Reg Environ Change (2011) 11:883–891
DOI 10.1007/s10113-011-0227-8

ORIGINAL ARTICLE

On the Nexus of SO_2 and CO_2 emissions in China: the ancillary benefits of CO_2 emission reductions

Xinye Zheng · Li Zhang · Yihua Yu · Shi Lin

Received: 15 October 2010 / Accepted: 20 April 2011 / Published online: 6 May 2011

Abstract This paper examines the issue of ancillary benefits by linking sulfur dioxide (SO_2) emission to CO_2 emission using a panel of 29 Chinese provinces over the period 1995–2007. In the presence of non-stationarity and cointegrating properties of these two data series, this paper applies the panel cointegration techniques to examine both the long-run and short-run elasticities of SO_2 with respect to CO_2. The major findings are that: (1) there exhibits a stable long-run equilibrium relationship between the SO_2 and CO_2 emission with the long-run elasticity being 2.15; (2) there exists a short-run relationship between these two emissions with the short-run elasticity being 0.04. In addition, following an exogenous shock that causes a deviation from the long-run equilibrium, it would take approximately 15 years for SO_2 emission to revert toward the long-run equilibrium path with an average annual convergence rate of 6.5%; (3) the derived ancillary benefits that is generated from one metric ton of CO_2 emission reduction, are 11.77 Yuan (approximately US$ 1.7) in the short run and 196.16 Yuan (US$ 30) in the long run. These findings are not only crucial from the econometric modeling perspective, but also have important policy implications.

Keywords Emission reduction · Ancillary benefits · Sulfur (Carbon) dioxide · Panel cointegration

X. Zheng · Y. Yu (✉) · S. Lin
School of Economics, Renmin University of China, Beijing 100872, China
e-mail: yihua.yu@ruc.edu.cn

L. Zhang
China Academy of Public Finance and Public Policy, Central University of Finance and Economics, Beijing 100081, China

Introduction

The issue of energy consumption is becoming increasingly important in public policy debate in the era of climate change. The combustion of fossil fuels emits both CO_2 and SO_2. CO_2 especially is seen as the principal culprit of the greenhouse effect. To prevent excessive concentrations of greenhouse gases (GHGs), several international CO_2-abatement agreements, such as the Kyoto Protocol, the Bali Action Plan, the Berlin Declaration, the Copenhagen Accord, and the Ambo Declaration, have been initiated in recent years. One strategy to fulfill every country's goal under the Kyoto Protocol is to replace fossil fuels by renewable energy sources to satisfy the demand for energy, heat, and fuel (Lasch et al. 2010).

In comparison to CO_2, whose importance is global in scope, SO_2 is a major pollutant that has a relatively short lifetime in the atmosphere and has negative effects that are regional or local in nature. The negative effects of SO_2 on health have been well documented in toxicological, human clinical and epidemiological studies (Electric Power Research Institute EPRI 2009). Specifically, repeated high exposures of SO_2 cause problems of bronchoconstriction, airway inflammation, and hyper-responsiveness in test animals. From the human clinical studies, a large body of epidemiological studies finds the association of SO_2 concentration with daily mortality (including cardiovascular and pulmonary mortality (Hedley et al. 2002), respiratory symptoms and diseases (Health Effects Institute 2004), emergency-room visits, and hospital admissions for asthma or chronic obstructive pulmonary disease (Anderson et al. 1997; Sunyer et al. 1997). In brief, the general findings are that exposure to high levels of SO_2 can cause problems of breathing difficulty, respiratory illness, emphysema, asthma, acute broncho spasm, chronic bronchitis, and heart diseases.

The relationship between CO_2 and SO_2 emissions should be solely determined by the process of energy consumption if the current emission technology is given (or the energy mix is unchanging) and if there is no any policy intervention in place. For example, the amount of CO_2 or SO_2 that is emitted by burning a metric ton of coal can be exactly determined for any particular plant or region given its current state of technology. However, due to the detrimental impacts of SO_2 emissions on the health of local communities, governments tend to intervene to cut SO_2 emissions. Hence, the emission ratio of SO_2 to CO_2 is indeterminate with government interventions. On the other hand, governments do not have incentives to cut CO_2 emissions voluntarily due to the public good nature of CO_2 emissions. As a result, the relationships between CO_2 and SO_2 emissions for different countries during different time periods exhibit different characteristics. In general, three patterns exist for the trends of SO_2 and CO_2 emissions in recent years. First, some countries, for instance Australia, Greece, and China, experienced a rise of both SO_2 and CO_2 emissions during 1990–2007 (United Nations 2010). Second, some countries such as United Kingdom and Germany or EEA-32 countries in general experienced a decline of both emissions (European Environment Agency, EEA 2010). Third, some countries such as the United States, Italy, and Spain experienced a decoupling of SO_2 and CO_2 emissions, i.e., as CO_2 emissions increased, SO_2 emissions decreased (United Nations 2010). The decoupling effect has been explained by various factors such as the introduction of flue gas desulfurization, the changes in the primary energy mix (i.e., switching from high sulfur-containing fuels to low sulfur fuels such as natural gas), environmental regulations (e.g., Clean Air Interstate Rule issued by the United States Environmental Protection Agency in 2005), etc.

It is well known that the combustion of fossil fuels, which contain mainly coal and crude oil with heavy sulfur content, emits not only CO_2 but also a range of other damaging pollutants including SO_2. Considering the coexistence between SO_2 and CO_2, reducing CO_2 emissions will reduce SO_2 emissions correspondingly, and such reduction obviously holds benefits for local jurisdictions. Hence, SO_2 emissions reduction can be regarded as an ancillary benefit of CO_2 emissions reduction. In the literature, the ancillary benefits, arguably, are regarded as important incentives to form an international coalition to reduce CO_2 emission (Ekins 1996a, b; Wagner 2001; IPCC 2007; Pittel and Rübbelke 2008; McEvoy and Stranlund 2009; Nemet et al. 2010), though ancillary benefits may also reduce incentives of leaving the agreement (Finus and Rübbelke 2008).

Considering that the CO_2 abatement is costly and the direct benefits associated with the CO_2 abatement are not obvious, policy makers have turned their attention to the size of the ancillary benefits. This is understandable as there is a need to know the size of the ancillary benefits in order to determine the appropriate magnitude of CO_2 abatement in a given country. The existing literature shows large variations in the size of ancillary benefits (Pittel and Rübbelke 2008; Nemet et al. 2010). For instance, based on 24 studies of developed countries, 22 studies show that the monetary value of ancillary benefit originated from CO_2 abatement varies from \$2/tCO_2 to \$128/tCO_2, with an average of \$44/tCO_2 (Nemet et al. 2010). Most of these studies used as the estimation strategy the computable general equilibrium (CGE) model, which is criticized to be sensitive to factors such as the methodologies, the discount factor, the parameter of price elasticity, and others chosen by the authors (Bell et al. 2008). These factors contribute to the variations in the estimated ancillary benefits.

Not only affected by the economic valuation methods, the value of ancillary benefits is also affected by the differing damage costs across countries. Rabl et al. (1998) estimated a damage cost arising from SO_2 emissions of US \$13.4 thousand per metric ton. Netcen (2004) estimated the damage cost to be between 3,072 and 27,930 Euro (1997 prices) per metric ton of SO_2 emissions. Sáez and Linares (1999) presented a range of damage costs arising from SO_2 emissions for 15 OECD countries in the base year 1995. The average damage cost due to SO_2 emissions is about 5,736 Euro. Belgium ranked the first, with the damage cost of between 11,388–12,141 Euro, while Finland had the smallest damage cost of between 1,027–1,486 Euro. The damage costs for Germany and United Kingdom are between 1,800–13,688 Euro and 6,027–10,025 Euro, respectively. Given that empirical results are inconsistent and that estimating accurately the size of the ancillary benefits of CO_2 abatement has important implications for policy decision-making, more research is worthwhile. To avoid problems in the existing CGE-based literature, we examine the size of ancillary benefits by linking SO_2 emissions to CO_2 emissions in an empirical econometric model. In other words, we attempt to investigate quantitatively how SO_2 emissions will be affected if one metric ton of CO_2 emission is reduced. More importantly, as shown later, with the data on the damage cost arising from SO_2 emission in China available, we are actually able to obtain the monetary costs that can be saved from SO_2 emissions reduction, pursuant on CO_2 emission reduction. This study is especially important in the context of China for the following reasons: China became the biggest emitter of CO_2 and SO_2 due to its rapid development in the economy as well as heavy dependence on coal in the past decade. The government responds to these two types of pollutions differently. The SO_2-related pollution has had significant adverse impacts on human health, ecosystems, and cultural resources, and has caused direct economic loss in millions RMB Yuan each year (Zheng et al. 2010). Huge efforts have been made by the Chinese government to control SO_2 emission. For example, in the fourth plenary session of the tenth National People's Congress, the central government explicitly mandated a 10% reduction in national SO_2 emissions to be accomplished by 2010. Under the new policy implemented in China's Eleventh Five-Year Plan (2006–2010), more advanced but costly technologies are used, and desulfurization devices are required to be installed in most power plants, which had 52% of total coal consumption in China (National Bureau of Statistics 2009). This new policy has effectively reduced SO_2 emission in China (Cao et al. 2009). However, the policy had no impact on reducing CO_2 emissions.

As a response to the international pressures calling upon China to cut its CO_2 emission, the Chinese government promised a 40–45% reduction on CO_2 emissions per unit GDP from 2005 levels by year 2020. However, even if this policy goal will be achieved, CO_2 emission in China could still increase. As a matter of fact, the Chinese government is unwilling to take actions that will reduce total emissions. The inactive response to CO_2 emission can be partly explained by the free rider problem inherent to CO_2 emission reduction, since CO_2 emission reduction will benefit not only China but also the rest of the world. Such reduction costs, which are estimated to be very high in some studies (Zou 2010), would be completely borne by China.

The policy outcomes resulted from these different policy orientations toward CO_2 and SO_2 emission reductions can be significantly improved. For example, if carbon taxation or a cap-and-trade system is used to control CO_2 emissions, energy efficiency will increase, and new energy technology will become economically viable and widely employed. Consequently, CO_2 and SO_2 emissions can be reduced accordingly. Nonetheless, as we have argued earlier, policies such as carbon taxation may exert negative impacts on economic growth. In order to assess how much CO_2 emission should be reduced, policy makers must weigh the benefits (for example, the ancillary benefits of SO_2 emission reductions) against the costs (for example, economic slowdown) pursuant on CO_2 emission reductions.

To examine the quantitative relationship between SO_2 and CO_2 emissions, we use panel data of 29 Chinese provinces over the period 1995–2007. The estimation proceeds in four general steps. First, we conduct several panel unit root tests to ascertain whether either emission series in question are non-stationary (or having a unit root). This step is important in that if these two data series are tested to be non-stationary, which means that the statistical properties such as the mean and variance of these two data series change with time, the traditional OLS estimation using these two non-stationary data could cause a spurious regression problem and lead to misleading inferences. Second, where the two series are found to be non-stationary, we proceed to use panel cointegration tests (Pedroni 1999a, 2004) to determine whether the SO_2 and CO_2 variables are cointegrated of the same order. That is to say, we test whether there exists a linear combination of the two non-stationary emission data that yields a stationary series. If this is the case, the cointegrating relationship can be interpreted as a stable long-run relationship between the SO_2 and CO_2 emission variables.

Third, the long-run relationship can then be estimated by applying the fully modified OLS technique (FMOLS hereafter, Pedroni 2000). The FMOLS method has the advantage to correct for bias found in standard OLS due to the endogeneity and serial correlation of regressors (Kao and Chiang 2000). Last, the existing long-run relationship between SO_2 and CO_2 emissions paves the way to allow us to estimate their short-run relationship through transforming the variables into first differences in an error correction model framework. Moreover, once the short-run and long-run relationships (elasticities) are identified empirically, the ancillary benefits can be derived.

The rest of the paper is organized as follows. Section 2 presents some stylized facts on the SO_2 and CO_2 emissions in China. Section 3 describes the data source used in this study. Section 4 specifies the empirical econometric model and a series of diagnostic tests on the SO_2 and CO_2 emissions data. Section 5 reports the empirical results. The last section concludes.

Stylized facts

Facts of CO_2 emissions in China

Over the last three decades, China has witnessed explosive economic growth, which has exerted a huge influence on the environment with change of CO_2 emissions serving as a prime example. In 1980, China's total CO_2 emissions amounted to 1.46 billion tons, accounting for only 7.89% of the world total (Fig. 1). By 2004, China emissions reached 18.1% of the world total, ranking China the second largest emitter behind the United States (22.1%). In 2007, China surpassed the US as the largest emitter, reaching 6.28 billion metric tons (20.01% of world total). Figure 1 shows China's emissions during the period 1980–2007, exhibiting a rapid increase in recent years.

As former director of the State Administration of Work Safety Baoming Zhang put it, China's abundant coal reserves (94%) dwarf its reserves of oil (5.4%) and natural gas (0.6%). China is also the world's largest consumer of coal, which generates over 70% of its energy supply and

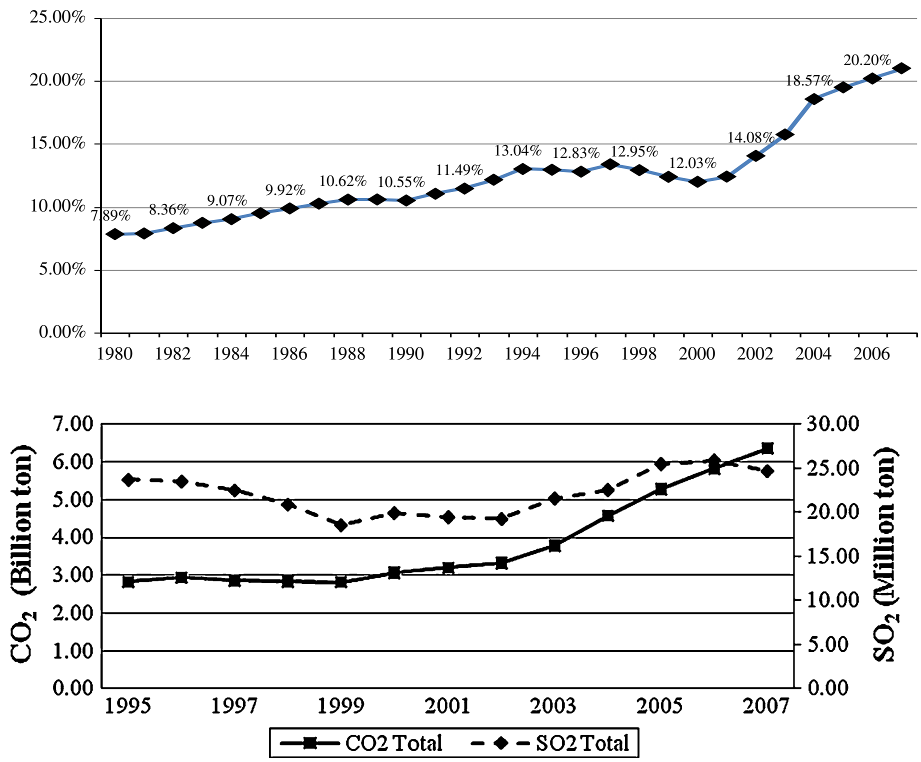

Fig. 1 China's share of the world's total CO_2 emission (1980–2007), and total CO_2 emission in China (1995–2007) China Statistical Yearbook 2008 (Top); Authors' calculation (Bottom, Tibet is excluded in calculation)

which is the most carbon-intensive fossil fuel. These facts help explain why CO_2 emission in China is higher than the rest of the world and also why China is facing severe environmental problems.

Facts of SO_2 emissions in China

SO_2 affects the environment in ways different from CO_2. First, the CO_2-induced pollution problem holds greater importance at a global level, while SO_2 has negative effects that are regional or local in nature (e.g., SO_2 induced acid rain). Second, SO_2 air pollution will cause respiratory and other health problems for human beings quickly, while CO_2 tends to have fewer noticeable short-term effects. SO_2 not only jeopardizes the well-being of mankind, but also erodes buildings, acidifies lakes and streams, damages forests and threatens the broader ecological balance.

China is among the countries that are heavily affected by SO_2 pollution. As deputy director of Ministry of Environmental Protection Lijun Zhang explained during the China Electricity Council's 2005 China Power Forum, "the current amount of SO_2 emissions in China has exceeded the capacity of the atmosphere by more than 80%." Acid rain is estimated to fall on over 30% of the total national land area (Stein 2008). Figure 1 illustrates that the total SO_2 emission appears to exhibit a fluctuating but increasing trend between 1995 and 2007. The total amount of SO_2 emissions increased from 23.7 in 1995 to 24.7 million metric tons by 2007, with an annual growth rate of 0.35%.

High levels of SO_2 emissions can be attributed to various reasons. One reason lies in that China uses the coal as the major energy source where both carbon content and sulfur content are comparably high. The other reasons include the relatively low energy efficiency of China's industries and households in coal usage, the costly desulfurization equipment and relatively slow speed in popularizing such equipment, and the failure to scrub waste gases before releasing them into the atmosphere. With heavy reliance on coal and its poor energy efficiency, the continuing growth of China's SO_2 emission is inevitable.

To sum up, China is the biggest emitter of CO_2 and SO_2 emissions as a result of China's industrial development process and coal-dominated energy infrastructure. As SO_2 and CO_2 in China are produced mostly from combustion coal, a relevant question is whether there exists a quantitative relationship between SO_2 and CO_2 emissions, and if the answer is yes, how strongly are they interrelated in the short and long term? What implications can be drawn from their relationship? These questions will be examined in this study.

Data source

Annual data on SO_2 emissions for each province are directly retrieved from the *China Statistical Yearbook*. Yet, there are no officially published data for CO_2 emissions. To obtain data for CO_2 emissions, we recognize that there are energy usage data for different sources of energy (coal, gas, oil, etc.) that are published by the *China Statistical Yearbook*. The published energy usage data enable us to use an internationally recognized method, i.e., the Intergovernmental Panel Climate Change (IPCC) method, to estimate the amount of CO_2 emitted by various types of fuel. The formula applied in estimating CO_2 emissions is prescribed by the *IPCC Guidelines for National Greenhouse Gas Inventories* (2006) and is defined as,

$$E_i = (FC_i)(CAL_i)(CC_i)(CO_i)(44/22) \tag{1}$$

where i denotes for various fuel types, FC denotes total amount of fuel consumption, CAL is the calorific value of the fuel (say, measured in joules), CC is the carbon content for each fuel, CO is the oxidization rate measuring how much of the carbon as a percentage in each fuel is oxidized into CO_2 during combustion. Hence, an oxidization rate of unity implies complete combustion. The last term is the ratio of the molecular weight of CO_2 (m. w. 44) to the molecular weight of carbon (m. w. 12). Therefore, complete combustion of 1 metric ton of the coal with a carbon content of 78% and a heating value of 14.76 million joules produces approximately 2.86 metric tons of CO_2.

Data for total consumption of each fuel type (coal, crude oil, and natural gas) and their corresponding calorific value can be obtained from the *China Statistical Yearbook*, and the carbon content and carbon oxidization rate of each fuel can be borrowed from the *IPCC Guidelines for National Greenhouse Gas Inventories*. As the combustion technology and operating conditions vary significantly from country to country, in the context of China, the oxidization rates for coal, crude oil, and natural gas are set to be 0.90, 0.98, and 0.98, respectively (Zhu et al. 2006). Recognizing that the oxidization rate of coal usage can vary from sector to sector in each province, For example, the oxidization rate can be different in the power sector from that in the manufacturing sector. Consequently, the oxidization rate used in this study could induce possible measure error on calculating provincial per capita CO_2 emissions (the dependent variable).

The data set for this study consists of a panel of 29 Chinese provinces (including three centrally administered municipalities—Beijing, Tianjin, and Shanghai, and four autonomous regions—Guangxi, Inner Mongolia, Ningxia, and Xinjiang) covering the period 1995–2007. The autonomous region, Tibet, is omitted from the sample due to data unavailability. Though considered as a centrally administered municipality since 1997, Chongqing is merged into Sichuan province in view of historical administrative relationship. Hence, eventually we have a total of 377 observations for this study.

Econometric model and diagnostic tests

The panel model is specified as,

$$SO_{2it} = \beta_0 + \beta_1 CO_{2it} + \varepsilon_{it} \tag{2}$$

where $i = 1, 2, \ldots, 29$ identifies each province in the panel; $t = 1, 2, \ldots 13$ represents the time period. SO_{2it} and CO_{2it} are per capita SO_2 and CO_2 emissions for province i at period t, respectively. There are several options that could be applied to estimate Eq. 2, namely, the classical regression model with a single constant term or the pooled ordinary least squares (OLS) model, the province-specific and/or time-specific models (or the fixed effect model), or the random province and/or time effect model (or the random effect model). The latter two models differ in the assumption that whether unobservable heterogeneity is correlated with exogenous variables in the regression equation. However, based on Fig. 1 which plots the annual SO_2 and CO_2 emission data from 1995 to 2007, we can see that both emission variables appear to be trending upwards over time. This implies that these two variables (or per capita) may be non-stationary, which means these two variables whose means and variance may change over time by definition. If that is the case, using classical estimation methods such as the OLS, to estimate relationships with unit root variables could cause spurious regression problem and hence give rise to misleading inferences, as we regress a non-stationary SO_2 series on another non-stationary CO_2 series.

Therefore, we will examine the role of CO_2 emission on SO_2 using recently developed panel cointegration methods. This method proceeds with four steps in general. First, we conduct several panel unit root tests to determine whether these two emission series in question are non-stationary. If the two series are found to be non-stationary, we proceed to the next step to implement several panel cointegration tests developed by Pedroni (1999a, 2004) to test whether the SO_2 and CO_2 emission variables in question are cointegrated of the same order. i.e., we test whether there exists a linear combination of the two non-stationary emission data that yields a stationary series. If this is the case, we can claim a stable long-run relationship exists between SO_2 and CO_2 emissions. In the third step, the long-run relationship (or long-run elasticity) can be estimated by applying the FMOLS technique to Eq. 2. Finally, the existence of the long-run relationship allows us to estimate the short-run relationship through transforming the variables into first differences in an error correction model framework.

Panel unit root tests

To avoid spurious regression and bias of the OLS estimation, and considering the panel data's nature of parameter heterogeneity, we perform five panel unit root tests to identify the stationarity properties of the SO_2 and CO_2 variables in question. These tests are Levin et al. (2002) test, henceforth LLC test Breitung (2000), test Im et al.

(2003) test, henceforth IPS test, Fisher-type augmented Dickey–Fuller (ADF, Dickey and Fuller 1981) test, henceforth ADF–Fisher test, Philips Perron test (henceforth PP test). As the traditional ADF unit root test has been criticized for having a problem of low power in rejecting the null hypothesis of stationarity of the time series (e.g., Hall 1994; Ng and Perron 2001), the literature suggests that the five panel unit root tests applied in this paper have higher power than the unit root test based on individual time series such as the ADF test.

However, all these panel unit root tests we used in this study are still based on the ADF process as follows,

$$\Delta y_{it} = \rho_{it} y_{i,t-1} + \delta_i + \theta_t + \eta_i t + \sum_{j=1}^{pi} r_{ij} \Delta y_{i,t-j} + \varepsilon_{it} \tag{3}$$

where $i = 1, 2, \ldots N$ identifies each province in the panel; $t = 1, 2, \ldots T$ represents the time period; δ_i is the individual constant; $\eta_i t$ is the individual time trend; θ_t is the common time effect; p_i is the number of lags in the ADF process. The null hypothesis of stationarity is that ρ is identical across provinces and equal to zero, while the alternative hypothesis of unit root is that $\rho < 0$ for at least one province.

All tests rely on the assumption that the error term ε_{it} is a stationary process (i.e., the expectation function $E[\varepsilon_{it}\varepsilon_{it}] = 0$ for all t, s and $t \neq s$), which is the assumption required for calculating common time-effects (Banerjee et al. 2005). The inclusion of the term $\sum_{j=1}^{pi} r_{ij}\Delta y_{i,t-j}$ accounts for serial correlation (possibly different across provinces) in the ADF regression errors. In addition, the tests have different assumptions about the degree of heterogeneity of ρ. Also, the inclusion of individual constants and time trends is optional. However, the Breitung's (2000) test requires that individual trends are included. In addition, LLC and Breitung tests assume a common unit root process, in other words, these two tests assumes that ρ to be identical for all provinces, while IPS, ADF–Fisher, and PP tests presented by Maddala and Wu (1999) assume an individual unit root process, that's, ρ is assumed to be different for each province.

Table 1 reports the results of panel unit root tests for the level model and the first-differenced model, which are performed using Eviews 6.0. The models are specified with fixed effects and provincial individual time trend given the obvious upward pattern of CO_2 (SO_2) emissions. The optimal number of lag length is determined by Schwartz Information Criterion (SIC). On the whole, these five panel unit root tests yield a consistent result that the level series of CO_2 and SO_2 are non-stationary, and their first differences appear to be stationary. This result indicates that these two emission data are non-stationary and are integrated of order one, denoted I (1).

Panel cointegration tests

Given that both variables contain a panel unit root, we proceed to examine whether there is a long-run relationship between CO_2 and SO_2 emissions using the Pedroni (1999a, 2004) panel cointegration test, which allows for cross-sectional interdependency of different individual effects. Pedroni's test is based on the following specification,

$$y_{it} = \alpha_i + \delta_t + \beta X_{it} + \varepsilon_{it} \tag{4}$$

where $i = 1, 2, \ldots N$ stands for each province in the panel; $t = 1, 2, \ldots T$ is the time period; α_i and δ_t allow for province-specific fixed effects and deterministic trends, X_{it} is a vector of explanatory variables. The disturbances are assumed to be independent and identically distributed. The variables y_{it} and X_{it} are assumed to be integrated of order one. Thus, under the null hypothesis of no cointegration

Table 1 Results of panel unit root tests

	LLC	Breitung	IPS	ADF–Fisher	PP
Levels					
SO_2	−3.957**	3.843	1.485	47.567	57.812
	(0.0000)	(0.9999)	(0.9313)	(0.8342)	(0.4823)
CO_2	−6.307**	0.237	−1.469*	75.597*	74.916*
	(0.0000)	(0.5936)	(0.0709)	(0.0602)	(0.0668)
First differences					
SO_2	−16.218***	−4.707***	−7.995***	154.305***	200.025***
	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)
CO_2	−16.590***	−4.621***	−8.071***	145.713***	176.660***
	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)

The tests are Levin et al. (2002), Breitung (2000), Im et al. (2003), ADF–Fisher χ^2 (ADF), PP χ^2 (PP) which is due to Maddala and Wu (1999). The statistics are asymptotically distributed as standard normal with a left-hand side rejection area. p values are reported in the parentheses. The optimal number of lag length included in the unit root test is determined by Schwartz Information Criterion (SIC). A ***, **, and * indicate respectively the rejection of the null hypothesis of non-stationarity at 1, 5, and 10% levels. Estimations are conducted by Eviews 6.0

(i.e., $\rho_i = 1$), the following unit root test is conducted on the residuals $\varepsilon_{it} = \rho_i \varepsilon_{it} + \mu_{it}$. Pedroni's panel cointegration test consists of seven different test statistics. They are panel v-statistic, panel ρ-statistic, panel PP-statistic, panel ADF-statistic, group ρ-statistic, group PP-statistic, and group ADF-statistic. The first four test statistics are based on the within-dimension approach that assumes common autoregressive coefficients across all cross sections (ρ), while the latter three test statistics are based on the between-dimension approach that assumes autoregressive coefficients to vary across the cross sections (ρ_i). All seven test statistics follow an asymptotically standard normal distribution and their detailed derivations can be found in Lee (2005) and Narayan and Narayan (2008).

For comparison purposes, we also use the Kao (1999) test to identify a common cointegrating relationship for all provinces that allows for fixed effects but no country-specific trends or common time-effects. Table 2 presents the results of Pedroni's (1999a, 2004) and Kao's (1999) panel cointegration tests. Most of the tests reject the null hypothesis of no cointegration at least on the 1% level of significance, indicating a long-run cointegrating relationship exists between the CO_2 and SO_2 emission data.

Empirical results of the long-run and short-run elasticities

Once the series are found to be panel non-stationary and cointegrated of order one, we can claim that a long-run relationship exists between CO_2 and SO_2 emissions. This forms the basis to conduct the short-run estimation through transforming the variables into first differences, and an adjustment mechanism can be included in the difference model to estimate the speed of adjustment of the SO_2 emissions to its equilibrium following a shock to the system.

As far as the coefficient estimation is concerned, the regressors in Eq. 2 are likely to be endogenous, which leads to biased estimators if the OLS method is applied (Kao and Chiang 2000). As a result, the emission equation can be estimated using the bias-correction methods, the FMOLS approach developed by Pedroni (1999b) or the dynamic OLS (DOLS) approach developed by Stock and Watson (1993). Yet, as it is shown that the estimates from either or DOLS are asymptotically equivalent (Banerjee 1999; Dreger and Reimers 2005), we report only the FMOLS results for heterogeneous cointegrated panels. We also present the panel OLS estimators for comparison purposes. The results for the long-run estimates are given in panel A of Table 3, while the results for the short-run estimates are given in panel B.

With respect to the long-run estimation results using the FMOLS estimation technique, the coefficient on CO_2 is 0.01 and statistically significant at 1% level of significance. This result suggests that as CO_2 emissions decline by 1%, SO_2 emissions are reduced by 0.01%. Calculating at the mean level of each variable, the elasticity of SO_2 with respect to CO_2 is 2.15, revealing that SO_2 emissions are very elastic. Furthermore, the coefficient estimate (0.01) is found to be larger than that using the panel OLS approach (0.006), implying that using the panel OLS approach instead of the efficient FMOLS approach fails to get a strong impact of CO_2 emission reduction.

Turning to the results of the short-run estimation, consistent with the long-run equilibrating result, the regressor CO_2 remains positive and statistically significant in the short run. However, the magnitude is found to be much smaller in the short run (about 0.0006). To be more

Table 2 Results of panel cointegration tests

Pedroni test	
Within dimension test statistics	
Panel v-statistic	2.721 (0.0033)***
Panel ρ-statistic	0.790 (0.7853)
Panel PP-statistic	−10.860 (0.0000)***
Panel ADF-statistic	−7.620 (0.0000)***
Between-dimension test statistics	
Group ρ-statistic	3.067 (0.999)
Group PP-statistic	−9.972 (0.0000)***
Group ADF-statistic	−8.182 (0.0000)***
Kao residual test	
Kao–ADF	−3.162 (0.0008)***
Residual variance	1809.918
HAC variance	2674.478

All reported values are asymptotically distributed as standard normal. The corresponding p value is reported in parentheses. ***, ** and * indicate respectively the rejection of the null hypothesis of no cointegration at 1, 5, and 10% levels. Estimations are performed by Eviews 6

Table 3 Long-run and short-run estimates

Panel A: long-run estimates	FMOLS	Panel OLS
CO_2	0.010 (22.41)***	0.006 (42.21)***
Constant	6.620 (3.18)***	–55.730 (4.83)***
Panel B: short-run estimates		
ΔCO_2	0.0006 (2.60)***	
ECT (error correction term)	−0.0652 (3.77)***	
Constant	0.0539 (3.28)***	

t values are reported in parentheses and critical values in brackets. ***, ** and * indicate significance at 1, 5 and 10% level of significance, respectively. Estimations are performed by RATS specific, whereas the elasticity of SO_2 in the long run is 2.15, the calculated elasticity is only 0.04 in the short run. This result suggests that the effect of CO_2 emissions on SO_2 emissions is smaller in the short run, but with time it tends to affect more on SO_2 emissions. One intuitive explanation would be that the firm managers or policy makers respond slowly to the decline of CO_2 emissions in the short run. Their slow response could indicate that some tools, such as the desulfurization facilities, are performed less efficient in the short run than in the long run, causing SO_2 emission to fall slowly in the short run.

In addition, based on the ECM-based causality test, we find evidence of unidirectional causality running from CO_2 emissions to SO_2 emissions in the long run. This result implies that policies aimed at reducing CO_2 emission would significantly reduce SO_2-induced damage costs, generating the so-called ancillary benefits. The result also implies that policies targeted at reducing SO_2 emission would not affect CO_2 emission abatement.

Last, we recognize that the single-period lagged ECT is statistically significant at 1% level with a magnitude of −0.065, implying that, after an exogenous shock to the system, the SO_2 emission will eventually converge to its long-run equilibrium path. However, the speed of adjustment is extremely slow. Specifically, it would take approximately 15 years (1/0.065) for SO_2 emission to revert toward the long-run relationship following an exogenous shock that causes a deviation from this long-run relationship. The extremely slow speed of adjustment toward the long-run relationship is reflected by the finding that the short run elasticity (0.04) of SO_2 emissions with respect to CO_2 emissions is small compared to the long-run elasticity (2.15).

Concluding remarks

This main purpose of this paper is to examine the issue of ancillary benefits by linking SO_2 emission to CO_2 emission in an empirical econometric model for a panel of 29 Chinese provinces during the 1995–2007 period. This paper applies the panel unit root and cointegration techniques to examine both the long-run and short-run elasticities of SO_2 emissions with respect to CO_2 emissions. We found that in the short run, reduction of one metric ton of CO_2 is accompanied by 0.0006 metric tons of SO_2 emission reduction, while in the long run the number of SO_2 emission reduction is 0.01. These results imply that the short-run and long-run elasticities of SO_2 emissions with respect to CO_2 emissions, evaluated at the mean, are 0.04 and 2.15, respectively.

According to a document released by the Ministry of Environmental Protection of China, SO_2 emissions amounted to 25.49 million metric tons in 2005, which resulted in a total economic cost of 500 billion (Heng 2006). This report indicates that the damage cost arising from SO_2 emission is about 19,615 Yuan per metric ton. Linking this number to the short-run and long-run estimates as mentioned above, it can be seen that the monetary costs that can be saved, or the ancillary benefits that are generated from one metric ton of CO_2 emission reduction, are 11.77 Yuan in the short run and 196.16 Yuan in the long run.

It is noteworthy, though, that the ancillary benefits vary across regions. For instance, the SO_2 emission permit price in December 2009 was 6,315 Yuan for Shuozhou city in Shanxi province (Shanxi Report 2010), 6,150 Yuan for Chongqing city in December, 2009 (Chongqing Evening News 2009), 4,200 Yuan for Guangdong province and 5,100 Yuan for Shaanxi province in September, 2009 (CNWEST News 2010), 4,200 Yuan for Northwestern regions in June 2010 (Xi'an Evening News 2010). Thus, the costs that can be saved from purchasing SO_2 emission permits generated from one metric ton of CO_2 emission reduction, are calculated to be, respectively, 3.79, 3.69, 2.52, 3.06, 2.52 Yuan in the short run, and 63.15, 61.50, 42.00, 51.00, 42.00 Yuan in the long run.

In the context of China, the findings in this study show that there are huge co-benefits for CO_2 emission reduction. In other words, CO_2 abatement policies aimed to prevent global warming can not only mitigate emission abatement pressures of other countries, but also generate environmental ancillary benefits for China's own sake such as reductions in SO_2 emissions. The results of the quantitative analysis of the relationship between SO_2 and CO_2 emissions from this study undoubtedly shed light on how to make better CO_2-abatement related decisions for China's policy makers.

Acknowledgments This study was supported by the Fundamental Research Funds for the Central Universities, and the Research Funds of Renmin University of China (projects 10XNK059). We are grateful to two anonymous referees and the editor for their valuable comments. The usual disclaimer applies.

References

Anderson HR, Spix C, Medina S, Schouten JP, Castellsague J, Rossi G, Zmirou D, Touloumi G, Wojtyniak B, Ponka A, Bacharova L, Schwartz J, Katsouyanni K (1997) Air pollution and daily admissions for chronic obstructive pulmonary disease in 6 European cities: results from the APHEA project. Eur Respir J 10:1064–1071

Banerjee A (1999) Panel data unit roots tests and cointegration: an overview. Oxf Bull Econ Stat 61:607–629

Banerjee A, Marcellino M, Osbat C (2005) Testing for PPP: should we use panel methods? Empir Econ 30:77–91

Bell ML, Davis DL, Cifuentes LA, Krupnick AJ, Morgenstern RD, Thurston GD (2008) Ancillary human health benefits of improved air quality resulting from climate change mitigation. Environ Health 7:41

Breitung J (2000) The local power of some unit root tests for panel data. In: Baltagi BH (ed) Advances in econometrics, vol. 15. Nonstationary panels, panel cointegration, and dynamic panels. JAY Press, Amsterdam, pp 161–178
Cao J, Garbaccio R, Ho MS (2009) China's 11th five-year-plan and the environment: reducing SO_2 emissions. Rev Environ Econ Policy 3:231–250
Chongqing Evening News (2009) Chongqing engaged the first time in emission permits trading: one ton of SO2 emission costs 6150 Yuan. http://news.sohu.com/20091226/n269219410.shtml. Viewed 20 May 2010
CNWEST News (2010) Shaanxi province auctioned emission permits again: 20.61 million Yuan for 4251 tons of SO2. http://news.cnwest.com/content/2010-09/28/content_3544957.htm. Viewed 20 May 2010
Dickey DA, Fuller W (1981) Distribution of the estimators for autoregressive time series with a unit root. Econometrica 49:1057–1072
Dreger C, Reimers H (2005) Health care expenditures in OECD countries: a panel unit root and cointegration analysis. Int J Appl Econom Quant Stud 2:5–20
Ekins P (1996a) How large a carbon tax is justified by the secondary benefits of CO2 abatement? Resour Energy Econ 18:161–187
Ekins P (1996b) The secondary benefits of CO2 abatement: how much emission reduction do they justify? Ecol Econ 16:13–24
Electric Power Research Institute (2009) Health effects of sulfur dioxide. Electric Power Research Institute, Inc. http://mydocs.epri.com/docs/public/000000000001018635.pdf
European Environment Agency (EEA) (2010) The European environment-state and outlook 2010. European Environment Agency, Denmark
Finus M, Rübbelke DTG (2008) Coalition formation and the ancillary benefits of climate policy. Stirling economics discussion papers 2008–2013. Department of Economics, University of Stirling
Hall A (1994) Testing for a unit root in time series with pretest data-based model selection. J Bus Econ Stat 12:461–470
Health Effects Institute (2004) Health effects of outdoor air pollution in developing countries of Asia: a literature review. Health Effects Institute, Boston
Hedley AJ, Wong CM, Thach TQ, Ma S, Lam TH, Anderson HR (2002) Cardiorespiratory and all-cause mortality after restrictions on sulfur content of fuel in Hong Kong: an intervention study. Lancet 360:1646–1652
Heng Y (2006) China's sulfur dioxide emissions caused annual economic loss of 500 Billion Yuan. Elite reference (in Chinese). http://news.sina.com.cn/c/2006-08-14/164710721789.shtml. Viewed 20 May 2008
Im KS, Pesaran MH, Shin Y (2003) Testing for unit roots in heterogeneous panels. J Econom 115:53–74
IPCC, 2007: Climate Change 2007: Mitigation of Climate Change, B. Metz, O. Davidson, P. Bosch, R. Dave and L. Meyer, Eds., Cambridge University Press, Cambridge
Kao C (1999) Spurious regression and residual-based tests for cointegration in panel data. J Econom 90:1–44
Kao C, Chiang M (2000) On the estimation and inference of a cointegrated regression in panel data. Adv Econom 15:179–222
Lasch P, Kollas C, Rock J, Suckow F (2010) Potentials and impacts of short-rotation coppice plantation with aspen in Eastern Germany under conditions of climate change. Reg Environ Change 10:83–94
Lee CC (2005) Energy consumption and GDP in developing countries: a cointegrated panel analysis. Energy Econ 27:415–427
Levin A, Lin CF, Chu CSJ (2002) Unit roots tests in panel data: asymptotic and finite sample properties. J Econom 108:1–24
Maddala GS, Wu S (1999) A comparative study of unit root tests with panel data and a new simple test. Oxf Bull Econ Stat 61:631–652
McEvoy DM, Stranlund JK (2009) Self-enforcing international environmental agreements with costly monitoring for compliance. Environ Resour Econ 42:491–508
Narayan PK, Narayan S (2008) Does environmental quality influence health expenditures? Empirical evidence from a panel of selected OECD countries. Ecol Econ 65:367–374
National Bureau of Statistics (2009) *China statistical yearbook*. China Stat Press, Beijing
Nemet GF, Holloway T, Meier P (2010) Implications of incorporating air-quality co-benefits into climate change policymaking. Environ Res Lett 5:014007
Netcen (2004) Benefits table database: estimates of the marginal external costs of air pollution in Europe. BeTa version E1.02a. Created for European Commission DG Environment by Netcen
Ng S, Perron P (2001) Lag length selection and the construction of unit root tests with good size and power. Econometrica 69:1519–1554
Pedroni P (1999a) Critical values for cointegration tests in heterogeneous panels with multiple regressors. Oxf Bull Econ Stat 61:653–670
Pedroni P (1999b) Fully modified OLS for heterogeneous cointegrated panels. Indiana University working paper, pp 1–40
Pedroni P (2000) Fully modified OLS for heterogeneous cointegrated panels. Adv Econom 15:93–130
Pedroni P (2004) Panel cointegration: asymptotic and finite sample properties of pooled time series tests with an application to the PPP hypothesis. Econom Theor 20:597–625
Pittel K, Rübbelke DTG (2008) Climate policy and ancillary benefits: a survey and integration into the modeling of international negotiations on climate change. Ecol Econ 68:210–220
Rabl A, Spadar JV, McGavran PD (1998) Health risks of air pollution from incinerators: a perspective. Waste Manage Res 16:365–388
Sáez RM, Linares P (1999) The national implementation in the EU of the framework. CIEMAT, Madrid
Shanxi Report (2010) Survey on 90 million Yuan trading in SO2 emissions in Shuozhou. http://www.21cbh.com/HTML/2010-5-10/xNMDAwMDE3NjMxNA.html. Viewed 20 May 2010
Stein M (2008) When technology fails: a manual for self-reliance, sustainability, and surviving the long emergency, 2nd edn. Chelsea Green Publishing, Vermont
Stock J, Watson M (1993) A simple estimator of cointegrating vectors in higher order integrated systems. Econometrica 6:783–820
Sunyer J, Spix C, Quenel P, Ponce-de-Leon A, Ponka A, Barumandzadeh T, Touloumi G, Bacharova L, Wojtyniak B, Vonk J, Bisanti L, Schwartz J, Katsouyanni K (1997) Urban air pollution and emergency admissions for asthma in four European cities: the APHEA Project. Thorax 52:760–765
United Nation (2010) The millennium development goals report 2010. United Nations, New York
Wagner UJ (2001) The design of stable international environmental agreements: economic theory and political economy. J Econ Surv 15(3):377–411
Xi'an Evening News (2010) Northwestern regions engaged the first time in emission permits trading: one ton of SO2 emission costs 4200Yuan. http://news.xinhuanet.com/politics/2010-06/06/c_12186879.htm. Viewed 20 May 2010
Zheng X, Yu Y, Zhang L, Zhang Y (2010) Do pollution drive up public health expenditure? A panel unit root and cointegration analysis. RUC Center of Energy Economics, working paper no. 2010003. School of economics, Renmin University of China. http://www.hanqing.ruc.edu.cn/admin/uploadfile/201005/20100520103320946.pdf
Zhu C, Ma Z, Wang C, Liu Z (2006) Analysis of variations of energy consumption related CO2 emissions in China. Ecol Environ 15(4):1029–1034 (in Chinese)
Zou J (2010) Scenario analysis on China's low-carbon development path and technology choice. RUC School of environment working paper, Renmin University of China