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Does corruption matter for the environment? Panel evidence from China

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Abstract
This article examines the income-energy-SO_2 emissions nexus by taking a corruption variable into account. To that end, the panel cointegration methods are applied to 29 Chinese provinces over 1999–2012. The authors´ empirical evidence shows that an increase in the number of anti-corruption cases tends to drive down SO_2 emissions in China. It is also found that income growth appears to have a beneficial effect on decreasing SO_2 emissions over the past two decades. Finally, energy consumption is found to increase SO_2 emissions.

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Keywords China; corruption; environment; EKC; panel; SO2

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1 Introduction

China has achieved rapid economic growth at an average rate of almost 10 percent annually over the past three decades. This economic success, however, comes at the cost of deterioration of the environment. One of the most severe environmental problems that China is currently facing is air pollution. For example, the State Environmental Protection Administration of China (SEPA) reports that about 70% of the 300 cities in China fail to meet the air quality standards set by the World Health Organization (WHO) and seven out of the ten most polluted cities in the world are located in China. The World Bank estimates that the direct cost of air pollution – such as acid-rain damage to crops, medical bills and job-loss from illness – ranges between 8 percent and 12 percent of China's GDP annually (World Bank, 2007). In addition, it is estimated that, because of heavy air pollution, more than three million people die prematurely each year and the average life expectancy is more than 5 years lower for residents in northern China than those living in the south (Wang, 2007; Pope III and Dockery, 2013).

The Chinese government has made substantial efforts to reduce air pollution by introducing various emission reduction measures such as environmental taxes/charges, pollution treatment programs and even closure of inefficient power/industrial facilities. Moreover, under the 12th Five-Year Plan (2011–2015), China has paid considerable attention to energy and climate change issues and has established a new set of targets and policies for the plan period. The main targets include a 16 percent reduction in energy intensity (energy consumption per unit of GDP), an increase in non-fossil energy up to 11.4 percent of total energy consumption and a 17 percent reduction in carbon intensity (carbon emissions per unit of GDP). In 2015, however, China still recorded the world's largest increment in energy consumption for the thirteenth consecutive year and became the world's largest emitter of both carbon dioxide (CO_2) and sulfur dioxide (SO_2) emissions. Therefore, a fundamental question would be certain to arise regarding China's environment: what are the main determinants affecting air pollution in China?

A number of studies have sought to isolate the independent effects of various factors on air pollution in China. Traditional specification of this subject includes a growth variable (i.e., income per capita) and investigates the *environmental Kuznets curve* (EKC) – an inverted U-shaped relationship between income per capita and certain types of pollutants (typically measured by CO_2 emissions). Then, as we glance through the literature more, we come across empirical studies that claim that energy consumption could be an important determinant of environmental outcomes and analyze the so-called income-energy-environment relationship. Examples include, but are not limited to, Song et al. (2008), Jalil and Mahmud (2009), Baek et al. (2009), Baek and Koo (2009), Jalil and Feridun (2011), Wang et al. (2011), Govindaraju and Tang (2013), Michieka (2014), Qu and Yan (2014), Yuan et al. (2015), Wang et al. (2016) and Li et al. (2016). The findings from these studies generally show that there is ambiguous evidence in favor of the EKC for China – e.g., Baek et al. (2009) and Baek and Koo (2009) for no evidence of the EKC, and Jalil and Mahmud (2009), Song et al. (2008) and Li et al. (2016) for evidence of the EKC – and strong evidence that China's growth in energy consumption indeed causes environmental degradation (e.g., Song et al., 2008; Govindaraju and Tang, 2013; Qu and Yan, 2014; Li et al, 2016).

Important but perhaps less widely recognized in the literature is the possibility that corruption could be an important factor of air pollution in China because China's deteriorating environment coincides with the soar of corruptibility. The fact is that domestic firms in China use a bribe to lobby officials to lower the environmental standards. Although China's government vows to fight corruption, it appears that the level of corruption has constantly increased over the past decade. For the years 2002 through 2009, for example, the people's procuratorates at all levels investigated more than 240,000 cases of embezzlement, bribery, dereliction of duty and infringements on rights. In 2009 alone, more than 3,000 people were punished for their criminal liability in offering bribes (Information Office of China's State Council, 2010). Thus, a corruption variable should be accounted for when estimating factors affecting China's deteriorating environment properly.

Up until now, many scholars have sought to address the impact of corruption on the environment. Examples include, but are not limited to, Lopez and Mitra (2000), Damania et al. (2003), Fredriksson et al. (2003), Fredriksson and Svensson (2003), Welsch (2004), He et al. (2007), Cole (2007), Woods (2008) and Leitao (2010). Damania et al. (2003), for example, examine the corruption-environment nexus in a panel data of developing and developed countries, and find that corruption indeed reduces environmental policy stringency. He et al. (2007) employ cross-country data and confirm the findings of Damania et al. (2003) in that a higher level of corruption always reduces the quality of environmental regulation. Woods (2008) reports that political corruption serves to systematically weaken state environmental programs in the United States. However, attention of most studies has been on cross-country data when investigating the corruption-environment nexus. Thus, the existing literature does not directly address the issue in China. This observation has motivated us to conduct this line of research.

The main objective of this paper is to take a measure of corruption into account in a model when examining the income-energy-environment relationship in China. Although China is currently the world's largest SO_2 emitting country along with CO_2 emissions, empirical studies have paid little attention to SO_2 emissions in their analyses.[1] Empirical focus is thus on assessing the effects of corruption, income and energy consumption on SO_2 emissions using panel data of 29 provinces in China from 1999 to 2012. To that end, the panel cointegration methods are utilized.

This paper is organized as follows: in Section II, we outline the empirical model to be estimated and the data used for the estimation. In Sections III and IV our empirical procedures and major findings are discussed, respectively. Finally, section V makes some concluding remarks.

[1] Baek et al. (2009) is perhaps the only study addressing the issue; they find that growth has a beneficial effect on reducing SO_2 emissions in China. However, they only examine the income-environment nexus.

2 Methodology

2.1 The model to be estimated

In examining factors affecting a country's environment, researchers generally rely on the so-called standard model of the income-energy-environment nexus (e.g., Iwata et al., 2010; Baek and Kim, 2013; Baek, 2015). For the empirical model adopted here, we extend the income-energy-environment nexus to include a measure of corruption. Letting *i* denote the cross-sectional unit (Chinese provinces in this paper) and *t* the time period, we can write a model in a log-linear form as:

$$\ln(so_2)_{it} = \alpha + \beta_1 \ln y_{it} + \beta_2 \ln y_{it}^2 + \beta_3 \ln ec_{it} + \beta_4 cor_{it} + \varepsilon_{it} \quad (1)$$

where $(so_2)_{it}$ is the sulfur dioxide (SO_2) emissions for province *i* in China; y_{it} is the 2000 real income for province *i*; y_{it}^2 is the square of the 2000 real income for province *i*; ec_{it} is the energy consumption for province *i*; cor_{it} is a measure of corruption for province *i* and is the number of anti-corruption cases; and ε_{it} is the error term.[2] The variables are measured on a per capita basis. We are particularly interested in the parameter β_4 – that is, the ceteris paribus effect of corruption on SO_2 emissions.

Given that numerous studies commonly show the crucial role of income plays in influencing environmental outcomes (e.g., Jalil and Mahmud, 2009; Iwata et al., 2010; Baek and Kim, 2013; Baek, 2015; Li et al., 2016), it would be proper to directly test the Environmental Kuznets Curve (EKC) hypothesis into our modeling. In Eq. (1), to the extent that $\beta_1>0$ and $\beta_2<0$, the EKC hypothesis is predicted to hold; that is, income has a diminishing effect on SO_2 emissions after the turning point (or maximum point of the income), achieving a parabolic shape. It is expected that $\beta_3>0$ due to the fact that an increase in energy consumption mainly driven by growth is likely to push SO_2 emissions up. Finally, it is expected that $\beta_4<0$ because the increasing number of anti-corruption cases is likely to result in improved environmental regulations, thereby reducing SO_2 emissions.

2.2 Data

SO_2 emissions are used as a proxy for a measure of air pollution. China is the world's largest coal consumer, accounting for about 50% of the world's total coal consumption. Coal combustion generates more than 90% of SO_2 emissions in China. As a result, China currently ranks the largest SO_2 emitter worldwide. To ensure comparability with income per capita in Eq. (1), the SO_2 emissions per capita for individual provinces (measured in 10 thousand metric tons) are calculated using their population size. The provincial gross domestic product per capita (measured in constant 2000 Chinese Yuan) is used as a proxy for real per capita income for each province. The energy consumption is measured in 10 thousand metric tons of coal equivalent

[2] It should be admitted that, since we use China's provincial data in estimating Eq. (1), it would be more desirable to control for many other provincial characteristics in the model. Unfortunately, however, the lack of data availability at the provincial level prevents us from considering more control variables in Eq. (1).

per capita. The number of anti-corrupt cases is used as a measure of the degree of corruptibility and is intended to allow for the likelihood that higher anticorruption efforts are likely to less environmental degradation.[3] The data on SO_2 emissions are collected from China Environmental Statistical Yearbooks. All the remaining variables are from China's Statistical Yearbooks.

Our (balanced) panel dataset contains the 29 Chinese provinces from 1999 to 2012 ($N*T$=406 observations, where N=29 provinces and T=14 years). This time period is chosen by availability of the data for all the variables. All variables are converted into natural logarithms. Table 1 summarizes our data.

Table 1 **–** Descriptive statistics

Variable	Obs.	Mean	Standard Deviation	Min	Max
so_2	406	0.015	0.011	0.002	0.061
y	406	18,000.95	13,390.93	2,537.02	68,296.06
ec	406	2.389	1.381	0.504	7.946
cor	406	1,271.88	788.83	110.00	3,954.00

Notes: so_2, y, ec and cor represent SO_2 emissions per capita for individual provinces (measured in 10 thousand metric tons), provincial gross domestic product per capita (measured in constant 2000 Chinese Yuan), energy consumption per capita for individual provinces (measured in 10 thousand metric tons of coal equivalent per capita), and the number of anti-corrupt cases, respectively.

3 Empirical results

The first requirement for estimating our model in Eq. (1) using the panel cointegration method is that the variables must be nonstationary $I(1)$ series. Accordingly, the panel cointegration modeling normally starts with testing whether a panel series follows a unit root. However, the possibility of cross-sectional dependence in panels is likely to invalidate the test statistics of conventional panel unit root tests such as the LLC (Levin et al., 2002) and IPS tests (Im et al., 2003). These tests commonly assume the cross-sectional independence in panels. Before applying a unit root test, therefore, we must test whether a panel series is cross-sectionally independent. A cross-sectional dependence (CD) test of Pesaran (2004) can be used to achieve

[3] The Chinese government has been making great efforts to combat corruption and to build a clean government using the following four measures over the past decades. The first measure is to select officials on the basis of democratic, open, competition, and preferred standards in order to prevent corrupted officials from being selected. The second measure is to establish a sound law system and regulations against corruption. The third measure is to put the power under the control by institutional innovation. The last is to build up a monitoring system including Chinese Communist Party inner-party supervision, supervision of the National People's Congress, democratic supervision of Chinese People's Political Consultative Conference, government supervision, judicial supervision, civil supervision and supervision by public opinion. Thus, anti-corruption cases should be relevant in using a proxy for corruption. Some scholars (e.g., Damania et al., 2003; Cole, 2007) use governmental honesty taken from the International Country Risk Guide (ICGR) as a proxy for corruption in their models. At the sub-national level, however, the data are not available.

this goal. The results show that the null hypothesis of no cross-sectional dependence can be strongly rejected (the *p*-values for all five variables are zero to two decimal places), providing compelling evidence of the cross-sectional dependence in the sample (Table 2).

Given that the panel series are found to be cross-sectionally dependent, it is no longer appropriate to use conventional panel unit root tests. Hence, we employ more powerful tests that allow for cross-sectional dependence such as the cross-sectionally Im-Pesaran-Shin (CIPS) and cross-sectionally augmented Dickey-Fuller (CADF) tests.[4] The CIPS test results show that we cannot (can) reject the unit root hypothesis for the levels (first differences) of all series, indicating that all five variables are $I(1)$ processes (Table 3). The CADF test largely confirms this finding, although we can reject a unit root in the level of cor_{it}; when cor_{it} is differenced, however, the null is strongly rejected and this leads us to believe that cor_{it} is $I(1)$ variable. The upshot of the unit root tests is that all five series in Eq. (1) appear to be nonstationary $I(1)$ processes.

When estimating a nonstationary panel model, there is serious concern about spurious regression. In one important case, a regression estimating nonstationary $I(1)$ series is not spurious, and that is when the series are cointegrated. Hence, the presence of cointegration relationship among the variables is tested using the various tests developed by Pedroni (1999 and 2004) and Kao (1999). The results show that the null hypothesis of no cointegration can be rejected even at the 1% level of significance for all five tests, evidence that SO_2 emissions and its determinants have a long-run relationship (Table 4). In other words, whenever deviations from the long-run equilibrium take place, they would be transient: there are economic forces that drive SO_2 and its main factors back to restore the long-run equilibrium relationship.

Having learned about a potential long-run relationship among the five series, we now apply the FMOLS and DOLS panel estimators of Mark and Sul (2003) and Kao and Chiang (2001) to Eq. (1) in order to estimate the long-run parameters. We also report the estimated effects of the fixed effects (FE) estimator here for comparison.[5]

Table 2 – Results of cross-sectional dependence (CD) test.

	$\ln(so)_2$	$\ln y$	$\ln y^2$	$\ln ec$	$\ln cor$
CD statistic	29.73**	74.27**	74.14**	72.79**	29.72**
p-value	0.00	0.00	0.00	0.00	0.00

Notes: ** denotes rejection of the null hypothesis at the 1% level.

[4] The resulting tests are known as the second generation tests for a panel unit root in order to distinguish them from the conventional tests or the first generation tests.

[5] It should be pointed out that the endogeneity of corruption could be a potential weakness of our work; our findings should thus be viewed with caution. To avoid this, what is needed is a good instrumental variable when estimating Eq. (1). But the existing literature on the topic does not offer a proper instrumental variable, which is exogenous yet highly correlated with corruption. Further, we realize that, even if the proposed instrument is available in the literature, our use of provincial-level panel data might have made it more difficult for the instrument to be collected in China. The relatively consistent findings based on the two dynamic panel estimators and traditional FE should somehow mitigate our concern with the endogeneity issue and strength the credibility of our findings.

Table 3 – Results of panel unit root tests.

Variable	CADF		CIPS	
	Level	First difference	Level	First difference
$\ln(so)_2$	–2.42	–2.96**	–2.89	–3.98**
$\ln y$	–1.70	–2.81**	–1.53	–3.38**
$\ln y^2$	–1.24	–2.80**	–1.21	–3.36**
$\ln ec$	–2.48	–3.01**	–2.63	–4.25**
$\ln cor$	–2.14	–3.17**	–3.11**	–3.84**

Notes: CADF and CIPS represent cross-sectionally augmented Dickey-Fuller and cross-sectionally Im-Pesaran-Shin tests, respectively. ** denotes rejection of null hypothesis at the 1% level.

Table 4 – Results of panel cointegration tests.

Test	Statistics
Panel PP statistic	–8.23**
Group PP statistic	–14.73**
Panel ADF statistic	–6.09**
Group ADF statistic	–5.77**
Kao test statistic	–2.29**

Notes: ** denotes rejection of the null hypothesis of no cointegration at the 1% level.

4 Discussion

Table 5 reports the long-run effects for all independent variables and for each of the three estimated models. The estimates generated by the models seem remarkably consistent. The signs of the coefficients are the same across models, and the same variables are generally statistically significant in each model.

The coefficients on the income and the quadratic term are similar across models. Because the coefficient on y_{it} is always positive and the coefficient on y_{it}^2 is always negative, this equation literally implies that, at low value of income, an additional rise in income tends to increase SO_2 emissions. At some point, the effect becomes negative, and the quadratic shape means that the elasticity of SO_2 emissions with respect to income is decreasing as income increases. In other words, the finding seems to be supportive of the EKC hypothesis for Chinese SO_2 emissions. It turns out, however, that all of the 29 provinces in the sample have more than the calculated turnaround values of income, and so the part of the curve to the left can be ignored. Thus, SO_2 emissions in fact have monotonically fallen with income growth in China over the past decade. From policy perspectives, this result can be interpreted that the Chinese government's policies targeted to reducing air pollution are likely to work effectively without costing economic growth.

The partial effect of energy consumption on SO_2 emissions is always positive, and the magnitudes of the coefficient estimates are very similar across all three models. In the FMOLS model, for example, a one percent increase in energy consumption is estimated to increase SO_2

emissions by about 0.80% in China. Given the fact that growth largely leads to an increase in energy use, this finding suggests that any favorable growth effect on air pollution could be offset by a detrimental energy consumption impact.

The key policy variable, cor_{it}, seems to have the desired effect. The estimated coefficient is negative for all three models. The statistical significance is high for the FMOLS and fixed effects, and lacking for the DOLS. For example, the FMOLS coefficient (–0.17) implies that, for other things being equal, China can reduce SO_2 emissions by about 0.17% as the number of anti-corruption cases increases by one percent. To our knowledge, this is a new finding that has not been documented yet in the empirical literature. As a policy matter, this suggests that effective anti-corruption measures would improve the environment through the enforcement of environmental regulations in China. From a methodological perspective, this finding explains why the complementary features of different modelling approaches would be desirable to draw more robust conclusion and thus better understand the corruption-environment nexus in China.

Finally, in addition to learning about the long-run relationship in Eq. (1), utilizing the notion of causality enriches our understanding of the variables by providing causal inference (i.e., direction of causality). For completeness, therefore, the bootstrap panel Granger causality test developed by Emirmahmutoglu and Kose (2011) is utilized. This method is most useful when dealing with cross-sectional dependence in panels as we identify in our model. The results show strong bidirectional causation for 4 cases and unidirectional causation for 5 cases (Table 6). For example, the relationships between SO_2 emissions and energy consumption, and SO_2 emissions and corruption are characterized by bidirectional causality. This means that SO_2 emissions are significantly affected by changes in energy consumption (corruption) and energy consumption (corruption) is also influenced by changes in SO_2 emissions. On the other hand, there is unidirectional causality running from income to SO_2 emissions. This suggests that SO_2 emissions are significantly affected by changes in income, while income is not affected by changes in SO_2 emissions. Together, these findings provide evidence that all independent variables can be used to forecast future SO_2 emissions and justify the use of our model in Eq. (1).

Table 5 – Results of long-run estimates

Variable	FMOLS		DOLS		Fixed effects	
	Coefficient	*p*-value	Coefficient	*p*-value	Coefficient	*p*-value
ln*y*	2.09**	0.00	2.15**	0.00	2.06**	0.00
lny^2	–0.28**	0.00	–0.28**	0.00	–0.26**	0.00
ln*ec*	0.80**	0.00	0.72**	0.00	0.80**	0.00
ln*cor*	–0.17**	0.00	–0.03	0.00	–0.13**	0.00
R^2	0.88		0.96		0.87	

Notes: ** denotes significance at the 1% level.

Table 6 – Results of bootstrap panel Granger causality test.

Hypothesis	Fisher statistics	*p*-value	Conclusion
ln$y$ → ln(so)$_2$	119.09**	0.00	One-way causality from ln$y$ to ln(so)$_2$
ln(so)$_2$ → ln$y$	42.44	0.93	
lny^2 → ln(so)$_2$	113.56**	0.00	One-way causality from lny^2 to ln(so)$_2$
ln(so)$_2$ → lny^2	43.75	0.89	
lnec → ln(so)$_2$	151.22***	0.00	Two-way causality between lnec and ln(so)$_2$
ln(so)$_2$ → lnec	129.77***	0.00	
lncor → ln(so)$_2$	158.06***	0.00	Two-way causality from lncor and ln(so)$_2$
ln(so)$_2$ → lncor	83.47**	0.02	
lnec → ln$y$	74.67*	0.07	Two-way causality between lnec and ln$y$
ln$y$ → lnec	150.30***	0.00	
lnec → lny^2	76.33**	0.05	Two-way causality between lnec and lny^2
lny^2 → lnec	132.19***	0.00	
lncor →ln$y$	66.83	0.19	One-way causality from ln$y$ to lncor
ln$y$ →lncor	84.72***	0.01	
lncor → lny^2	66.81	0.19	One-way causality from lny^2 to lncor
lny^2 → lncor	92.69***	0.00	
lnec → lncor	94.19***	0.00	One-way causality from lnec to lncor
lncor → lnec	66.27	0.21	

Notes: ***, ** and * denote significance at the 1%, 5% and 10% levels, respectively.

5 Concluding remarks

Although corruption can have an effect on the environment in China, no study has directly addressed this issue empirically. In this short article, therefore, we take a corruption variable into account in a dynamic panel model when estimating the income-energy-SO_2 emissions nexus. Our results show that anti-corruption cases seem to have a beneficial effect on reducing SO_2 emissions in China. Other findings show that income growth tends to lower SO_2 emissions, while energy consumption increases SO_2 emissions.

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