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Abstract

This paper examines the relationship between the logarithms of CO₂ emissions and real GDP in China by applying fractional integration and cointegration methods. The univariate results indicate that the two series are highly persistent, their orders of integration being around 2, whilst the cointegration tests (using both standard and fractional techniques) imply that there exists a long-run equilibrium relationship between the two variables in first differences, i.e. their growth rates are linked together in the long run. This suggests the need for environmental policies aimed at reducing emissions during periods of economic growth.

JEL-Codes: C220, C320, Q560.

Keywords: CO₂ emissions, GDP, China, persistence, fractional integration, fractional cointegration.

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1. Introduction

China has been experiencing very rapid economic growth in recent decades, during which it has relied mainly on fossil fuels (coal and oil) that generate greenhouse CO₂ emissions. The relationship between economic growth and the environment is often understood in terms of the so-called environmental Kuznets curve (EKC) with its typical U-shape (as in the case of the original curve describing the relationship between income inequality and GDP per capita – see Kuznets, 1955). The basic idea is that emissions in countries in the early stages of development are low and therefore environmental quality indicators are good; the subsequent industrialisation process initially damages the environment, but as income per capita increases environmental legislation is introduced to reduce emissions and pollution (see Grossman and Krueger, 1991, and Shafik and Bandyopadhyay, 1992, among others).

A recent study by Riti et al. (2017) analyses the relationship between emissions, economic growth and energy consumption in China using a variety of econometric techniques such as ARDL, FMOLS and DOLS that produce similar results providing support for an EKC but finding a different turning point compared to other studies.

The present paper revisits these issues by applying fractional integration and cointegration methods that are more general than those based on the classical I(0) v. I(1) dichotomy found in most studies, and provide more accurate estimates of persistence. Having examined first the properties of the individual series, it then tests for the existence of a long-run equilibrium relationship linking them in the context of a (fractional) cointegration framework.

The layout of the paper is as follows. Section 2 reviews the relevant literature. Section 3 outlines the empirical framework. Section 4 describes the data and discusses the empirical findings, while Section 5 offers some concluding remarks.

2. Literature Review

The relationship between economic growth and CO₂ emissions has been analysed in numerous papers using a variety of approaches. Several of them have investigated the existence of an EKC (Grossman and Krueger, 1991, 1995; Shafik and Bandyopadhyay, 1992; Shafik, 1994; Dinda and Coondoo, 2006; Heil and Selden, 2001; Coondoo and Dinda, 2002; Friedl and Getzner, 2003; Barassi and Spagnolo, 2012; etc.), examining different factors of environmental degradation, specifically emissions of sulfur dioxide (SO₂) (Day and Grafton, 2003 and Llorca and Meunié, 2009), nitrous oxide (N₂O) (Cho et al., 2014), and methane (CH₄) (Cho et al., 2014), Total Suspended Particles (TSP) (Day and Grafton, 2003), and water wastes (Haisheng et al., 2005). CO₂ emissions are among the most widely used indicators of environmental degradation (Apergis and Payne, 2009; Lean and Smyth, 2010; Du et al., 2012; Shahbaz et al., 2013b; and Tiwari et al., 2013). Their relationship with the use of energy and the type of energy source, manufacturing and urbanization rate is investigated in many studies (Iwata et al., 2010; Al-Mulali et al., 2012; Martinez-Zarzoso and Maruotti, 2011; Zhu et al., 2012; Zhang and Lin, 2012; Li and Wei, 2015; Zhang et al., 2017; Sinha and Bathacharya, 2017; etc.).

Some studies use different indicators and examine the relationships between financial development and environmental pollution (Jalil and Feridun, 2011; Shahbaz et al., 2013a), whilst others investigate the linkages between economic growth, CO₂ emissions and energy consumption (Saboori and Sulaiman, 2013a,b; Chandran and Tang, 2013), and analyse different countries and regions, specifically America (Day and Grafton, 2003; Plassmann and Khanna, 2006; Apergis and Payne, 2009; Zilio and Recalde, 2011; and HamitHaggar, 2012), Europe and Central Asia (Ang, 2009; Atici, 2008; Acaravci and Ozturk, 2010; Pao and Tsai, 2011; Shahbaz et al., 2013c; and Ozturk and Acaravci, 2013), the Middle East and North Africa (Fodha and Zaghoud,

2010; Ozcan, 2013 ; Farhani et al., 2014; and Shahbaz et al., 2014), South Asia (Nasir and Rehman, 2011; Ahmed and Long, 2012; Shahbaz and Lean, 2012; and Tiwari et al., 2013), Sub-Saharan Africa (Orubu and Omotor, 2011 and Osabuohien et al., 2014), or for the MINT (Lin and Nelson, 2018). A few papers test for the existence of a long-run EKC (Markandya, Golub and Pedroso-Galinato 2006; Lindmark 2002; etc.).

Other studies have incorporated into the original EKC indicators such as commercial opening, direct foreign investment, environmental regulation, research and development intensity, and energy efficiency (Antweiler et al., 1998; Cole, 2004; Pao and Tsai, 2011; Jalil and Feridun, 2011 Sheng and Lü, 2012; Ren et al., 2014), and provided mixed results. Some report a positive relationship between CO₂ emissions and economic growth (Nasir and Rehman 2011; Ozturk and Acaravci, 2013; Bakhsh and al., 2017; Ahmad and Du, 2017; Wang et al. 2016, Ma et al. 2016, Saidi and Mbarek, 2016, and Chaabouni et al., 2016); these results appear to be sensitive to the production structure (Rafindadi, 2016; Apergis and Ozturk, 2015) or the degree of openness of the economy (Magazzino, 2015; Asumadu-Sarkodie and Owusu, 2016; Díaz and Cancelo, 2009). Other studies find instead a negative relationship between CO₂ emissions and economic growth (Roca et al. 2001; Azomahou et al., 2006; Ajmi and al., 2015; Salahuddin et al., 2016; Lin et al. 2017; Dogan and Aslan, 2017; Baek and Pride, 2014). There is also some evidence that this relationship is at first positive, and then becomes negative (Shahbaz et al., 2014; Shahbaz et al., 2016; Riti, Song, Shu & Kamah, 2017).

The results are also inconclusive in the specific case of China. Some papers find evidence supporting the EKC hypothesis (Roumasset, Burnett and Wang, 2008; Song, Zheng and Tong, 2008; He, 2009; Govindaraju and Tang, 2013; Haisheng et al., 2005; Jalil and Mahmud, 2009; and Jalil and Feridun, 2011), whilst others do not (Du et al., 2012; Wang et al., 2011; Llorca and Meunié, 2009; Govindaraju and Tang, 2013;

Onafowora and Owoye, 2014; etc.).¹ Gonvinsraju and Tang (2013) find a long-term relationship between CO₂ emissions, economic growth and coal consumption in China; their Granger causality tests provide strong evidence of unidirectional causality running from economic growth to CO₂ emissions.

A second strand of the literature performs decomposition analysis of the factors that affect the environment (Peters et al., 2007; Liu and Jiang, 2011; Li and Wei, 2015; Zhang et al., 2015; Xie and al. 2018) using methods such as Laspeyres and Divisia indices (Sun, 1998; Ang and Zhang, 1999). In the case of China, some studies find that the decreases in CO₂ emissions are due to increases in energy intensity (Wang et al., 2005; and Liu et al., 2007), and others that the main factors driving CO₂ emissions are economies of scale, population and / or energy structure (Xu et al., 2014; Xie et al., 2018). Ang et al. (1998) use the logarithmic Divide Index (LMDI) decomposition method to analyse the relationship between energy consumption and carbon emissions in China's industrial sector during the years 1985-1990, and conclude that the relationship between carbon emissions and energy consumption in this sector is positive, whilst the relationship between energy intensity and carbon emissions is negative. Jalil and Mahmud (2009) report that the main factors influencing carbon emissions in China in the long run are income and energy consumption. Zhang et al. (2009) find that economic activity increases carbon emissions and energy intensity decreases them.

Using structure decomposition (SDA) Peters et al. (2007) conclude that in China the growth in CO₂ emissions from infrastructure construction, household consumption in cities, the urbanization process and lifestyle has been greater than the savings from efficiency improvements. On the other hand, Li and Wei (2015) find that the impact of

¹Al-Mulali et al. (2012) provide a useful summary table with the factors of environmental degradation and the variables used in each study (Table 1, pp. 124-126).

the industrial structure on CO₂ emissions is gradually changing from positive to negative and that the main driver of the reduction of CO₂ emissions in China is carbon intensity. Zhang et al. (2015) carry out SDA analysis of the factors that influence China's pollutants and conclude that increasing efficiency and intensity of emissions are important factors in reducing industrial pollution.

A third line of research investigates the causal relationship between energy consumption and economic growth (see the seminal study by Kraft and Kraft, 1978), finding long-run unidirectional causality running from GDP to energy consumption in the case of the US over the period 1947-1974. Several studies have been carried out providing mixed results (Akarca, 1979; Akarca and Long, 1980; Yu and Jin, 1992; Shiu and Lam, 2004; Glasure and Lee, 1998); this is due to differences in the methodology and the data used, as well as country-specific characteristics (Chen et al., 2007). Some papers report evidence supporting the hypothesis of neutrality, according to which there is no causality between energy consumption and GDP; others find unidirectional causality from economic growth to energy consumption or in the opposite direction, or bidirectional causality (see Ozturk, 2010, for a comprehensive survey of this literature).

Other studies focus on gross fixed capital formation, labour force and carbon dioxide to investigate causality between energy consumption and economic growth: Ghali and El-Sakka (2004), Huang et al. (2008), and Apergis and Payne (2009). Chontanawat et al. (2008) use data for 30 OECD countries and 78 non-OECD countries over the period between 1971 and 2000, and conclude that bi-directional causality between aggregate energy consumption and GDP is more frequent in the OECD developed countries than in the non-OECD developing countries. Lin and Nelson (2018) focus on the MINTs and find a two-way causal relationship between economic

growth and energy consumption, as well as between the former and FDI flows, and a unidirectional causal relationship between FDI and energy consumption.

In the case of China, evidence of causality between electricity consumption and economic growth is provided by Huang (1993), Shiu and Lam (2004), Yuan et al. (2007, 2008), and Chen et al.(2007). Regarding the direction of causality, Auffhammer and Carson (2008) suggest that China's planned trajectory of CO₂ emissions increased dramatically between 2002 and 2007 and reject the specification of the static environmental Kuznets curve, while Ang (2009) tries to find the determinants of CO₂ emissions in China using aggregate data for more than half a century, and finds that the intensity of research, technology transfer and economic absorption capacity to assimilate foreign technology are negatively related to carbon emissions in China, and that greater use of energy, growth of revenue and commercial openness tends to increase CO₂ emissions. Finally, Chang (2010) argues that energy consumption and CO₂ emissions increased as a result of targeting higher economic growth in China, and Nguyen et al. (2017) conclude that investment and trade openness play an important role in the relationship between carbon emissions, energy consumption and growth in China.

3. Fractional Integration and Cointegration

The order of integration of a time series is the differencing parameter required to make it stationary I(0). Specifically, a covariance stationary process $\{x_t, t = 0, \pm 1, \dots\}$ is said to be I(0) if the infinite sum of all its autocovariances, defined as $\gamma_u = E[(x_t - Ex_t)(x_{t+u} - Ex_{t+u})]$ is finite, i.e.,

$$\sum_{u=-\infty}^{\infty} |\gamma_u| < \infty. \quad (1)$$

Then, a process is said to be integrated of order d or I(d) if it can be represented as:

$$(1 - L)^d x_t = u_t, \quad t = 0, \pm 1, \dots, \quad (2)$$

where L is the lag operator ($L^k x_t = x_{t-k}$), and u_t is $I(0)$. One can use a Binomial expansion in equation (2) such that then, if d is fractional, x_t can be expressed as

$$x_t = d x_{t-1} - \frac{d(d-1)}{2} x_{t-2} + \frac{d(d-1)(d-2)}{6} x_{t-3} - \dots + u_t.$$

In other words, x_t is a function of all its past history, and the higher its value is, the higher is the level of dependence between observations distant in time. Thus, the parameter d measures the degree of persistence of the series. A very interesting case occurs when d belongs to the interval $[0.5, 1)$, which implies non-stationary but mean-reverting behaviour, with shocks having transitory but long-lived effects.

To estimate d for the individual series we use the Whittle function in the frequency domain (Dahlhaus, 1989) following the procedure described in Robinson (1994) (see also Gil-Alana and Robinson, 1997). The bivariate analysis is based on the concept of fractional cointegration, and uses the two-step approach of Engle and Granger (1987) extended to the fractional case as in Cheung and Lai (1993) and Gil-Alana (2003) as well as the Fractional CVAR (FCVAR) model proposed by Johansen and Nielsen (2010, 2012).

4. Data and Empirical Results

We use quarterly data on real GDP and CO₂ emissions in China, from 1978 to 2015, obtained from the Eikon database, which merges data from different sources into a single platform. Both variables appear to be highly trended (see Figure 1).

[Insert Figure 1 about here]

As a first step we examine the orders of integration of the two individual series, i.e. of the logs of CO₂ emissions and real GDP respectively. For this purpose, we consider the following model:

$$y_t = \alpha + \beta t + x_t, \quad (1 - B)^d x_t = u_t, \quad t = 1, 2, \dots, \quad (3)$$

and test the null hypothesis:

$$H_o: d = d_o, \quad (4)$$

in (3) for d_o -values of -2, -1.99, ..., -0.01, 0, 0.01, ..., 1.99 and 2 under two alternative assumptions for the I(0) error term u_t , namely that it follows a white noise and a weakly autocorrelated process as in the exponential spectral model of Bloomfield (1973) respectively. The latter fits extremely well in the framework suggested by Robinson (1994) and it is stationary for all values unlike the AR case (see, e.g. Gil-Alana, 2004). As for the deterministic terms, we consider the three cases of i) no terms, ii) a constant, and iii) a constant and a linear time trend, and choose the specification with statistically significant coefficients. The results are displayed in Table 1.

[Insert Table 1 about here]

The two individual series appear to be highly persistent. In the case of white noise residuals the estimated values of d are 1.87 and 1.92 respectively for the log CO₂ and log GDP series, and a significant positive time trend is found in the latter case. When allowing for autocorrelation, the estimated values are 1.91 and 1.82, and the null hypothesis of I(2) behaviour cannot be rejected since the 95% confidence intervals include the value of 2 for both series.

[Insert Table 2 about here]

Table 2 displays the estimates of d using the “local” Whittle semi-parametric approach of Robinson (1995).² When using this method, the estimates must be in the range $(-0.5, 0.5)$, and therefore we carry out the analysis using the second differences. The null of $I(0)$ behaviour cannot be rejected in any case regardless of the bandwidth parameter. Thus, both the parametric and semi-parametric results indicate that the two series are non-stationary with orders of integration around 2.

[Insert Table 3 about here]

Next we examine the possibility of fractional cointegration by using in the first instance the method suggested by Gil-Alana (2003), which is an extension of the Engle and Granger (1987) approach to the fractional case. Thus, in the first step, we test for the order of integration of the two variables (in first differences). Since the previous results imply that they are $I(1)$, in the second step, we regress one variable against the other and test whether the residuals are integrated of order $d - b$, these two parameters corresponding to the orders of integration of the two variables of interest. We display in Table 3 the results for the two cases of uncorrelated and autocorrelated errors for three different estimation approaches for the coefficients in the regression model:

i) OLS in the time domain, i.e.,

$$\frac{\sum_{t=1}^T (1-L) \log CO2_t (1-L) \log GDP_t}{\sum_{t=1}^T ((1-L) \log GDP_t)^2}; \quad (5)$$

ii) OLS in the frequency domain, i.e.,

² It is semi-parametric in the sense that no specific model is assumed for the $I(0)$ error term. This method (Robinson, 1995) was later extended and improved by Phillips and Shimotsu (2005) and Abadir et al. (2007) among others, but the latter approaches require other user-chosen parameters in addition to the bandwidth and the results are very sensitive to those.

$$\frac{\sum_{j=0}^{T-1} I_{(1-L)\log GDP (1-L)\log CO2}(\lambda_j)}{\sum_{j=0}^{T-1} I_{(1-L)\log GDP}(\lambda_j)}, \quad (6)$$

where $\lambda_j = 2\pi j/T$, $j = 1, \dots, T$ are the Fourier frequencies, and where for arbitrary sequences, w_t and v_t , we define the cross periodogram and periodogram respectively as

$$I_{wv}(\lambda) = \omega_w(\lambda) \omega_v(-\lambda)^T; \quad \text{and} \quad I_w(\lambda) = I_{ww}(\lambda),$$

with $\omega(\lambda)$ being the discrete Fourier transform of w_t : $\omega_w(\lambda) = \frac{1}{\sqrt{2\pi T}} \sum_{t=1}^T w_t e^{i\lambda t}$;

iii) finally, we also employ a narrow band least squares (NBLS) estimator, which is related to the band estimator proposed by Hannan (1963), and which is given by

$$\frac{\sum_{j=0}^m s_j \operatorname{Re} \left[I_{(1-L)\log GDP (1-L)\log CO2}(\lambda_j) \right]}{\sum_{j=0}^m s_j I_{(1-L)\log GDP}(\lambda_j)}, \quad (7)$$

where $1 \leq m \leq T/2$, $s_j = 1$ for $j = 0, T/2$ and $s_j = 2$ otherwise, the motivation for this third approach being that, since cointegration is a long-run phenomenon, when estimating the slope coefficient in the regression model one might be concentrating only on the low frequencies, which are those corresponding to the long-run, hence neglecting information about the high frequencies, which might distort the estimation results (see Gil-Alana and Hualde, 2009).

In the first of these cases, the estimated values of d are 0.83 and 0.87 respectively, and while the $I(1)$ hypothesis cannot be rejected with autocorrelated errors, it is rejected in favour of $I(d, d < 1)$ with white noise residuals, i.e. in the latter case we find mean reversion and fractional cointegration, though with a very slow rate of adjustment; however, when using the frequency domain least squares estimator, the values are much smaller, providing evidence of fractional cointegration in the two cases

of uncorrelated and autocorrelated errors; finally, when using the NBLs estimator in (7) the estimates are very sensitive to the choice of the bandwidth parameter and with $m = (T)^{0.5}$ the null of standard cointegration, i.e. $d = b = 1$, cannot be rejected. Thus, the results seem to be very sensitive to the estimation method used for the cointegrating regression and the bandwidth parameter.

Given the lack of robustness of the above results, we also apply the FCVAR method of Johansen and Nielsen (2010, 2012), first under the assumption that $d = b$, these two parameters being the order of (fractional) integration of the individual series, which implies that the cointegrating errors will be $I(d - b) = I(0)$. Their estimated order of integration is 1.024, which supports the hypothesis of classical cointegration, with the individual series being $I(1)$ and the cointegrating errors $I(0)$. Further, the null $d = b$ cannot be rejected by means of a LR test, which again implies standard cointegration. This finding is in contrast to the previous test results, which implied that standard cointegration should be rejected in favour of fractional cointegration (i.e., $d - b > 0$), and suggests that the earlier tests might be biased in favour of higher degrees of integration because of the method used for the estimation of the coefficients in the cointegrating regression (see Gil-Alana, 2003; Gil-Alana and Hualde, 2009). Classical cointegration between the two series, is also supported by the tests of Johansen (1988, 1996) and Johansen and Juselius (1990) (these test results are not reported for reasons of space). Therefore there is evidence of a stable long-run equilibrium relationship between the growth rates of CO₂ emissions and real GDP in China.

5. Conclusions

This paper has analysed the relationship between the logarithms of CO₂ emissions and real GDP in China by applying fractional integration and cointegration methods. The

univariate results indicate that the two series are non-stationary and highly persistent, their orders of integration being around 2, whilst the cointegration tests (using both standard and fractional techniques) imply that there exists a long-run equilibrium relationship between the two variables in first differences, i.e. their growth rates are linked together in the long run. This suggests the need for environmental policies aimed at reducing emissions during periods of economic growth. Future work will examine in greater depth the functional form of this relationship in order to gain a better understanding of the issues of interest.

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Figure 1: Time series plots

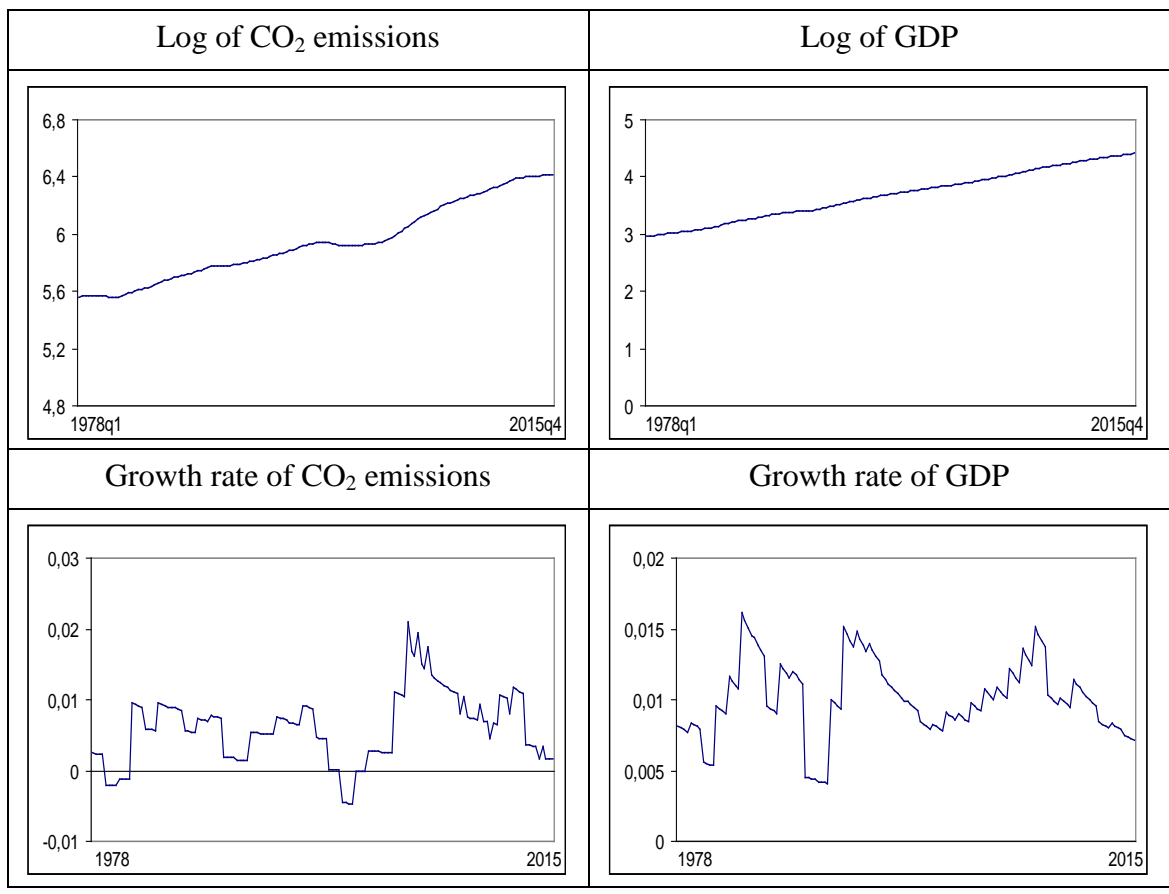


Table 1: Estimates of d on the individual series. Parametric approach

Series	d (95% band)	Intercept	A time trend
i) No autocorrelation			
Log CO ₂	1.87 (1.77, 2.01)	12.806 (3344.17)	-----
Log GDP	1.92 (1.79, 2.06)	6.797 (2226.49)	0.018 (4.64)
ii) With autocorrelation			
Log CO ₂	1.91 (1.69, 2.21)	12.803 (2519.22)	-----
Log GDP	1.82 (1.59, 2.14)	6.791 (2144.61)	0.019 (4.75)

Table 2: Estimates of d on the individual series. Semi-parametric approach

Bandwidth	$(1 - L)^2 \text{Log CO}_2$	$(1 - L)^2 \text{Log GDP}$	I(0) interval
10	-0.223*	-0.282	+/- 0.260
11	-0.123*	-0.180*	+/- 0.247
$12 = (T)^{0.5}$	-0.186*	-0.099*	+/- 0.237
13	-0.100*	-0.047*	+/- 0.228
14	-0.097*	0.025*	+/- 0.219
15	-0.034*	0.104*	+/- 0.212

*: Evidence of I(0) behaviour in the second differences.

Table 3: Cointegrating regression using first differenced data

	Constant	Slope	d - b (White noise)	d - b (Autocorrelation)
LSTD	-0.0101 (-3.29)	1.0200 (7.97)	0.83 (0.73, 0.96)	0.87 (0.62, 1.20)
LSFD	-0.0079 (-3.29)	1.1932 (2.32)	0.46 (0.22, 0.64)	0.44 (0.19, 0.69)
NBLS	-0.0043 (-3.29)	1.0121 (3.65)	0.19 (-0.04, 0.36)	0.08 (-0.02, 0.42)

LSTD means Least Squares in the time domain (5); LSFD is Least Squares in the Frequency Domain (6), and NBLS is Narrow Band Least Squares (7).

Table 4: FCVAR results for the growth rate series

	d-b	Cointegrating equation beta	
		Rate CO ₂	Rate GDP
a) Model for the growth rates of CO ₂ and GDP imposing $d=b$	d = 1.024 (0.092)	1.000	-2.662
	$\Delta^d \begin{pmatrix} \text{Rate CO}_2 \\ \text{Rate GDP} \end{pmatrix} - \begin{bmatrix} 0.002 \\ 0.009 \end{bmatrix} = L_d \begin{bmatrix} 0.001 \\ 0.057 \end{bmatrix} v_t + \sum_{i=1}^2 [\hat{\Gamma}_i \Delta^d L_d^i (X_t - \mu)] + \varepsilon_t$		
b) Model for the growth rates of CO ₂ and GDP not imposing $d=b$	d = 0.948 (0.106) b = 0.986 (0.000)	1.000	-2.735
	$\Delta^d \begin{pmatrix} \text{Rate CO}_2 \\ \text{Rate GDP} \end{pmatrix} - \begin{bmatrix} 0.002 \\ 0.009 \end{bmatrix} = L_d \begin{bmatrix} 0.005 \\ 0.049 \end{bmatrix} v_t + \sum_{i=1}^2 [\hat{\Gamma}_i \Delta^d L_d^i (X_t - \mu)] + \varepsilon_t$		