Trade Openness and Carbon Emissions: Evidence from Central and Eastern Europe

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Abstract

Using a composite trade share measure of trade openness, we examined the effects of trade openness on carbon dioxide emission for a sample of 17 Central and Eastern European (CEE) countries over the period 1994 to 2014. We found that high trade openness is associated with low carbon emission in the long run. When compared with the simple measure of trade openness (i.e. total trade as a percentage of GDP), the composite measure indicates that the effect of openness on carbon emission in the long run is smaller in absolute terms. Moreover, while high openness is associated with high emission in the short run using the simple measure, this association is non-existent when using the composite measure. These findings are robust to two historically closed economies and the recent Global Financial Crisis. When testing the Environmental Kuznets Curve (EKC) hypothesis using the composite measure, we found evidence in support of this in the long run. This finding connotes that high openness is associated with low emission in the long run, but up to a certain level of openness. That is, there is a turning point for openness beyond which further openness may spur high emission. Overall, our findings clearly suggest, to a large extent, that the measure of trade openness matters.

Keywords: Measuring Trade Openness; Carbon Emissions; Central and Eastern Europe.
JEL Codes: Q56

1. Introduction

The question of whether trade openness is beneficial or harmful to the environment remains elusive in the literature (Copeland & Taylor, 2005; Frankel & Rose, 2005; Chintrakarn & Millimet, 2006; Frankel, 2009). Hence, in order to push the debate towards consensus, further empirical exploration of the question at hand is required. In this paper, we attempt to answer the question by focusing on the CEE countries. These countries are historically in the Eastern bloc, which is located to the west of the post-World War II border with the former Soviet Union, the three Baltic States of Estonia, Latvia and Lithuania, and the independent states in the former Yugoslavia (World Bank, 2008; Iyke, 2017). What makes the CEE countries appealing for this study is that they were previously communist states, which have pursued drastic economic, political and institutional reforms over the last three decades. Virtually all of them have pursued trade liberalisation policies to make their economies accessible to the rest of the world. Therefore, it would be intriguing to know how such openness-oriented policies effect the environment during this period of transition.

Apart from the lack of consensus regarding the openness-environment nexus, an issue necessitating this paper is that previous studies have favoured simple outcome-oriented measures of trade openness in their empirical analysis. Although outcome-oriented measures are robust because they are relatively objective and the data for constructing them is publicly available at reputable sources such as the IMF and the World Bank, the simple outcome-oriented measures popularly used in various studies suffer from one limitation. As argued by Squalli and Wilson (2011) and later corroborated by Iyke (2017), they only account for a country's share in world trade. They fail to account for a country's interaction and interconnectedness with the rest of the world. This limitation implies that previous studies using simple outcome-oriented measures of openness could be enhanced by using better measures – in a quest to report fair size of the effects of openness on, for instance, carbon dioxide (CO_2) emission (Cole, 2004; Managi, Hibiki & Tsurumi, 2009; Omri, 2013; Al-Mulali, Saboori & Ozturk, 2015; Shahbaz, Kumar & Zakaria, 2017). The policy implications stemming from the openness-emission debate are critical (see Copeland & Taylor, 1995; 2005). Therefore, efforts to establish the fair size of the effects of openness on emission are certainly worthwhile.

To overcome the above limitation, this paper proposes that the openness-environment or openness-emission nexus should be tested using a recently developed composite trade share measure of openness that does not only account for a country's contribution to world trade, but also reflects its interaction and interconnectedness with the rest of the world. A change in this direction could enhance our understanding regarding the openness-emission nexus. The composite trade share measure of openness has been developed by Squalli and Wilson (2011) for a cross-section of countries. In this paper, we extend the idea of a panel data setting in order to gain from the rich cross-sectional and time series dynamics of panel data. Beyond constructing a panel composite trade share index in order to shed further light on the openness-emission relationship, we carefully identified a group of relatively homogenous countries (CEE countries) that have charted a common path to trade liberalisation contemporaneously. Then, we modelled the openness-emission nexus by using a distributed lag approach that allows us to restrict countries to a homogeneous long-run path, but that allows flexible heterogeneous short-run adjustments to the equilibrium.

As a preview of our findings, we found that, when measured by the composite trade share, high trade openness is associated with low carbon emission in the long run. As compared to the simple measure of trade openness widely used in the literature (i.e. the sum of exports and imports as a percentage of GDP), the new measure suggests that the effect on emission of openness in the long run is smaller in absolute terms. In addition, while high openness is associated with high emission in the short run using the simple measure, such an association is non-existent when using the new measure. Furthermore, these findings are robust to two historically closed economies and the recent Global Financial Crisis. These findings clearly suggest, to a large extent, that the measure of trade openness does matter. We push the empirical analysis a little further by testing the EKC hypothesis widely studied in the literature. We found evidence in support of an EKC in the relationship between openness and emission in the long run. This finding shows that high openness is associated with low emission in the long run, but only up to a certain level of openness. Explained differently, the finding implies that there is a turning point for openness beyond which further openness may spur high emission. We did not further explore the estimate of this turning point, but relegated it to future studies. Beyond showing the importance of accounting for the two dimensions of trade openness, our study shed light on the need for policymakers in the CEE countries to be aware of the consequences if the turning point of openness is superseded. At present, these countries are in a favourable position to embark on further openness-oriented policies at no cost of degrading environmental quality. To the extent that our study adds to the existing literature, our findings do not offer one-size-fits-all evidence of the openness-emission nexus for these countries. Further empirical explorations are encouraged in order to pin down a concrete relationship. If and when data becomes available, it would be intriguing to see how our findings compare with estimates at city level across the CEE countries.

By focusing on the measures of trade openness, our study adds to the literature on the definition, measures and implications of trade openness, including Krueger (1978), Leamer (1988), Anderson and Neary (1992), Harrison (1996), Pritchett (1996), Cavallo and Frankel (2008), Chang, Kaltani and Loayza (2009), Frankel (2009), Squalli and Wilson (2011), and Iyke

(2017). We examined the effects of openness on carbon dioxide emission. In this sense, our study contributes to the literature on the openness-emission nexus such as Copeland and Taylor (2005), Frankel and Rose (2005), Managi et al. (2009), Omri (2013), Al-Mulali et al. (2015), Shahbaz et al. (2017). In addition, we tested the EKC hypothesis. Hence, our study contributes to that literature (Cole, 2004; Managi & Jena, 2008; Jalil & Mahmud, 2009; Nasir & Rehman, 2011; Ozturk & Acaravci, 2013; Shahbaz et al., 2017, Shahzad et al., 2017). Finally, we differentiated the short-run effects from the long-run effects, thereby contributing to previous studies focusing on that aspect, including Catao and Solomou (2005), Catao and Terrones (2005), Frank (2009), Kim and Lin (2010), Chudik et al. (2017).

The rest of the paper is organised as follows. In the next section, we review the related literature on trade openness and carbon dioxide emission. In section 3, we discuss the measures of trade openness, the data and the empirical model. Section 4 presents the empirical results and section 5 concludes the paper.

2. Literature review

The pioneering studies on the EKC, such as studies done by Grossman and Krueger (1991), Shafik and Bandyopadhyay (1992), and Panayotou (1993) indicated an inverted U-relationship between environmental quality and income per capita. Various extensions of the EKC literature underscored international trade as one of the most important factors influencing environmental quality (Antweiler, Copeland & Taylor, 2001; Dinda, 2004; Cole, 2004; Frankel, 2009). According to Grossman and Krueger (1991), trade openness affects the environment through three channels: scale, technique and composition effects. The scale effect shows that trade is likely to increase pollution as more outputs and pollutants are produced due to an increase in market access and market activities (Dinda, 2004; Cole, 2004). However, the technique effect demonstrates that trade openness reduces pollutions (Martin & Wheeler, 1992). As technologies advance due to trade liberalisation, the obsolete and dirty production processes are replaced by cleaner ones, thereby improving environmental quality (Martin & Wheeler, 1992; Reppelin-Hill, 1999). Finally, the composition effect states that as the structure of an economy changes, the levels of pollution also vary (Grossman & Krueger, 1991). As a result of trade liberalisation, the structure of an economy changes according to the comparative advantage of economy. If an economy has a comparative advantage in pollution-intensive production, then trade would promote such production, thereby increasing the emission of pollutants. This argument is inherent in the Pollution Haven Hypothesis (PHH), which states that, following a reduction in trade barriers, firms in the pollution-intensive industry would move from countries with strong environmental regulations to countries with weak environmental regulations. In other words, weak environmental regulations become the source of a comparative advantage, which often occurs in developing countries (Dinda, 2004; Cole, 2004; Cherniwchan, Copeland & Taylor, 2017).

Trade could also induce emissions or hurt environmental quality through the so-called "race to the bottom" hypothesis. According to this hypothesis, reduction in trade barriers encourages multinational firms to relocate to the countries with lower environmental regulations. Such rising capital outflows would force the government to adopt less stringent environmental regulations to maintain international competitiveness. In essence, the adoption of less stringent environmental regulations owing to international trade increases environmental degradation (Jaffe, Peterson, Portney & Stavins, 1995; Dinda, 2004; Frankel, 2009).

The above theoretical arguments suggest that the relationship between free trade and pollutant emissions can be positive or negative. The theoretical ambiguity in trade-environment nexus is also in accord with the results of empirical studies. Some studies show that trade openness reduces pollutant emissions. For example, while examining the impact of trade openness on environmental quality, Managi (2004) found that there is a positive relationship between them in both developed and developing countries. Frankel and Rose (2005) explored the effect of trade on the environment and found that free trade reduces air pollutants. Recently, Ling, Ahmed, Muhamad & Shahbaz (2014) analysed the impact of trade openness on carbon dioxide emissions in Malaysia and concluded that trade lowers carbon dioxide emissions. In contrast, other studies found a positive association between trade openness and pollutant emissions. For example, Frankel (2009) found empirical evidence that trade could exacerbate environment degradation when it is measured by carbon dioxide emissions. Hossain (2011), while exploring the causal relationship of trade and carbon dioxide emissions for a panel of countries, found trade to cause carbon dioxide emissions in the short run. Other studies such as Takeda and Matsuura, (2006), Bombardini and Li (2016), and Shahbaz et al. (2017) also found a positive and significant association between trade openness and pollutant emissions.

Some studies show that the association between trade openness and pollutant emissions depends on the stage of economic development of the country. For example, while estimating the scale, technique and composition impacts of trade openness on air pollution, Antweiler et al. (2001) found that trade openness tends to worsen air quality in rich countries but improve air quality in poor countries. Contrary to the results of Antweiler et al. (2001), Cole (2004), using data on North-South trade flows for pollution-intensive products to test the Pollution Haven Hypothesis, concludes that trade increases the migration of pollution-intensive industries from the developed countries to the developing countries. Similarly, Managi et al. (2009) found that trade increases carbon dioxide emissions in non-OECD countries but lowers emissions in OECD countries. In addition, some studies conclude that trade openness does not have any significant impact on the environment (see, for example, Grossman & Krueger, 1991; Shafik, 1994; Copeland & Taylor, 2005; Soytas et al., 2007; Levinson, 2009; Jalil & Mahmud, 2009; Omri, 2013).

Some studies investigated the relationship between trade openness and carbon dioxide emissions by specifically focusing on transition economies. For example, Dean (2002) tested whether lenient environmental standards in China would worsen the quality of the environment as measured by water pollution. The results from her study suggested that freer trade in fact benefits the environment. When exploring the impact of the opening of 10 CEE economies to international trade on environmental quality at firm level during the period of 1990-1997, Andonova (2003) found that openness provides limited support in improving the environment. Al-Mulali et al. (2015) examine the influence of trade openness on pollution in 23 European countries, including the Czech Republic, Hungary, Poland, Romania, Slovak Republic and Slovenia during the period 1990 to 2013. Their results indicated that trade reduces pollutant emissions. Recently, Ahmed et al. (2016) examined the causal relationship between trade openness and carbon dioxide emissions in Brazil, India, China and South Africa. They found that trade openness in these countries induces higher emissions.

Our survey suggests that most of the existing studies employed simple outcome-oriented measures of trade openness, when examining the effects of trade on emissions or the environment. While some studies use the sum of exports and imports as percentage of GDP as a measure of trade openness (Frankel & Rose, 2005; Frankel, 2009; Hossain, 2011; Jalil & Mahmud, 2009; Omri, 2013), others use the volume of exports per capita and imports per capita

(Shahbaz et al., 2012; Ahmed et al., 2016, Shahbaz et al., 2017). These simple outcome-oriented measures of trade openness are generally limited because they fail to account for a country's interaction and interconnectedness with the rest of the world. In this paper, we overcome this limitation by extending a recently developed trade openness measure to the panel data setting. We then employed a flexible model to distinguish short-run effects from long-run effects. Our paper thus offers a fresh perspective of an old debate.

3. Methods and data

3.1. Measuring trade openness

The concept of trade openness has been widely discussed in the trade literature and varies from one author to the other. For example, while Krueger (1978) defines trade openness as the pursuance of favourable export-oriented policies by an economy, Anderson and Neary (1992) argue that trade openness indicates the degree of distortion of an economy due to tariff and nontariff barriers. Harrison (1996) links trade openness to the degree of neutrality of the incentives between savings from imports and earnings from exports. Others, including Leamer (1988) and Pritchett (1996), define trade openness as a measure of the trade intensity of an economy.

Clearly, the idea of trade openness can be measured using policies or trade outcomes. In application, various policy-based measures of openness have been developed (Edwards, 1998; Lee et al., 2004). Nevertheless, the policy-based measures are subjective and therefore compromise the empirical results. In other words, these measures are influenced by the researcher's prior knowledge regarding the meaning of openness. Studies using the policy-based measures have received strong criticism. The proposed alternative is the outcome-based measure of openness although several studies have used different outcome-based openness measures (Dollar & Kraay, 2003; Yanikkaya, 2003; Alcalá & Ciccone, 2004; Cavallo & Frankel, 2008; Chang et al., 2009; Frankel, 2009). The main advantage of the outcome-based openness measures is that they are objectively constructed using publicly available trade data. The most commonly employed outcome-based measure of openness is the trade intensity ratio or share (TS), calculated as the sum of exports and imports, divided by GDP (Leamer, 1988; Chang et al., 2009). This is defined mathematically as, $X + M/GDP$, where X, M and GDP denote exports, imports and gross domestic product, respectively. The World Bank compiles this measure annually.

The disadvantage of the existing outcome-based measures of openness, including the simple trade intensity ratio (TS), is that they only capture one dimension of trade openness: a country's share in world trade. They do not account for the advantages enjoyed by a country due to its interaction and interconnectedness with the rest of the world (Squalli & Wilson 2011; Iyke, 2017). A better measure of openness should account for both a country's share of trade, interaction and interconnectedness with the rest of the world. Squalli and Wilson (2011) developed a measure of openness that takes into account both dimensions of trade in a cross-sectional setting. This measure has been extended to a panel setting in order to capture cross-sectional and time dynamics by Iyke (2017). In this paper, we follow the latter paper in constructing a measure of openness developed by Squalli and Wilson (2011). The new measure of openness is calculated as follows:

$$CTS_i = \frac{(X+M)_i}{\frac{1}{n}\sum_{j=1}^{n}(X+M)_j}\frac{(X+M)_i}{GDP_i} \quad (1)$$

where CTS is the composite trade share defined as TS adjusted by the proportion of a country's trade relative to the average world trade; $TS_i = (X+M)_i/GDP_i$; i is a given country which belongs to *j* a set of countries {1, …., n}. By definition,

$$WTS_i = \frac{(X+M)_i}{\sum_{j=1}^{n}(X+M)_j}. \quad (2)$$

$nWTS_i > 1$ if a country is a major contributor to world trade and its trade is higher than the world average. In this case, TS_i must adjust upwards. The main difference between the CTS and TS is that the former penalises smaller countries, while the latter penalises larger countries (Squalli & Wilson, 2011).[1] In our empirical application, we calculated the CTS for each country for all years used in the sample.

3.2. Empirical model

To examine the effects of trade openness on carbon emissions within the CEE, we follow the recent studies (e.g. Cole, 2004; Managi et al., 2009; Omri, 2013; Al-Mulali et al., 2015; Shahbaz et al., 2017) and specify the following empirical model:

$$lnCO_{it} = \alpha_0 + \alpha_1 lnY_{it} + \alpha_2 lnTR_{it} + \epsilon_{it} \quad (3)$$

where CO, Y and TO denote carbon dioxide emissions, real income and trade openness, respectively; ln is the natural logarithm operator; αs are the parameters to be estimated; ϵ denotes the *iid* error term; i and t denote the cross-sectional and time subscripts, respectively.

The limitation of the empirical model in Eq. (3) is that it does not take into account the short-run dynamic movements of the variables. Hence, it is impossible for the researcher to evaluate the short-run effects of openness on carbon emissions. The researcher is only able to assess the long-run effects. To obtain the short-run dynamics, we first need to recast Eq. (3) into a distributed lag model of the form:

$$y_{it} = \mu_i + \sum_{j=1}^{p}\lambda_{ij}\,\Delta y_{it-j} + \sum_{j=0}^{q}\delta'_{ij}\,X_{it-j} + \varepsilon_{it}. \quad (4)$$

By suitably reparametrizing Eq. (4), we obtain the following error-correction model:

$$\Delta y_{it} = \mu_i + \phi_i(y_{it-1} - \theta'_i X_{it}) + \sum_{j=1}^{p-1}\lambda^*_{ij}\,\Delta y_{it-j} + \sum_{j=0}^{q-1}\delta'^*_{ij}\,\Delta X_{it-j} + \varepsilon_{it} \quad (5)$$

where y and X are the dependent variable (carbon dioxide emissions, CO) and the explanatory variables (real income, Y; openness, TO), respectively; μ and ε denote the individual fixed

[1] Further details are provided in Squalli and Wilson (2011) and Iyke (2017).

effects and the *iid* error term, respectively; λ_{ij} and δ_{ij} denote scalars and coefficient vectors, respectively. Furthermore, $\phi_i = -\left(1-\sum_{j=1}^{p}\lambda_{ij}\right)$; $\theta_i = \sum_{j=0}^{q}\delta_{ij}/(1-\sum_k \lambda_{ik})$; $\lambda_{ij}^* = -\sum_{m=j+1}^{p}\lambda_{im}$, $j = 1, 2, \ldots, p-1$; $\delta_{ij}^* = -\sum_{m=j+1}^{q}\delta_{im}$, $j = 1, 2, \ldots, q-1$. ϕ_i denote the error-correction term, indicating the rate of adjustment of the variables to equilibrium whenever they depart from it. The evidence of cointegration or long-run relationship among the variables is supported if the estimated value of ϕ_i is negative and statistically significant. Finally, θ_i' denotes the cointegrating vector, showing the number of cointegration relationships in the model.

The carbon dioxide emissions model in Eq. (5) offers various policy insights, thereby making it suitable for empirical analysis. Firstly, the policymaker is able to differentiate the short-run effects from the long-run effects of trade openness on carbon dioxide emissions. Secondly, the policymaker can model the persistence and the adjustment to equilibrium paths of openness and carbon emissions. Thirdly, contemporaneous feedback causality, which often bias empirical estimates, is taken into consideration in the specification. Finally, the policymaker is able to model cross-sectional heterogeneities in the emissions-openness relationship by allowing the parameters in specification vary. Various studies offer similar explanations (Catao & Solomou, 2005; Catao & Terrones, 2005; Frank, 2009; Kim & Lin, 2010; Chudik et al., 2017).

The carbon dioxide emissions model in Eq. (5) can be estimated by using three popular estimators. The first two are extreme cases, while the last is the intermediate of the two. Supposing only the intercept parameters are heterogeneous, then we can estimate the model using the dynamic fixed-effects (DFE) estimator. However, if we assume that the parameters are heterogeneous across countries, then the mean group (MG) estimator proposed by Pesaran and Smith (1995) is preferred for estimating the model. If, instead, we assume that the intercept, short-run coefficients, and the error terms vary but the long-run coefficients are the same across countries, then the preferred estimator is the pooled mean group (PMG) estimator developed by Pesaran et al. (1999). In most empirical applications, the PMG estimator is preferred because it combines the pooling advantages of the DFE estimator and the averaging advantages of the MG estimator. In fact, Pesaran et al. (1999) have shown that, due to its flexibility, the PMG estimator performs better than both the MG and DFE estimators.[2]

To avoid pitfalls, it is natural to report results for all three and test estimator preference using the standard Hausman test (Pesaran et al., 1999). In our analysis – although we prefer the PMG estimator because it allows us to model a common cross-sectional long-run relationship between carbon emissions and openness vis-à-vis estimating short-run heterogeneous adjustments of the markets to equilibrium across countries – we report results for all three and adjudged the best results using the Hausman test.

3.3. Data

Table 1 shows the definitions of the variables and their sources, while table 2 shows the descriptive statistics. The dependent variable is carbon dioxide emissions (CO_2 in metric tons per capita). The independent variables are trade openness (TO) and real income. These variables appear in most of the literature on the emissions-openness nexus (see, example, Omri,

[2] If there is cross-sectional variation of the slope coefficients, the DFE estimator yields inconsistent results. Similarly, if the long-run coefficients are homogeneous, the MG estimator yields inconsistent results. The PMG estimator yields consistent results in both cases (see Pesaran et al., 1999).

2013; Shahbaz et al., 2017). We covered 17 CEE countries[3] for the period 1994 to 2014. This was the longest sample span available at the time we carried out the study. Because observations on CO_2 are missing for Montenegro and Serbia in some years, the dataset is unbalanced. However, this has an immaterial effect on the empirical results.

Table 1: List of variables and their sources

Variable	Name	Source
CO	CO_2 emissions (metric tons per capita)	World Bank national accounts data and OECD National Accounts data files.
Y	GDP per capita (constant 2005 US$)	World Bank national accounts data and OECD National Accounts data files.
X	Exports of goods and services (constant 2005 US$)	World Bank national accounts data and OECD National Accounts data files.
M	Imports of goods and services (constant 2005 US$)	World Bank national accounts data and OECD National Accounts data files.
GDP	GDP (current US$)	World Bank national accounts data and OECD National Accounts data files.
TS	Trade (% of GDP)	World Bank national accounts data and OECD National Accounts data files.
WTS	World Trade Share	Calculated as in *Eq.* (2) using data from World Bank national accounts data and OECD National Accounts data files.
CTS	Composite Trade Share	Calculated as in *Eq.* (1) using data from World Bank national accounts data and OECD National Accounts data files.

Note: X, M, GDP and TS were used in calculating the trade openness measures discussed above.

Table 2: Descriptive Statistics

Variable	Mean	Standard Deviation	Minimum	Maximum
lnCO	1.5588	0.6883	-0.7206	2.7154
lnY	8.6379	0.7362	6.2534	9.9517
lnTS	4.5470	0.3386	3.1448	5.2118
lnCTS	8.4848	1.3514	5.1547	10.7787

Note: ln denotes the natural logarithm operator.

3 These countries are Albania, Bosnia and Herzegovina, Bulgaria, Croatia, Czech Republic, Estonia, Georgia, Hungary, Latvia, Lithuania, Macedonia FYR, Montenegro, Poland, Romania, Serbia, Slovak Republic and Slovenia.

4. Empirical results

4.1. The basic results

We begin our empirical analysis by presenting the basic results. These are the coefficient estimates obtained when trade openness is measured as a simple trade intensity ratio (TS). Most studies on the openness-emissions nexus have used this measure (Omri, 2013; Al-Mulali et al., 2015; Shahbaz et al., 2017). Table 3 shows these results. The inclusion of lags is an important task when estimating the error-correction framework discussed above. Studies such as those done by Kim and Lin (2010), and Loayza and Ranciere (2006) explained that if our interest is in the long-run parameters, the optimal lags in the model should be chosen based on consistent information criteria on a country-by-country basis. In contrast, they argued, if we are interested in both the short- and long-run parameters, it is optimal to impose a common lag order across countries. In our empirical analysis, we are interested in both the short- and long-run coefficients. Hence, we imposed a common lag across countries. Specifically, we imposed a maximum lag of one in order to avoid over-specification, since the data is annual. To avoid bias against any of the estimators, we reported the estimates for the PMG, MG and DFE estimators and established the best estimator using the Hausman test.

The results indicated that the error-correction term is negative, significant and lower than unity in absolute terms, for all three estimates. We can conclude that the variables are cointegrated or share a stable long-run relationship. This means that the variables tend to move closely in the long run if they drift apart in the short run. In addition, the Hausman test suggests that the PMG estimator is the best estimator, since the p-values are considerably larger than the conventional significance levels (i.e. 1%, 5% and 10%).[4] Now, focusing on the results based on the PMG estimator, we find that openness is associated with high emissions in the short run. However, this positive effect of openness on emissions becomes negative in the long run. These findings appear to corroborate the conflicting literature (Cole, 2004; Managi et al., 2009; Shahbaz et al., 2017). A caveat of these results is that the simple trade intensity ratio measure of trade openness is limited. This measure only accounts for a country's share in world trade but fails to account for the advantages enjoyed by a country due to its interaction and interconnectedness with the rest of the world (Squalli & Wilson 2011; Iyke, 2017). In other words, the results may not be a fair reflection of the effects of openness on emissions within these countries. We would address this issue next. Looking at the results for real income, it is evident that the level of emissions rises with rising income both in the short and long run. Alternatively, richer countries in the CEE appear to emit more carbon dioxide. Although these results are consistent with the literature, this conclusion may not necessarily be the entire picture. Hence, our results should be interpreted with caution.

Table 3: The Basic Results

Variable	PMG	MG	DFE
Long-run Estimates			
lnY	0.3584	0.3771	0.9653
	(0.0000)	(0.0520)	(0.0000)
lnTS	-0.4050	-0.2449	-0.8707
	(0.0000)	(0.1660)	(0.0000)
Short-run Estimates			

[4] The rest of the paper reports results based on the PMG estimator.

ECT	-0.3249	-0.5847	-0.1991
	(0.0000)	(0.0000)	(0.0040)
ΔlnY	0.8394	0.7620	0.5732
	(0.0000)	(0.0000)	(0.0020)
ΔlnTS	0.1758	0.1369	0.2319
	(0.0130)	(0.0560)	(0.0310)
Constant	0.0913	-0.6748	
	(0.0500)	(0.4990)	
Hausman Test ($\chi2$)		0.3900	0.1200
P-value		(0.8240)	(0.9433)

Notes: P-values are in the parentheses; Δ is the first difference operator; ECT is the error-correction term.

4.2. Comparing the basic results to the new results

As pointed out above, the limitation of the simple trade intensity ratio as a measure of openness connotes that the coefficient estimates may not reflect the appropriate effect of openness on carbon emissions. A good measure of openness should account for both a country's share of trade, interaction and interconnectedness with the rest of the world. The composite trade share or ratio measures these dimensions of openness. Hence, we improve upon previous studies by employing the composite trade share measure openness developed by Squalli and Wilson (2011) for cross-sectional data. However, since our data is longitudinal, we followed Iyke (2017) and constructed the longitudinal equivalent. Further details are presented earlier. The results, using the composite trade share (CTS), are shown in table 4. In order to highlight the relevance of the CTS, we reported the basic results as well. From the results, it is clear that there exists a stable long-run relationship among the variables. This is because the error-correction term is negative, statistically significant and smaller than unity in absolute sense.

Considering the coefficient estimates, we can see that openness has no significant effect on carbon emissions in the short run if openness is measured by the composite trade share. This opposes the short-run effect of openness on emissions using the simple trade share. Similarly, the long-run effect of openness on emissions is negative using the composite trade share, the size of the effect is smaller when compared with the estimate using the simple trade share. This evidence substantiates our earlier claim that care must be taken when interpreting the basic results. Clearly, the measure of trade openness is critical when examining the effects of openness on carbon emissions. We may conclude that openness is associated with declining carbon emissions in the long run. Although the opposite is the case in the short run, the effect is not significant in statistical terms.

Table 4: Comparing the Basic Results to the New Results

Variable	Basic Results	New Results
Long-run Estimates		
lnY	0.3584	0.1607
	(0.0000)	(0.0010)
lnTS	-0.4050	
	(0.0000)	
lnCTS		-0.1931
		(0.0000)

Short-run Estimates		
ECT	-0.3249	-0.3218
	(0.0000)	(0.0000)
ΔlnY	0.8394	0.8257
	(0.0000)	(0.0000)
ΔlnTS	0.1758	
	(0.0130)	
ΔlnCTS		0.0604
		0.2370
Constant	0.0913	0.5914
	(0.0500)	(0.0000)

Notes: P-values are in the parentheses; Δ is the first difference operator; ECT is the error-correction term.

4.3. The effects of historically closed countries

The results reported so far hinge on the assumption that the CEE countries are all open. Could this assumption be driving the results? Quite recent studies have documented that Croatia and Estonia are closed countries (Wacziarg & Welch, 2003; 2008; Nannicini & Billmeier, 2011). Therefore, it would be empirically useful to see what happens to our results when these countries are removed from the sample. Table 5 shows these results precisely. For comparison purposes, we also reported the results based on the full sample. A careful look at these results showed that the two closed countries do not drive the results significantly. The findings are essentially the same as the previous ones. That is, openness is associated with declining carbon emission in the long run. In the short run, although openness enhances carbon emission, the effect is statistically insignificant.

Table 5: Controlling for Historically Closed Countries

	Basic Results		New Results	
Variable	All Countries	Open Countries	All Countries	Open Countries
Long-run				
lnY	0.3584	0.4818	0.1607	0.2568
	(0.0000)	(0.0000)	(0.0010)	(0.0000)
lnTS	-0.4050	-0.5338		
	(0.0000)	(0.0000)		
lnCTS			-0.1931	-0.2943
			(0.0000)	(0.0000)
Short-run				
ECT	-0.3249	-0.3202	-0.3218	-0.3064
	(0.0000)	(0.0000)	(0.0000)	(0.0000)
ΔlnY	0.8394	0.8761	0.8257	0.8895
	(0.0000)	(0.0000)	(0.0000)	(0.0000)
ΔlnTS	0.1758	0.1986		
	(0.0130)	(0.0070)		
ΔlnCTS			0.0604	0.0635
			(0.2370)	(0.2970)
Constant	0.0913	-0.0864	0.5914	0.5592
	(0.0500)	(0.0360)	(0.0000)	(0.0000)

Notes: P-values are in the parentheses; Δ is the first difference operator; ECT is the error-correction term.

4.4. The effects of the global financial crisis

The recent Global Financial Crisis (GFC) has affected aggregate demand and therefore factors of production. It is possible that general emissions from plants and overall energy usage may have been negatively affected during the peak of the crisis. Moreover, the volume and frequency of trade across countries may have been negatively affected as well (Milesi-Ferretti & Tille, 2011; Anderson & Nelgen, 2012). The converse of the response of these variables is also plausible (Peters et al., 2012). Specifically, the GFC was an extreme event that has shaped many variables, including the levels of carbon emissions and the degree of trade openness across countries. In that sense, could GFC be driving our results? To present clean estimates, it is relevant to remove the GFC effect. The full-scale GFC was felt during 2008 (Chor & Manova, 2012; Fratzscher, 2012; Peters et al., 2012). Therefore, we excluded 2008 from the sample and re-estimated the empirical model. Table 6 shows these results. Again, for comparison purposes, we reported the results for the full sample as well. From these results, it is evident that the crisis has little effect on our findings.

Table 6: Controlling for the Recent Global Financial Crisis

	Basic Results		New Results	
Variable	Crisis	No Crisis	Crisis	No Crisis
Long-run				
lnY	0.3584	0.3579	0.1607	0.1495
	(0.0000)	(0.0000)	(0.0010)	(0.0020)
lnTS	-0.4050	-0.4183		
	(0.0000)	(0.0000)		
lnCTS			-0.1931	-0.1961
			(0.0000)	(0.0000)
Short-run				
ECT	-0.3249	-0.3327	-0.3218	-0.3172
	(0.0000)	(0.0000)	(0.0000)	(0.0000)
ΔlnY	0.8394	0.8019	0.8257	0.7677
	(0.0000)	(0.0000)	(0.0000)	(0.0000)
ΔlnTS	0.1758	0.1903		
	(0.0130)	(0.0090)		
ΔlnCTS			0.0604	0.0739
			(0.2370)	(0.1600)
Constant	0.0913	0.1187	0.5914	0.6166
	(0.0500)	(0.0170)	(0.0000)	(0.0000)

Notes: P-values are in the parentheses; Δ is the first difference operator; ECT is the error-correction term.

4.5. Testing the Environmental Kuznets Hypothesis

One popular hypothesis tested in the openness-emission literature is the Environmental Kuznets Hypothesis (EKH) (Cole, 2004; Managi et al., 2009; Shahbaz et al., 2012; Lau et al., 2014). Researchers want to establish whether there is an EKC in the openness-emission nexus. The evidence of EKC suggests that the relationship between openness and emission is nonlinear. In other words, there is a given range of openness, for which openness is negatively related to emission, and a range for which the relationship is positive. Establishing such a relationship has important policy implications. The most prominent one being that policymakers would be able to estimate an optimal level (or turning point) of trade openness (Shahbaz et al., 2017), beyond (below) which openness would be undesirable (desirable).

Previous studies have tested the EKH using the simple trade share measure. The limitation of this measure of openness has already been discussed. Hence, it is useful to test the EKH using the new measure of openness, the composite trade share. We take our analysis a bit further by re-testing the EKH using the composite trade share measure of openness. Table 7 shows these results. We reported the basic results for comparison purposes. The results show evidence in support of the EKH. Note that we documented a negative effect of openness on emissions in the long run. Hence, ideally, if there is an EKC, it should be a normal U-curve. This is unlike the studies documenting a positive effect of openness on emissions, thereby establishing the evidence of an inverted U-curve (Ozturk & Acaravci, 2013; Shahbaz et al., 2017; Shahzad et al., 2017). Both the basic and new results support evidence of EKC with a normal U-shape in the long run. The EKC vanishes in the short run. Our findings imply that there is a turning point beyond which openness may enhance emission. This turning point is not of interest to us. We leave that for future studies.

Table 7: The Evidence of Environmental Kuznets Curve

Variable	Basic Results	New Results
Long-run Estimates		
lnY	0.9536	0.5141
	(0.0000)	(0.0000)
lnTS	-7.7246	
	(0.0000)	
$\ln TS^2$	0.8537	
	(0.0000)	
lnCTS		-1.3594
		(0.0000)
$\ln CTS^2$		0.0943
		(0.0000)
Short-run Estimates		
ECT	-0.2265	-0.3089
	(0.0240)	(0.0000)
ΔlnY	0.5907	0.6347
	(0.0010)	(0.0000)
ΔlnTS	1.8105	
	(0.6190)	
$\Delta\ln TS^2$	-0.1455	
	(0.7090)	
ΔlnCTS		-0.7204
		(0.7180)
$\Delta\ln CTS^2$		0.0479
		(0.6820)
Constant	2.4471	-2.3154
	(0.0260)	(0.0000)

Notes: P-values are in the parentheses; Δ is the first difference operator; ECT is the error-correction term.

5. Conclusion

In this paper, we studied the effects of trade openness on carbon dioxide (CO_2) emission in the CEE countries. Outcome-based measures of trade openness are very popular in the literature, mainly because they are relatively objective. Besides, the data for constructing them are publicly available at reputable sources such as the IMF and the World Bank. So far, all the available outcome-based measures of openness have the limitation that they only capture the share of a country's trade with the rest of the world. They fail to account for a country's interaction and interconnectedness with the world economy. Owing to this limitation, it is plausible that previous studies are not estimating the appropriate size of the effects of openness on carbon emission. The policy implications stemming from the openness-emission debate are critical. Therefore, efforts to establish the fair size of the effects of openness on emission are certainly worthwhile.

Recently, a composite trade share measure of trade openness has been developed to account for the two dimensions of openness, namely a country's contribution to world trade and its interaction and interconnectedness with the rest of the world. Using this new measure would definitely enhance our understanding regarding the openness-emission nexus. The new measure of openness is obtained for a cross-section of countries by its proponents. In this paper, we extend the idea to a panel data setting in order to gain from the rich cross-sectional and time series dynamics of panel data. We carefully identified a group of relatively homogenous countries that have charted a common path to trade liberalisation contemporaneously – the CEE countries. We then modelled the openness-emission nexus by using a distributed lag approach that allowed us to restrict countries to a homogeneous long-run path but that also allowed flexible heterogeneous short-run adjustments to equilibrium.

By using the new measure of trade openness, we found that low carbon emission is associated with high openness in the long run. When compared with the simple measure of trade openness popularly used in the literature (i.e. the sum of exports and imports as a percentage of GDP), the new measure shows that the effect on emission of openness in the long run is smaller in absolute terms. Furthermore, while high emission tends to be associated with high trade openness in the short run using the simple measure of openness, such an association is non-existent using the new measure. Generally, our findings are in line with the existing literature. In order to ensure that our findings are not driven by other factors, we explored a couple of options. Firstly, we controlled for two countries noted in various studies to be historically closed. Following this estimation, the results remained significantly unaffected. Secondly, we controlled for the recent GFC. In this case, our results also came out robust. These findings clearly suggested that future studies should be cautious when exploring the openness-emission nexus. To a large extent, the measure of trade openness matters.

We took our empirical analysis a little further by testing the EKH widely investigated in the literature. We found evidence in support of an EKC in the relationship between openness and emission in the long run. In particular, we found that high openness is associated with low emission in the long run, but up to a certain level of openness. In other words, there is a turning point for openness beyond which further openness may spur high emission. Although, one would be interested in knowing that particular turning point, estimating it is beyond this study. We leave that for future studies.

Beyond showing the relevance of capturing all dimensions of trade openness, our study shed light on policy options for the CEE countries. In negotiating further trade liberalisation, policymakers in these countries should also be aware of the consequences if the turning point of openness is superseded. For now, these countries are in a favourable position to embark on further openness-oriented policies at no cost of degrading the environmental quality. Of course, our findings do not offer one-size-fits-all evidence on the openness-emission nexus for these countries. Further empirical explorations are encouraged in order to pin down a concrete relationship. If and when data becomes available, it would be intriguing to see how our findings compare with estimates at the city level across the CEE countries.

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