermediate economic zone, and Western economic zone). Results are not clear-cut with respect to urbanization. The country panel and the Eastern economic zone display feedback effects, whereas a unidirectional positive linkage is found running from growth to urbanization for the Intermediate economic zone. In the Western economic zone links are different: positive causality from economic growth to urbanization and negative causality from urbanization to economic growth. Hence, regional differences seem to matter and deserve to be addressed at the provincial level.

## 3. Data and methodology

In this study, we analyze the short- and long-run linkages among electricity consumption, urbanization, GDP, and trade using data for 28 provinces during the period 1995–2016.[1] Electricity consumption and GDP in constant 2000 prices are in per capita terms (NBSC, 1999-2019). Urbanization is measured by the proportion of the urban population (in cities and towns) to the total population by province. According to demographic statistics from the National Bureau of Statistics of China, urban population refers to people living in a specific urban area for more than six months within a statistical year. Trade openness is given by the usual Trade Openness Ratio (TOR) computed as nominal imports plus nominal exports relative to nominal GDP. We take the logs of all the previous variables so that estimated coefficients can be interpreted as elasticities.

First, we check whether the series under investigation are cross-sectionally dependent. To accomplish this task, we employ Pesaran's (2015) CD test together with the exponent of CSD introduced by Bailey et al. (2016). After confirming that series under consideration are cross-sectionally dependent, we rely on second generation unit root tests to identify the order of integration of the variables. We opt for the PANIC and PANICCA approaches (Bai and Ng, 2010; Reese and Westerlund, 2016), which allow to test for unit roots separately in the common factors and idiosyncratic components. Then, we assess the short-run causal relationships between the stationary variables making use of the Granger causality test developed by Dumitrescu and Hurlin (2012). This is a simplified version of the Granger (1969) noncausality test for heterogeneous panel data models with fixed coefficient. This test can be applied in the presence of CSD when all the variables in the panel are stationarity at a common level. The underlying model is:

$$y_{i,t} = \alpha_i + \sum_{k=1}^{K} \beta_{i,j} y_{i,t-k} + \sum_{k=0}^{K} \gamma_{i,j} x_{i,t-k} + \varepsilon_{i,t} \tag{1}$$

where $y_{i,t}$ and $x_{i,t}$ are observations of two stationary variables for region $i$ in period $t$. Coefficients can differ across regions, but are assumed to be time invariant, whereas the lag order $K$ is the same across provinces. The Schwartz Bayesian Information criterion is used to find the optimal lag length. The null and the alternative hypotheses are defined as follows:

$$H_0: \quad \gamma_{i,1} = \cdots = \gamma_{i,K} = 0 \qquad \forall i = 1, \ldots, N$$

$$H_1: \quad \gamma_{i,1} = \ldots = \gamma_{i,K} = 0 \qquad \forall i = 1, \ldots, N_1$$

$$\gamma_{1,j} \neq 0 \;\; \text{or} \ldots \text{or} \;\; \gamma_{1,j} \neq 0 \quad \forall i = N_1 + 1, \ldots, N$$

where $N_1$ is unknown but satisfies the condition $0 \leq N_1 \leq N$. We assume $N_1 \leq N$ otherwise there is no causality for any of the units in the panel, while there is causality for all the regions when $N_1 = 0$. If the null is accepted the variable $x$ does not Granger cause the variable $y$ for all the regions. By contrast, rejecting $H_0$ does not exclude that there is no causality for some provinces. This possibility is examined with the average of individual Wald statistics (Dumitrescu and Hurlin, 2012; Lopez and Weber, 2017).

[1] Data are from the China and Provincial Statistical Yearbooks compiled by the National Bureau of Statistics of China and published by China Statistical Press. These have been often updated and we mostly used data retrieved on January 2019 from http://www.stats.gov.cn/tjsj/ndsj. However, some older provincial data are only available on paper. In China there are 27 regions and 4 municipalities: Beijing, Shanghai, Tianjin and Chongqing. The latter was created in 1997 and has been included in Sichuan in the present study. Hainan and Tibet are excluded for lack of reliable information as in Herrerias et al. (2013).

Since both the Kao and Pedroni tests rejects the null of no cointegration[2] we apply the vector error correction model (VECM) version of panel Granger causality to detect the long-run direction of causal relationships among the variables under scrutiny (Apergis, 2018). This is a two-step process whose first step is to estimate the cointegrating relationship using Fully Modified Ordinary Least Squares (FMOLS). The one period lagged residual term, which comes from the potential long-run model, measures the speed of adjustment to equilibrium. In the second step, the lagged error term obtained from the first step is added to the following ECM system:

$$\Delta le_{i,t} = \delta_{1i} + \varphi_{1i}\hat{\varepsilon}_{i,t-1} + \sum_{j=1}^{K} \phi^{e}_{1i,j}\Delta le_{i,t-j} + \sum_{j=1}^{K} \phi^{u}_{1i,j}\Delta lu_{i,t-j} + \sum_{j=1}^{K} \mu^{g}_{1i,j}\Delta ly_{i,t-j} + \sum_{j=1}^{K} \mu^{o}_{1i,j}\Delta lt_{i,t-j} + u_{1i,t} \tag{2}$$

$$\Delta lu_{i,t} = \delta_{2i} + \varphi_{2i}\hat{\varepsilon}_{i,t-1} + \sum_{j=1}^{K} \phi^{e}_{2i,j}\Delta le_{i,t-j} + \sum_{j=1}^{K} \phi^{u}_{2i,j}\Delta lu_{i,t-j} + \sum_{j=1}^{K} \mu^{g}_{2i,j}\Delta ly_{i,t-j} + \sum_{j=1}^{K} \mu^{o}_{2i,j}\Delta lt_{i,t-j} + u_{2i,t} \tag{3}$$

$$\Delta ly_{i,t} = \delta_{3i} + \varphi_{3i}\hat{\varepsilon}_{i,t-1} + \sum_{j=1}^{K} \phi^{e}_{3i,j}\Delta le_{i,t-j} + \sum_{j=1}^{K} \phi^{u}_{3i,j}\Delta lu_{i,t-j} + \sum_{j=1}^{K} \mu^{g}_{3i,j}\Delta ly_{i,t-j} + \sum_{j=1}^{K} \mu^{o}_{3i,j}\Delta lt_{i,t-j} + u_{3i,t} \tag{4}$$

$$\Delta lt_{i,t} = \delta_{4i} + \varphi_{4i}\hat{\varepsilon}_{i,t-1} + \sum_{j=1}^{K} \phi^{e}_{4i,j}\Delta le_{i,t-j} + \sum_{j=1}^{K} \phi^{u}_{4i,j}\Delta lu_{i,t-j} + \sum_{j=1}^{K} \mu^{g}_{4i,j}\Delta ly_{i,t-j} + \sum_{j=1}^{K} \mu^{o}_{4i,j}\Delta ly_{i,t-j} + u_{4i,t} \tag{5}$$

We also include cross-section averages to allow for the presence of common factors as advocated by Eberhardt and Teal (2013). Within this setting all the variables in this ECM system are stationary and we can replace the error correction term with its estimate due to the superconsistency of the estimator of the cointegrating relationship. Then, we may perform standard hypothesis tests for the estimated coefficients. Yet, the reliability of individual tests is scarce due to the small sample size with only 22 years of data across 28 regions. Hence, we prefer to adopt to the two panel tests proposed by Canning and Pedroni (2008) and applied in the energy literature too (Narayan and Popp, 2012; Herrerias et al., 2013; Liddle and Lung, 2014; Mahalingam and Orman, 2018). These test statistics are the Group Mean (GM) based test and the Lambda–Pearson (LP) based test. The former takes the average of the provinces and has a standard normal distribution under the null hypothesis of no long-run causal effect for the panel. The latter is derived by the $p$-values associated with each province's $t$-test statistic. The Lambda–Pearson statistic is distributed as a $\chi^2$ with 2 $N$ degrees of freedom under the null of no long-run causation for the panel. Actually, the null hypothesis is the same for both the GM and LP tests under the assumption that coefficients are homogenous and equal to zero across regions, while the alternative states they are different from nil for some non-negligible portions of the Chinese provinces.

## 4. Findings

Cross-sectional units in our panel are likely interdependent as common shocks and/or national policies can affect regional electricity consumption, output, urbanization, and trade openness but with a different degree. We address weak versus strong dependence by making

[2] Results are available upon request.

**Table 2**
CD test and exponent of Cross Sectional Dependence.

| | CD-test | *p*-value | Mean and 90% confidence bands | | |
|---|---|---|---|---|---|
| | | | $\hat{\alpha}_{0.05}$ | $\hat{\alpha}$ | $\hat{\alpha}_{0.95}$ |
| *le* | 91.1 | 0.00 | 0.935 | 1.007 | 1.079 |
| *dle* | 70.6 | 0.00 | 0.961 | 0.997 | 1.032 |
| *ly* | 90.8 | 0.00 | 0.912 | 1.007 | 1.102 |
| *dly* | 86.2 | 0.00 | 0.960 | 1.007 | 1.055 |
| *lt* | 59.2 | 0.00 | 0.934 | 1.007 | 1.079 |
| *dlt* | 47.4 | 0.00 | 0.946 | 1.007 | 1.068 |
| *lu* | 90.9 | 0.00 | 0.849 | 1.007 | 1.165 |
| *dlu* | 52.3 | 0.00 | 0.914 | 0.986 | 1.058 |

use of the CD test and the estimated confidence bands of $\alpha$, i.e. the exponent of CSD (Pesaran, 2004; Bailey et al., 2016). The null hypothesis of the CD test states the variable is weakly cross-sectional dependent. This is an appealing feature since only strong cross-sectional dependence makes estimates inconsistent (Chudik et al., 2011). The exponent supplements the analysis as it shows the degree of cross-sectional dependence. The exponent $\alpha$ is a measure of the strength of the factors which can take any value in the range [0,1] (Bailey et al., 2016). Values of $\alpha < ½$ correspond to different degree of weak cross-sectional dependence, whilst values of $\alpha \geq ½$ point to strong sectional dependence. Pesaran (2015) shows the null of the CD test is a function of the degree to which *T* expands relative to *N*. In our setting, as in macro panels, both should diverge roughly at the same rate and the implicit null is $0 \leq \alpha < 1/4$.

Results of cross-sectional dependence analysis are presented in Table 2 which show the null hypothesis of weak cross-sectional dependence is always strongly rejected at the 1% level of significance.[3] Furthermore, the bias corrected estimates of the exponent are close to unity for all the variables in levels and first differences, while the 90% confidence bands are approximately $\pm 0.07$ and largely above 0.5.[4] Hence, this evidence points to the presence of strong cross-sectional dependence in all the variables under scrutiny.

To investigate the persistence properties of the data we follow Bai and Ng (2004), who deem $z_{i,t} = a_i^z + \psi_i f_t + v_{i,t}$, and assume for the unobservable $f_t$ and $v_{i,t}$ the following data generating processes: $(\mathbf{I} - L)f_t = \mathbf{C}(L)e_t$, $(1 - \rho_i)v_{i,t} = \mathbf{B}(L)v_{i,t}$ where $\mathbf{C}(L) = \sum_{j=0}^{\infty} C_j L_j$, $\mathbf{B}(L) = \sum_{j=0}^{\infty} B_j L_j$ are polinomials in the lag operator. Hence, we can test $\rho_i = 0$ without imposing stationarity of the common factors and vice-versa. To accomplish the first task, i.e. checking stationarity of the idiosyncratic components, we can pool individual ADF *t* statistics from defactored residuals. The PANIC approach proxies factors by principal components (Bai and Ng, 2010), whereas PANICCA relies on cross-section averages augmentation (Reese and Westerlund, 2016). However, PANIC requires to preliminary establish the number of common factors needed to represent the cross-sectional dependence in the data. These should not be very numerous in a macroeconomic setting (Stock and Watson, 2002, 2005) and a small number should be sufficient to explain most of the variation while usual selection methods, such as the Akaike or the Bayesian information criteria, are instead prone to overestimate the number (Westerlund and Urbain, 2015). Moreover, in the light of the small dimension of our sample, we assume there is explicitly at most one common factor. Unit-root tests of the idiosyncratic components can be the Pa and Pb statistics and the PMSB test statistic (Bai and Ng, 2010), whereas a simple ADF-type test for nonstationarity can be used for the single factor (Reese and Westerlund, 2016). The former three test statistics converge to a standard normal under the unit root null hypothesis ($\rho_i = 1$). Provided that the alternative is formulated as $|\rho_i| < 1$ for some *i*, all three statistics are left-tailed, and the appropriate 5% critical value is therefore given by 1.645.

Results of unit root analysis are presented in Table 3. When we add a trend, all the variables at their levels are non-stationary with dynamics driven by stochastic trends in both the single common factor and the idiosyncratic components. Nevertheless, electricity consumption, urbanization and trade openness are I(1). Without a trend, urbanization is once again I(1), whereas the evidence is mixed with respect to electricity consumption and trade openness. However, we have to consider that, in small samples, the Pa and Pb statistics tend to over-reject the null, whilst PMSB test tends to under-reject it (Reese and Westerlund, 2016). Hence, electricity per capita may be stationary while trade openness is I(1) at usual significance level. Finally, the behavior of GDP per capita is puzzling. When we only deem a constant, idiosyncratic components seem to be stationary, while the common factor displays a unit root. When we add a trend, this variable is not stationary whereas evidence is mixed with respect to its first difference.

To shed light on this issue we employ the PANICCA approach with additional variables since all the variables under investigation are decomposed in the same way and the data generating process can be easily expressed combining common factors and idiosyncratic components. Hence, GDP per capita analysis is extended to include the other variables available in our dataset. Results are displayed in Table 4 once more for one common factor. Without a trend, output may be stationary or at most I(1). With a trend, the empirical evidence is still mixed, but it supports the hypothesis GDP per capita is integrated of order one. The inclusion of these additional variables suggests the source of non-stationarity is due to both the common factor and the idiosyncratic components.[5] All in all, we can safely conclude all the variables under scrutiny are I(1).[6]

We can now apply the Granger causality test devised by Dumitrescu and Hurlin (2012) on the first differences of the series under examination, where the Bayesian information criterion selects the lag length and *p*-values are based on a bootstrap procedure. Table 5 shows that, in the short-run, a change in the GDP per capita, urbanization as well as in the trade openness ratio does not affect the change in electricity consumption in all the panels. On the contrary, electricity consumption does Granger cause output in accordance with the findings by Shiu and Lam (2004), Yuan et al. (2007), Yuan et al. (2008), and Chang (2010). Furthermore, changes in international trade may be due to variations in the urbanization rate and electricity per capita, but only at the 10% significance level. Finally, urbanization appears to be caused by trade and output. Notably, just urbanization and trade seem to display network effects in the short-run.

The long-run direction of causation is investigated with the two-step procedure suggested by Canning and Pedroni (2008). In the first step, we use FMOLS to estimate the cointegrating relationship and construct the disequilibrium term. We also check whether these terms are stationary according to Maddala-Wu or CIPS test when CSD is detected. Since the stationary requirement is fulfilled, we can estimate system (6)–(9) as suggested by Eberhardt and Teal (2013). Results are displayed in Table 6.

Group Mean *t*-tests reveal no causality link in all the specifications, whereas Lambda–Pearson statistics indicate causality is running from urbanization, trade, and GDP to electricity consumption as well as from urbanization, trade, and electricity consumption to output. These differences highlight the large heterogeneity in the error correction term coefficients across provinces, since the GM is a two-sided test

[3] From now all the variables are in logs.

[4] These are given by Eq. (13) of Bailey et al. (2016).

[5] We also performed CIPS tests with Wald test to select individual dynamics (Burdisso and Sangiácomo, 2016). They confirm electricity per capita, trade openness, and urbanization are I(1) at 1% significance level. When the Portmanteau test is embraced, only electricity per capita appears to be stationary at 5% significance level. Results are available in Table A.1 of the appendix.

[6] Finally, we also use Bai and Carrion-I-Silvestre (2009) panel unit root tests by allowing for three endogenous breaks. Results are shown in Table A.2 of the appendix and support the findings that we obtained from the PANIC and PANICCA analyses.

**Table 3**
PANIC and PANICCA unit root tests.

| | | Without trend | | | | With trend | | | |
|---|---|---|---|---|---|---|---|---|---|
| | | Idiosyncratic Components | | | | Idiosyncratic Components | | | |
| Var | CF | ADF | Pa | Pb | PMSB | ADF | Pa | Pb | PMSB |
| *le* | 0 | −2.47[b] | −2.59[a] | −1.58[c] | −1.20 | 2.78 | 0.46 | 0.50 | 0.53 |
| | 1 | −2.52[b] | −2.69[a] | −1.66[b] | −1.23 | 3.31 | 0.46 | 0.50 | 0.53 |
| *dle* | 0 | −3.77[a] | −22.92[a] | −7.51[a] | −2.28[b] | −3.96[a] | −13.09[a] | −7.07[a] | −2.58[a] |
| | 1 | −3.94[a] | −21.87[a] | −7.28[a] | −2.24[b] | −4.05[a] | −13.02[a] | −7.08[a] | −2.63[a] |
| *ly* | 0 | 4.69 | −5.54[a] | −3.90[a] | −1.78[b] | 4.69 | 1.79 | 2.27 | 2.78 |
| | 1 | 4.69 | −5.43[a] | −3.85[a] | −1.75[b] | 4.69 | 1.82 | 2.32 | 2.86 |
| *dly* | 0 | −0.67 | −10.08[a] | −5.22[a] | −2.61[a] | −0.62 | −2.76[a] | −2.05[b] | −1.26 |
| | 1 | −4.49[a] | −7.99[a] | −4.24[a] | −2.71[a] | −4.52[a] | −0.36 | −0.35 | 0.26 |
| *lt* | 0 | −1.52 | −1.84[b] | −1.56[c] | −0.89 | −1.58 | 0.79 | 0.87 | 0.95 |
| | 1 | −1.56 | −1.86[b] | −1.58[c] | −0.90 | −1.62[c] | 0.78 | 0.85 | 0.93 |
| *dlt* | 0 | −4.41[a] | −19.37[a] | −7.71[a] | −2.92[a] | −4.34[a] | −8.00[a] | −4.67[a] | −2.02[b] |
| | 1 | −4.40[a] | −19.31[a] | −7.68[a] | −2.90[a] | −4.33[a] | −8.11[a] | −4.72[a] | −2.03[b] |
| *lu* | 0 | 4.69 | 0.36 | 0.42 | 1.34 | 4.69 | −1.58[c] | −1.37[c] | −1.01 |
| | 1 | 2.41 | 1.01 | 1.25 | 2.22 | 4.69 | −1.25 | −1.12 | −0.85 |
| *dlu* | 0 | −4.58[a] | −32.19[a] | −9.89[a] | −2.54[a] | −4.44[a] | −17.74[a] | −9.16[a] | −2.71[a] |
| | 1 | −4.54[a] | −41.95[a] | −11.95[a] | −2.91[a] | −4.56[a] | −22.71[a] | −11.54[a] | −3.22[a] |

PANICCA when CF = 0, PANIC with one common factor when CF = 1.
[a] Significant at 1%.
[b] Significant at 5%.
[c] Significant at 10%.

**Table 4**
GDP per capita PANICCA unit root tests with additional variables.

| | Without trend | | | | With trend | | | |
|---|---|---|---|---|---|---|---|---|
| | Idiosyncratic Components | | | | Idiosyncratic Components | | | |
| Var | ADF | Pa | Pb | PMSB | ADF | Pa | Pb | PMSB |
| *le* | −3.89[a] | −3.43[a] | −2.40[a] | −1.93[b]C | 3.63 | 1.09 | 1.22 | 1.30 |
| *lt* | −1.65[c] | 5.20 | 18.39 | 58.87 | −1.61[c] | 1.28 | 1.46 | 1.57 |
| *lu* | 4.69 | −5.24[a] | −3.78[a] | −1.68[b] | 4.69 | 1.79 | 2.28 | 2.80 |
| *le&lt* | 4.69 | 6.20 | 14.86 | 25.72 | −1.54 | 1.26 | 1.44 | 1.54 |
| *le&lu* | −3.77[a] | −3.28[a] | −2.38[a] | −1.76[b] | 3.63 | 1.08 | 1.21 | 1.28 |
| *lt&lu* | −1.61[c] | 5.18 | 18.29 | 58.19 | −1.60 | 1.28 | 1.46 | 1.57 |
| *all* | 4.69 | 6.14 | 13.74 | 21.84 | −1.52 | 1.26 | 1.44 | 1.54 |
| *dle* | −3.96[a] | −9.75[a] | −4.95[a] | −2.84[a] | −4.07[a] | −1.90[b] | −1.57[b] | −1.11 |
| *dlt* | −4.39[a] | −11.16[a] | −5.26[a] | −2.97[a] | −4.32[a] | −2.20[b] | −1.80[b] | −1.26 |
| *dlu* | −4.29[a] | −9.04[a] | −4.61[a] | −2.82[a] | −4.02[a] | −1.26 | −1.12 | −0.85 |
| *dle&dlt* | −4.36[a] | −11.14[a] | −5.30[a] | −3.01[a] | −4.29[a] | −2.20[b] | −1.80[b] | −1.26 |
| *dle&dlu* | −3.96[a] | −9.68[a] | −4.91[a] | −2.83[a] | −4.05[a] | −1.84[b] | −1.53[c] | −1.09 |
| *dlt&dlu* | −4.40[a] | −11.15[a] | −5.26[a] | −2.97[a] | −4.33[a] | −2.20[b] | −1.78[b] | −1.26 |
| *all* | −4.36[a] | −11.10[a] | −5.29[a] | −3.01[a] | −4.29[a] | −2.20[b] | −1.80[b] | −1.26 |

[a] Significant at 1%.
[b] Significant at 5%.
[c] significant at 10%.

and can take positive or negative values under the alternative hypothesis, whilst the LP is a one-sided test that can only take positive values. Since all the two-tailed test fails to reject the null, while the Lamba-Pearson succeeds in rejecting it in two cases, we can conclude the adjustment coefficients are on average zero, but not pervasively zero when we deem electricity consumption and GDP per capita. In other words, a long-run causal effect is present, even if the effect is positive in some Chinese provinces whilst it is negative in others. Hence, per capita GDP, the urbanization rate, and the trade openness ratio Granger cause per capita electricity consumption over the long-run as well as migration from rural areas into cities, trade, and electricity are triggering output. Summing up these results we can conclude that trade and

**Table 5**
Short-run panel causality.

| test | Zbar | Prob | test | Zbar | Prob | test | Zbar | Prob | test | Zbar | Prob |
|---|---|---|---|---|---|---|---|---|---|---|---|
| *dly* → *dle* | 1.61 | 0.28 | *dle* → *dly* | 5.97 | 0.02 | *dle* → *dlt* | 2.60 | 0.09 | *dle* → *dlu* | −0.74 | 0.59 |
| *dlt* → *dle* | −0.97 | 0.48 | *dlt* → *dly* | 0.02 | 0.99 | *dly* → *dlt* | −0.99 | 0.53 | *dly* → *dlu* | 3.58 | 0.05 |
| *dlu* → *dle* | −0.76 | 0.57 | *dlu* → *dly* | 0.45 | 0.77 | *dlu* → *dlt* | 2.81 | 0.08 | *dlt* → *dlu* | 4.28 | 0.02 |

**Table 6**
Long run causality tests.

| | *GM* | *(p)* | *LP* | *(p)* | *Mean* $\tilde{\varphi}_i$ | *z-stats* |
|---|---|---|---|---|---|---|
| *le* | −1.597 | 0.11 | 170.5 | 0.00 | −0.631 | −8.21 |
| *lu* | 0.331 | 0.74 | 60.6 | 0.31 | 0.081 | 1.38 |
| *ly* | 0.371 | 0.71 | 85.6 | 0.01 | 0.201 | 1.04 |
| *lt* | 0.193 | 0.85 | 55.3 | 0.50 | 0.784 | 1.05 |

urbanization are Granger causing per capita GDP and electricity consumption, whereas the latter are likely to display feedback effects.

## 5. Conclusions and policy implications

The objective of this paper is to investigate whether urbanization, output, and trade openness are key factors to explain the rapid increase in per capita electricity consumption in a sample of 28 Chinese provinces during the period 1995–2016. The main contribution to the literature is the application of a sound econometric technique that takes into consideration the presence of strong cross-sectional dependence in a heterogeneous panel. The time series properties of the variables under scrutiny are as follows. The PANIC and PANICCA approaches show urbanization and trade openness are always driven by stochastic trends both in common factors and idiosyncratic components as well as electricity consumption when we add a trend. Since GDP dynamic is somehow puzzling, we also employ the PANICCA approach with additional variables. Empirical evidence is mixed, but it supports the hypothesis all the variables under investigation are I(1).

Granger causality tests on first differences show changes in electricity consumption Granger cause changes in output, as suggested by Yuan et al. (2008) among others, while urbanization appears to be caused by trade and output. Only urbanization and trade display feedback effects in the short-run. Long-run Granger causality is investigated with the method suggested by Canning and Pedroni (2008). Herrerias et al. (2013) analyze Chinese regions with the same procedure and find feedback effects between electricity consumption and per-capita GDP. However, our analysis includes also urbanization and trade and reveals a large heterogeneity in causal effects. We detect a unidirectional causal link running from urbanization, output, and trade to electricity consumption and another one from electricity, urbanization, and trade to output. Hence, over the long-run output and electricity consumption appear to move together pushed by urbanization and trade openness.

Previous studies on China focused primarily either on the energy-urbanization or on the growth-electricity nexus, whereas the novel aspect of this work is to merge them using a rich dataset at the regional level. Getting a better understanding of the electricity-output-urbanization-trade relationship is of utmost importance as China's energy saving at the provincial level has important policy implications for the mitigation of the greenhouse gas emissions in the world too. Menegaki and Tsani (2018) claim the energy-growth nexus studies are not particularly verbose to whether energy strategies and targets can be fulfilled nor make distant future projections and forecasting. Since a substantial reduction in the Chinese GDP growth rate is not foreseeable, environmental sustainability is in peril when the energy generation mix will continue to use mostly coal as nowadays. China should get over traditional dependence on fossil fuel by increasing electricity supply diversification with a preference for cleaner, renewable, and cost-effective energy such as hydropower, nuclear power, and solar or wind force. Actually, the national government has made the supply of clean energy a high priority increasing considerably the generation of renewable energy. However, according to Wang et al. (2016), the lack of coordination and consistency in regional clean energy development planning has produced a mismatch between supply and demand with an excess capacity is some areas. The national and local governments should also accelerate the construction of natural gas networks and power grids expanding its coverage in the Southwestern provinces, which are endowed by abundant hydropower and natural gas resources. Moreover, policy makers should also provide wind solar energy and solar water heat collectors to mutually complement power stations where agricultural and pastoral activities dominate or in remote Western provinces.

At the same time, China should adopt an alternative economic development model and develop new strategic industries to readjust its economic structure as advocated by Zhong et al. (2019) among others. Zheng and Walsh (2019) also suggest China should give up part of its heavy industry to expand knowledge intensive sectors. However, the share of heavy industry has accounted for about 70% of the total Chinese industrial value in the last decades and reconsidering its role is far from being easy in the short-run. However, to tackle the dilemma between electricity consumption and growth, the central and local governments should focus on electricity generation mix amelioration, production structure optimization with the premise of enhancing electricity security and keeping regular economic growth.

The findings presented in this paper show that electricity consumption might be reduced not only via strategic planning but with an increase in urban consumption too. This is a novel view, as some authors claim urbanization is the third most important factor in the rise of energy consumption and the second most important factor in the increase in CO2 emissions (Wang et al., 2016). On the contrary, we argue that urban consumers can reallocate their expenditures by purchasing less electricity-intensive goods and services. Actually, Zheng and Walsh (2019) claim there exists a "U-shaped" relationship between urbanization and economic growth and our results suggest this is the case in several Chinese provinces. If growth and electricity consumption are positively related whereas urbanization negatively affects electricity utilization, an increase in urbanization may be beneficial for a country's healthy development.[7] Such an increase must be targeted to specific provinces yet. This view is shared by Ahmad and Zhao (2018) who find high variation in the impact of urbanization on economic growth at province/city level, suggesting that policies at the province/city level will be more effective than at the aggregate level. Such an approach is confirmed by our findings and also by Wang et al. (2018), who state that some large megacities, such as Beijing and Shanghai, should continue innovation-driven development and engage in energy-saving and environmentally friendly development, whereas under-urbanized regions should reduce their dependence on energy-intensive industries, and update their manufacturing sectors.

[7] The "National New-type Urbanization Plan (2014–2020)" sets China will have 60% of its people living in cities by 2020. The current rate is about 54%.

## Appendix A

**Table A.1**
Panel Unit Root Test (CIPS) with individual lags.

| Variable | Without trend | | With trend | |
|---|---|---|---|---|
| | Wald | Portmanteau | Wald | Portmanteau |
| *le* | −2.04 | −2.17** | −2.17 | −2.29 |
| *dle* | −3.99*** | −4.01*** | −4.30*** | −4.28*** |
| *ly* | −2.03 | −1.93 | −2.57 | −2.31 |
| *dly* | −2.71*** | −2.74*** | −3.02*** | −2.93*** |
| *lt* | −1.65 | −1.60 | −2.14 | −2.01 |
| *dlt* | −3.75*** | −3.75*** | −3.85*** | −3.78*** |
| *lu* | −1.17 | −1.08 | −2.06 | −2.03 |
| *dlu* | −3.74*** | −3.74*** | −3.91*** | −3.91*** |

Critical values without trend at 10%, 5%, 1%: −2.07, −2.15, −2.30.
Critical values with trend at 10%, 5%, 1%: −2.58, −2.66, −2.81.

**Table A.2**
Panel unit root test results with breaks.

| Variables | Break in the mean | | | Break in the trend | | | | | |
|---|---|---|---|---|---|---|---|---|---|
| | Z | P | Pm | Z | P | Pm | Z* | P* | Pm* |
| *le* | 14.541 | −4.635 | 9.229 | −0.540 | 2.050 | 82.459 | 5.071 | 1.249 | 73.689 |
| *ly* | 10.335 | −4.273 | 13.188 | 8.282 | −0.971 | 49.365 | 22.674 | −1.857 | 39.652 |
| *lt* | 22.681 | −4.825 | 7.135 | 6.399 | −2.841 | 28.881 | 10.382 | −3.919 | 17.063 |
| *lu* | −0.317 | −1.643 | 41.997 | 1.317 | −1.071 | 48.265 | 1.984 | −1.896 | 39.233 |

Notes: (a) Z, P and Pm denote the test statistics developed by Bai and Carrion-i-Silvestre (2009), the 5% critical values of which are 1.645, −1.645 and 50.998, respectively. (b) Z*, P*, and Pn* refer to the corresponding statistics obtained using the p-values of the simplified MSB statistics. (c) The number of common factors is one. (d)The maximum number of breaks allowed is three. To determine the breaks, the procedure of Bai and Perron (1998) is used (for details see Bai and Carrion-i-Silvestre, 2009).

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