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ANALYSIS

The environmental consequences of globalization: A country-specific time-series analysis

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ABSTRACT

The dynamic relationships among trade, income and the environment for developed and developing countries are examined using a cointegration analysis. Results suggest that trade and income growth tend to increase environmental quality in developed countries, whereas they have detrimental effects on environmental quality in most developing countries. It is also found that for developed countries, the causal relationship appears to run from trade and income to the environment — a change in trade and income growth causes a consequent change in environmental quality. For most developing countries, on the other hand, the causality is found to run from the environment to trade and income; however, the opposite causal relationship holds for China.

1. Introduction

Economists have vigorously debated the environmental consequences of trade liberalization (Copeland and Taylor, 1994, 2004; Copeland, 2005). Proponents of trade liberalization argue that, because environmental quality is a normal good, trade-induced income growth causes people to increase their demand for a clean environment, which in turn encourages firms to shift towards cleaner techniques of production.[1] Thus, free trade provides a win–win situation in the sense that it improves both environment and economy. Opponents of globalization, on the other hand, fear that if production techniques do not change, then environmental quality must deteriorate as trade increases the scale of economic activity. Moreover, if environmental quality is a normal good, then developing economies tend to adopt looser standards of environmental regulations. Given inequalities in the world distribution of income, trade liberalization may lead to more growth of pollution-intensive industries in developing countries as developed countries enforce strict environmental regulations, thereby having a significant adverse effect on environmental quality.

Since the seminal work by Grossman and Krueger (1991), many scholars have attempted to examine the effect of trade openness on the environment.[2] For example, Lucas et al. (1992) investigate the influence of trade openness on the growth rate of the toxic intensity of output. They find that among rapidly growing economies, increased trade openness reduces the growth rate of toxic intensity of output. Gale and Mendez (1998) analyze the relationship between trade, income growth and the environment; they find that an increase in income has a detrimental effect on environmental quality, but effect of trade liberalization on pollution is not significant. Dean (2002) examines the effect of trade liberalization on environmental damage. She finds that increased openness to international markets aggravates environmental damage through the terms of trade, but mitigates it through income growth. More recently, Frankel and Rose (2005) estimate the effect of trade on the environment for a given level of income per capita; they conclude that there is little evidence that trade openness causes significant environmental degradation.

Previous studies have undoubtedly expanded our understanding of the environmental consequences of economic growth and international trade. However, earlier studies have mostly adopted a reduced-form model to examine the presence of significant statistical associations for trade openness and income growth with environmental

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[1] It should be noted that economists have long argued that environmental quality is a luxury good; that is, income elasticity of environmental quality demand is in excess of unity. However, empirical studies have mostly found that demand for environmental quality indeed increases with income (mainly concerning health related pollutants), but not as a luxury good — income elasticity is smaller than 1 (e.g., Flores and Carson, 1997; Kristrom and Riera, 1996).

[2] Grossman and Krueger (1991) investigate the environmental impacts of the North American Free Trade Agreement (NAFTA) in an NBER working paper, which was later published in 1993 (Grossman and Krueger, 1993), and yield two novel results; (1) environmental quality first deteriorates and then improves with per capita income, which is known as the Environmental Kuznets Curve (EKC), and (2) trade liberalization tends to improve environmental quality via income growth.

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quality. Previous studies have paid little attention to the causal relationship among trade liberalization, income growth, and environment quality (Coondoo and Dinda, 2002; Chintrakarn and Millimet, 2006). More specifically, with the treatment of trade and income as being exogenous variables in a reduced-form model, past studies specify measures of environmental quality/damage (e.g., sulfur dioxide and carbon dioxide emissions) as a function of trade openness (usually defined as the sum of exports and imports divided by GDP) and income (usually per capita GDP). This approach implicitly assumes a unidirectional causal relationship; that is, a change in the level of trade openness and income causes a consequent change in the environmental quality, but the reverse does not hold. Accordingly, this presumption neglects the possibility of endogeneity of trade and income in the model. Because environmental quality and income may jointly (simultaneously) affect trade, causality could run in other directions (Frankel and Rose, 2005; Chintrakarn and Millimet, 2006). For example, trade can improve environmental quality via income growth, whereas strict environmental regulations can induce efficiency and encourage innovations, which may eventually positively affect a firm's competitiveness and thus trade volume, which is known as the Porter hypothesis (Porter and van der Linde, 1995). In addition, previous studies have typically used cross-section or panel data of a group of countries for their analyses. This approach assumes that a single country's experience (for example, economic development trajectory) over time would mirror the pattern revealed by a group of countries at different stages of development at a point in time (Dean, 2002; Coondoo and Dinda, 2002). However, considering wide cross-country variations observed in social, economical and political factors, the time path for individual countries may not follow a pattern of a group of countries.

Given the time-series properties of datasets on measures of economic activity, such as income and trade, and corresponding environmental change, a multivariate time-series analysis such as a vector autoregression (VAR) model is well suited to deal with the issue of endogeneity problem and/or causal mechanisms. More specifically, the VAR approach can determine both the short- and long-run dynamic effects of selected variables and test the endogeneity of them. One can interpret the results of this procedure as indicating potential impacts of shocks in a variable on all other variables. Compared to a reduced-form equation, therefore, the VAR approach allows us to address the endogeneity of income and trade, as well as to identify presence and direction of causality among variables without *a priori* theoretical structure. No previous study has attempted to directly address the potential endogeneity of income, trade and the environment with individual country-specific data and time-series models.[3]

In this paper, we use the Johansen cointegration analysis to examine the dynamic effect of trade liberalization on the environment using a time-series dataset of sulfur emissions (SO_2), income and trade openness for 50 individual countries over the last five decades. The Johansen approach features multivariate autoregression and maximum likelihood estimation; it is a convenient tool to examine dynamic interactions when variables used in the model are non-stationary and cointegrated. In addition, the cointegration approach is used to find the long-run equilibrium relationships among the selected variables. Given that the environmental consequences of income growth and liberalized trade are essentially a long-run concept (Dinda and Coondoo, 2006), using the cointegration method is indeed desirable to examine the true relationship among the environment, trade and income. Moreover, coefficients of the long-run relationships can be tested to determine whether any variable can be treated as a weakly exogenous variable, which is thus interpreted as a driving variable that influences the long-run movements of the other variables but is not affected by the other variables in the model. Hence, these dynamic interactions will provide an explanation for the causal mechanism among the selected variables. The remaining sections present the theoretical framework, empirical methodology, empirical findings, and conclusions.

2. A theoretical framework

To illustrate theoretical relationships among trade openness, (per capita) income and environmental quality, we first define income (Y) as a function of trade openness (T) and other exogenous variables (Z_1) as follows:

$$Y = f(T, Z_1) \tag{1}$$

Trade openness leads to an increase in the scale of economic activity and, consequently, an increase in income in a country. Trade openness thus shows a positive monotonic relation with income ($\partial Y/\partial T > 0$) as shown in the first-quadrant of Fig. 1. If trade openness does not positively affect an increase in income, it is hypothesized that the country may not involve in globalization via the World Trade Organization (WTO) and other regional and/or bilateral trade treaties. We then define pollution (E) as a function of income (Y) and production technology (Z_2) as follows:

$$E = f(Y, Z_2) \tag{2}$$

Pollution levels increase with growing income up to a threshold level (turning point) beyond which pollution levels decrease with higher income levels. The combination of these two effects ($\partial E/\partial Y > 0$ and $\partial E/\partial Y < 0$) creates the inverted U-shaped relationship between income and pollution levels as shown in the second-quadrant of Fig. 1; economists call this relationship as the Environmental Kuznets Curve (EKC). Finally, we substitute Eq. (1) into Eq. (2), which yields the following relationship:

$$E = g(T, Z_1, Z_2) \tag{3}$$

The relationship between pollution and trade openness depends on the relationships derived from Eqs. (1) and (2). For example, trade openness leads to an increase in real income and brings about a proportionate increase in pollution levels in early stages of economic development.

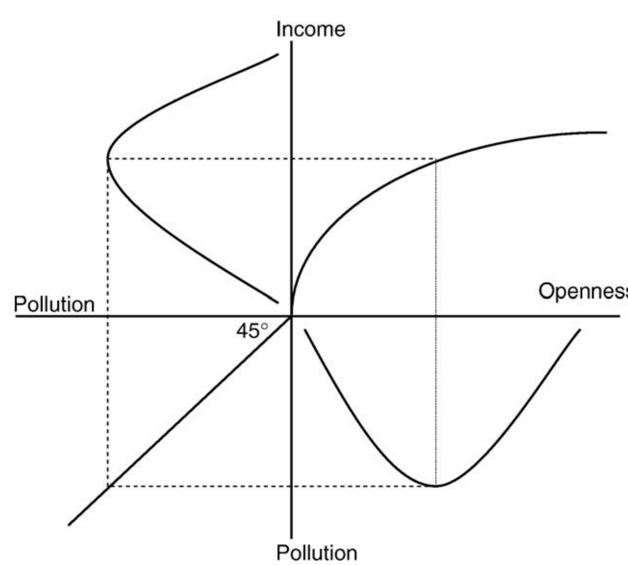

Fig. 1. Effect of trade openness on income and the environment.

[3] Frankel and Rose (2005) have directly addressed the endogeneity problem between trade, income and environmental quality in their analysis. However, they use the instrumental variable (IV) estimates based on cross-section data. On the other hand, some studies (for example, Coondoo and Dinda, 2002) have adopted time-series econometric techniques (e.g., Granger causality test and bivariate cointegration analysis) to examine causal relationship only between income and the environment.

However, because environmental quality is a normal good, an increase in income induced by trade causes people to increase their demand for a clean environment and eventually improves environmental quality. Consequently, a country tends to follow a pattern of rising pollution levels as trade openness proceeds ($\partial E/\partial T>0$), followed by declining pollution levels at more advanced stages of trade liberalization ($\partial E/\partial T<0$) (fourth-quadrant of Fig. 1). Therefore, the observed correlation between pollution and openness is a U-shaped curve as shown in the fourth-quadrant of Fig. 1.

Because individual countries experience different levels of income and openness corresponding to their process of development, the true form of the pollution–income–openness relationship mainly depends on where an economy is currently placed in a development trajectory. For example, as for individual countries that move beyond the EKC curve turning points and have higher degrees of openness, trade-induced increase in income may result in structural change towards less polluting industries, increased environmental awareness, and enforcement of environmental regulations, thereby improving environmental quality. On the other hand, if individual countries have not reached income levels high enough to be able to reach their turning points, trade liberalization may lead to more rapid growth of pollution-intensive industries in those countries and deterioration of environmental quality. As a result, the effect of trade liberalization on income and the environment is essentially an empirical question and varies according to circumstances such as individual countries' stages of development and openness levels.

3. Empirical methodology

3.1. Development of empirical time-series models

We use the cointegrated vector autoregression (CVAR) model developed by Johansen to examine dynamic interrelationship between trade, income and SO_2 emissions (Johansen, 1988; Johansen and Juselius, 1992). The Johansen method uses a statistical model involving up to k lags as follows:

$$\boldsymbol{y}_t = \boldsymbol{\mu} + A_1\boldsymbol{y}_{t-1} + \dots + \mathbf{A}_k\boldsymbol{y}_{t-k} + \boldsymbol{u}_t \tag{4}$$

where $\boldsymbol{y}_t$ is a (3×1) vector of endogenous variables — in this analysis, for example, $\boldsymbol{y}_t=[Openness_t, Income_t, Emission_t]$; $\mathbf{A}_k$ is a (3×3) matrix of parameters; $\boldsymbol{\mu}$ is a vector of constant; and $\boldsymbol{u}_t$ is a vector of normally and independently distributed error terms, or white noise. Eq. (4) is in reduced form with each variable $\boldsymbol{y}_t$ regressed on only lagged variables of both itself and all the other variables in the system. Thus, ordinary least squares (OLS) will produce efficient estimates.

We emphasize that the possibility of unit roots in time-series data raises issues about parameter inference and spurious regression (Wooldridge, 2000). For example, OLS regression involving non-stationary series no longer provides the valid interpretations of the standard statistics such as t-statistics, F-statistics and confidence intervals. To avoid this problem, non-stationary variables should be differenced to make them stationary. However, Engle and Granger (1987) show that, even in the case that all the variables in a model are non-stationary, it is possible for a linear combination of integrated variables to be stationary. In this case, the variables are said to be cointegrated and the problem of spurious regression does not arise. Hence, the first requirement for cointegration analysis is that the selected variables must be non-stationary.

If all variables in $\boldsymbol{y}_t$ are non-stationary, a test for cointegration is identical to a test of long-run equilibrium. Following Johansen (1988), Eq. (4) can be reformulated into a vector error-correction (VEC) form to impose the cointegration constraint as follows:

$$\Delta\boldsymbol{y}_t = \mu + \boldsymbol{\Gamma}_1\Delta\boldsymbol{y}_{t-1} + \dots + \boldsymbol{\Gamma}_{k-1}\Delta\boldsymbol{y}_{t-k+1} + \boldsymbol{\Pi}\boldsymbol{y}_{t-k} + \boldsymbol{u}_t \tag{5}$$

where Δ is the difference operator; $\boldsymbol{\Gamma}_1,\dots\boldsymbol{\Gamma}_{k-1}$ are the coefficient matrices of short-term dynamics; and $\boldsymbol{\Pi}=-(I-\boldsymbol{\Pi}_1+\dots+\boldsymbol{\Pi}_k)$ are the matrix of long-run coefficients. If the coefficient matrix $\boldsymbol{\Pi}$ has reduced rank — i.e., there are $r\le(n-1)$ cointegration vectors present, then the $\boldsymbol{\Pi}$ can be decomposed into a matrix of loading vectors, $\boldsymbol{\alpha}$, and a matrix of cointegrating vectors, $\boldsymbol{\beta}$, such as $\boldsymbol{\Pi}=\boldsymbol{\alpha}\boldsymbol{\beta}'$. For three endogenous non-stationary variables in our analysis, for example, $\boldsymbol{\beta}'\boldsymbol{y}_{t-k}$ in Eq. (5) represents up to two linearly independent cointegrating relations in the system. The number of cointegration vectors, the rank of $\boldsymbol{\Pi}$, in the model is determined by the likelihood ratio test (Johansen, 1988).

When the number of cointegration vectors (r) has been determined, it is possible to test hypotheses under r by imposing linear restrictions on the matrix of cointegration vectors, $\boldsymbol{\beta}$, and loadings, $\boldsymbol{\alpha}$ (Johansen and Juselius, 1992). The tests for these linear restrictions are asymptotically χ^2 distributed. For example, testing for weak exogeneity is formulated by establishing all zeros in row i of $\boldsymbol{\alpha}_{ij}$, $j=1,\dots,r$, indicating that the cointegration vectors in $\boldsymbol{\beta}$ do not enter the equation determining $\Delta\boldsymbol{y}_{it}$. This means that, when estimating the parameters of the model ($\boldsymbol{\Gamma}_i$, $\boldsymbol{\Pi}$, $\boldsymbol{\alpha}$, $\boldsymbol{\beta}$), there is no loss of information from not modeling the determinants of $\Delta\boldsymbol{y}_{it}$; thus, this variable is weakly exogenous to the system and can enter on the right-hand side of the VAR model (Harris and Sollis, 2003).

3.2. Data

We compiled annual time-series data on sulfur emission (SO_2), income and trade openness for 50 countries for the period 1960–2000. The estimated sulfur emissions for 50 countries are obtained from a large database constructed by David Stern (Stern, 2005, 2006), which is known as the David Stern's Datasite (available at the web site http://www.rpi.edu/~sternd/datasite.html). To ensure comparability with per capita real GDP in the model, per capita SO_2 emissions for individual countries (measured in kg) are calculated using their population sizes. The per capita real GDP (measured in real PPP-adjusted dollars) is used as a proxy for income and is taken from the Penn World Table (PWT 6.2) (available at the web site http://pwt.econ.upenn.edu/php_site/pwt62/pwt62_form.php). The degree of openness of an economy (defined as the ratio of the value of total trade to real GDP) is used as a proxy for trade openness and is obtained from the Penn World Table.[4]

The data on sulfur emissions (SO_2) used in empirical studies have almost invariably come from a single source, the ASL and Associates database (ASL and Associate, 1997; Lefohn et al., 1999), which compiles annual time-series data on SO_2 for individual countries from 1850 to 1990. However, the unavailability of data after 1990 has limited continued use of these estimates for further research. Hence, David Stern has developed global and individual country estimates of sulfur emissions from 1991 to 2000 or 2002 (most OECD countries) combined with estimates from existing published and reported sources for 1850–1990 (see Stern, 2006 for more details). In addition, following the World Bank's country classification, 50 countries used in our analysis are divided into two groups on the basis of 2005 gross national income (GNI) per capita: (1) 25 developing economies, \$876–\$10,725; and (2) 25 developed economies, \$10,726 or more.[5]

[4] A well-known problem with this measure is that it does not take into account other factors (i.e., country size or trade policies) affecting trade flows; other things being equal, for example, large countries tend to have smaller trade shares. For this reason, some scholars have sought possible solutions to this problem. For example, Leamer (1988) and Pritchett (1991) have developed a theoretical model to correct the measure of openness for underlying trade policies such as tariffs and quotas. However, the time-series dataset on these improved approaches are not available. As such, the traditional measure of trade openness is used for our analysis.

[5] The World Bank's typical categories for classifying economies include low income (or least developed countries), middle income (or less developed/developing countries) and high income economies (developed countries). However, since the data set does not include any low income economies (\$875 or less based on 2005 GNI per capita), we simply divide 50 countries into developing and developed countries.

3.3. Econometric procedure

As noted earlier, the first requirement for the use of the Johansen cointegration method is that the variables must be non-stationary. The presence of a unit root in $\boldsymbol{y}_t$ ($Openness_t$, $Income_t$, $Emission_t$) for 50 countries is tested using the Dickey–Fuller generalized least squares (DF-GLS) test (Elliot et al., 1996). This test optimizes the power of the conventional augmented Dickey–Fuller (ADF) test by detrending. The DF-GLS test works well in small samples and has substantially improved power when an unknown mean or trend is present (Elliot et al., 1996). The results show that the levels of all the series (150 series) are non-stationary, while the first differences are stationary.[6] From these findings, we conclude that all the series are non-stationary and integrated of order 1, or $I(1)$; therefore, cointegration analysis can be pursued on them.

It should be noted that, before implementing the cointegration test, the important specification issue to be addressed is the determination of the lag length for the VAR model, because the Johansen procedure is quite sensitive to changes in lag structure (Maddala and Kim, 1998).[7] The lag length (k) of the VAR model is determined based on the likelihood ratio (LR) tests. This method compares the models of different lag lengths sequentially to see if there is a significant difference in results (Doornik and Hendry, 1994). Of the 50 countries, for example, the hypothesis that there is no significant difference between a two- and a three-lag model cannot be rejected for 19 countries. Thus, two lags ($k=2$) are used for those countries in our cointegration analysis. Diagnostic tests on the residuals of each equation and corresponding vector test statistics support the VAR model with two lags as a sufficient description of the data (Tables 1 and 2). In the residual serial correlation and heteroskedasticity tests, the null hypotheses of no serial correlation and no heteroskedasticity cannot be rejected at the 5% significance level. Although the null hypothesis of normality is rejected for some cases at the 5% significance level, non-normality of residuals does not bias the results of the cointegration estimation (Gonzalo, 1994). For the remaining 31 countries, on the other hand, both the VAR lag selection criterion and diagnostic tests consistently support $k=1$ as the most appropriate lag length for the VAR model.

4. Empirical results

With the selected lag lengths ($k=1$ or $k=2$) in non-stationary VAR models, the Johansen cointegration procedure is used to determine the number of cointegrating vectors among the variables. The results indicate that one cointegration vector is found for 25 countries at the 5% significance level, whereas no cointegration is found for 25 countries (Tables 1 and 2). More specifically, of the 25 developed countries, the trace tests show that the hypothesis of no cointegration ($r=0$) is rejected and that of one cointegration vector ($r=1$) is accepted at the 5% level for 17 countries (Table 1). For the remaining 8 countries, on the other hand, the trace statistics are well below the critical value and $r=0$ cannot be rejected at the 5% level, indicating that the three variables are not cointegrated.[8] The results thus, by and large, support the hypothesis that cointegration between SO_2 emissions, income and openness is pervasive across developed countries and a long-run equilibrium relationship among these variables also exists. In contrast, of the 25 developing countries, the trace tests show that only 7 countries have a cointegration rank of one ($r=1$), while the remaining 18 countries have $r=0$ (Table 2). This finding provides relatively weak support for the contention that the presence of long-run relationship among the three variables is pervasive across developing countries.

When determining the existence of cointegration relationship, the cointegration vectors ($\boldsymbol{\beta}_j$) estimated from Eq. (5) represent the long-run relationship among the selected variables. More specifically, having obtained only one cointegration relationship between SO_2 emissions, income and openness in the 24 countries that include developed and developing economies, the first eigenvector ($\boldsymbol{\beta}_1$) of the three eigenvectors is most highly correlated with the stationary part of the process $\Delta\boldsymbol{y}_t$ when corrected for the lagged values of the differences. Thus, $\boldsymbol{\beta}_1$ represents the cointegration vector determined by the CVAR model (Johansen, 1988). After normalizing the coefficient

Table 1
Results of Johansen cointegration tests and diagnostic tests for developed countries.

Continent	Country	Cointegration	Diagnostic tests		
			Serial correlation	Heteroskedasticity	Normality
Asia	Japan	Yes	1.47	0.70	15.91*
	Korea	Yes	1.65	1.03	7.65
	Israel	Yes	1.06	0.86	29.17**
	Singapore	Yes	1.25	1.37	6.56
North America	USA	Yes	0.88	0.90	7.60
	Canada	Yes	1.44	0.54	15.48*
Western Europe	Austria	No	1.32	0.94	7.66
	Belgium	No	1.17	0.47	11.49
	Denmark	Yes	0.71	0.74	12.62*
	Finland	Yes	1.25	1.02	12.59*
	France	Yes	1.56	1.13	6.64
	Greece	Yes	0.74	0.38	6.23
	Iceland	No	1.13	0.94	28.69**
	Ireland	Yes	1.75	0.82	14.21*
	Italy	Yes	0.72	1.09	8.07
	Luxembourg	No	1.41	0.55	8.66
	Netherlands	Yes	1.36	1.15	5.46
	Norway	No	1.29	0.60	11.63
	Portugal	Yes	0.91	1.02	19.37**
	Spain	Yes	0.97	0.72	3.65
	Sweden	Yes	1.18	1.11	12.16*
	Switzerland	No	1.08	0.72	18.90**
	UK	Yes	1.19	1.16	7.48
Oceania	Australia	No	1.24	0.71	7.13
	New Zealand	No	1.40	1.09	6.37

Note: The system specification tests based on F-tests indicate that the VAR model should include both an intercept and a time trend in most cases, with exceptions of the inclusion of an intercept only in United States, Israel and Luxembourg. Diagnostic tests are conducted using both the residuals of each equation (SO_2, Income and Openness) and corresponding vector test statistics; for brevity, however, the results of system diagnostic tests are only reported here. ** and * denote rejection of the null hypothesis (no serial correlation, homoskedasticity and non-normality) at the 1% and 5% levels, respectively.

[6] The results of unit roots are not reported here for brevity but can be obtained from the authors on request.

[7] The sample size could be another issue of concern for Johansen procedure, because finite-sample analyses can bias the cointegration test toward finding the long-run relationship either too often or too infrequently. Hakkio and Rush (1991) note: "Our Monte Carlo studies show that the power of a cointegration test depends more on the span of the data rather than on the number of observations. Furthermore, increasing the number of observations, particularly by using monthly or quarterly data, does not add any robustness to the results in tests of cointegration." Following these authors, therefore, the annual data used in this study (41 observations for the period 1960–2000) are considered to be long enough to reflect the long-run relationship between trade, income and SO_2 emissions.

[8] 23 developed countries (except Australia and New Zealand) show that their SO_2 emissions levels tend to grow first, reach a peak and then start declining after a turning point. For completeness, therefore, we also examine if a structural break occurs in the long-run relationship among the three variables in these 23 developed countries. For this purpose, we use the most recent Johansen cointegration technique that allows for structural breaks at known points in time (Johansen et al., 2000). For the United States, Japan and France, for example, the plausible breaks seem to occur at 1969, 1970 and 1973, respectively, for the SO_2 emissions series (Fig. 2). We find the same results as those obtained from the standard Johansen method. The similarity of the long-run relationships for both cases suggest that the plausible breaks in the SO_2 emissions series have not altered the long-run equilibrium relationship among the three variables in these 23 developed countries.

of SO_2 emissions, for example, the long-run equilibrium relation ($\boldsymbol{\beta}_1$) between the three variables in the United States can be represented as the following reduced form; $Emission_t = -0.98\ Income_t - 0.11\ Openness_t$ (Table 3). In this equation, a negative coefficient of income on sulfur emissions suggests that environmental quality improves as the U.S. income increases. Similarly, a negative coefficient of openness on SO_2 implies that trade liberalization tends to reduce SO_2 emissions in the United States. We emphasize here that, in this study, we do not interpret the coefficients of the long-run relationship as long-run elasticities because such an interpretation may ignore the dynamics of the system (Lütkepohl, 2005). A 1% increase in the U.S. real income, for example, may not cause a long-term decline in SO_2 emissions by 0.98% because an increase in the U.S. income is likely to have an effect on trade openness as well that may interact in the long-run.

4.1. Analyzing long-run relationship

As noted earlier, the cointegration vector, $\boldsymbol{\beta}_1$, estimated from Eq. (5) is used to describe the long-run relationship between SO_2 emissions, income and openness after normalizing the coefficients of SO_2 emissions and rearranging in reduced forms (Table 3). The results show that, of the 17 developed countries in which all three variables are cointegrated, 13 countries show a negative long-run relationship between SO_2 emissions and per capita income, suggesting that pollution levels tend to decrease as a country's economy grows. For the remaining 4 countries (Israel, Singapore, Greece and Portugal), on the other hand, SO_2 emissions have a positive long-run relationship with per capita income, indicating economic growth tends to worsen environmental quality. This phenomenon could be explained using what is known in the literature as emissions intensity. More specifically, emissions intensity is generally defined as the ratio of a measure of environmental quality (e.g., sulfur dioxide and carbon dioxide emissions) to a measure of economic output. In this study, emissions intensity thus can be expressed as the ratio of per capita SO_2 emissions to per capita income. As an economy generally moves beyond the EKC threshold level of income, SO_2 emissions tend to decrease with higher income per capita, thereby improving emission intensity (a decrease in the ratio) (see the second-quadrant of Fig. 1). Under this circumstance, therefore, SO_2 emissions have a negative relationship with income. When an economy has not reached the EKC turning points, on the other hand, SO_2 emissions tend to increase with higher income per capita, thereby deteriorating emissions intensity (an increase in the ratio).[9] This can be interpreted to mean that SO_2 emissions thus have a positive relationship with income. In fact, the 13 economies that have a negative emission–income relationship are shown to have crossed a wide range of the EKC turning points from approximately $11,000 to $19,000 per capita income (in 2000 U.S. dollars) between 1969 and 1975 (Fig. 2). Accordingly, the emissions intensities of those countries have significantly improved over the last 50 years (Fig. 3). It is shown, on the other hand, that the 4 economies indicating a positive emission–income relationship have not reached income levels high enough to be able to derive the EKC turning point so that emission level tends to increase with higher income growth (Fig. 4).[10] Accordingly, the emissions intensities of those countries have improved little (Israel and Singapore) or even have deteriorated (Greece and Portugal) over the last 50 years (Fig. 2). Hence, these

Table 2
Results of Johansen cointegration tests and diagnostic tests for developing countries.

Continent	Country	Cointegration	Diagnostic tests		
			Serial correlation	Heteroskedasticity	Normality
Asia	China	Yes	0.97	0.61	8.56
	India	No	1.19	0.71	3.59
	Indonesia	No	1.49	1.29	20.53**
	Jordan	No	1.34	0.98	6.54
	Philippines	No	0.85	1.02	12.46*
	Sri Lanka	Yes	1.09	0.68	25.17**
	Thailand	No	1.75	1.27	22.91**
	Turkey	Yes	1.84	1.38	1.05
Central America	Costa Rica	No	1.03	0.90	36.82**
	El Salvador	No	1.19	1.28	29.09**
	Guatemala	Yes	0.41	0.84	64.67**
	Honduras	No	1.52	0.80	21.06**
	Mexico	Yes	0.59	1.26	3.95
	Nicaragua	No	0.82	0.82	24.87**
	Panama	No	0.70	0.84	35.80**
South America	Argentina	No	1.39	1.15	14.09*
	Bolivia	No	1.47	0.45	28.25**
	Brazil	No	1.24	0.69	6.19
	Chile	No	0.70	1.33	26.96**
	Columbia	No	1.55	0.68	8.81
	Ecuador	No	1.35	0.87	12.84*
	Paraguay	No	1.21	1.26	21.78**
	Peru	Yes	0.77	0.61	11.49
	Uruguay	Yes	0.96	0.99	21.64**
	Venezuela	No	1.67	1.14	14.43*

Note: The system specification tests based on *F*-tests indicate that the VAR model should include both an intercept and a time trend in most cases, with exceptions of the inclusion of an intercept only in India and Thailand. Diagnostic tests are conducted using both the residuals of each equation (SO_2, Income and Openness) and corresponding vector test statistics; for brevity, however, the results of system diagnostic tests are only reported here. ** and * denote rejection of the null hypothesis (no serial correlation, homoskedasticity and non-normality) at the 1% and 5% levels, respectively.

Table 3
Results of long-run relationship between SO_2 emissions, income and openness for developing and developed countries.

Continent	Country	Income	Openness
Developed countries			
Asia	Japan	−2.69	−0.95
	Korea	−2.54	−3.86
	Israel	+3.78	−0.63
	Singapore	+4.45	−0.45
North America	USA	−0.98	−0.11
	Canada	−2.23	−1.08
Western Europe	Denmark	−3.45	−6.07
	Finland	−1.38	−3.85
	France	−3.81	−1.38
	Greece	+1.02	−5.18
	Ireland	−1.65	+1.43
	Italy	−4.76	−3.08
	Netherlands	−0.87	−0.27
	Portugal	+1.01	−0.82
	Spain	−1.11	−1.03
	Sweden	−1.52	−0.90
	UK	−2.94	−1.04
Developing countries			
Asia	China	−1.13	−1.22
	Sri Lanka	+9.46	+0.51
	Turkey	+6.46	+4.97
Central America	Guatemala	+4.38	+1.72
	Mexico	+3.11	+2.74
South America	Peru	+0.68	+1.51
	Uruguay	+0.55	+4.87

Note: Since the long-run equilibrium relation ($\boldsymbol{\beta}_1$) is normalized to SO_2 emissions, coefficients indicate the negative and positive relationships between SO_2 emissions and the two variables; for example, in the case of Japan, SO_2 emissions have negative relationships with both income and openness.

[9] This can be true only if the growth rate of per capita SO_2 emissions is faster than that of per capita real income. In other words, although an economy has not crossed the EKC threshold, emission intensity may improve as income grows faster than emission level.

[10] Interestingly, Singapore and Israel have so-called N-shape curves, which exhibit the inverted-U curve initially, but beyond a certain income level, the relationship between SO_2 emissions and income turns positive again (Fig. 4). Israel (Singapore), for example, has a secondary turning point between income levels of $12,332 ($13,163) and $14,422 ($15,883).

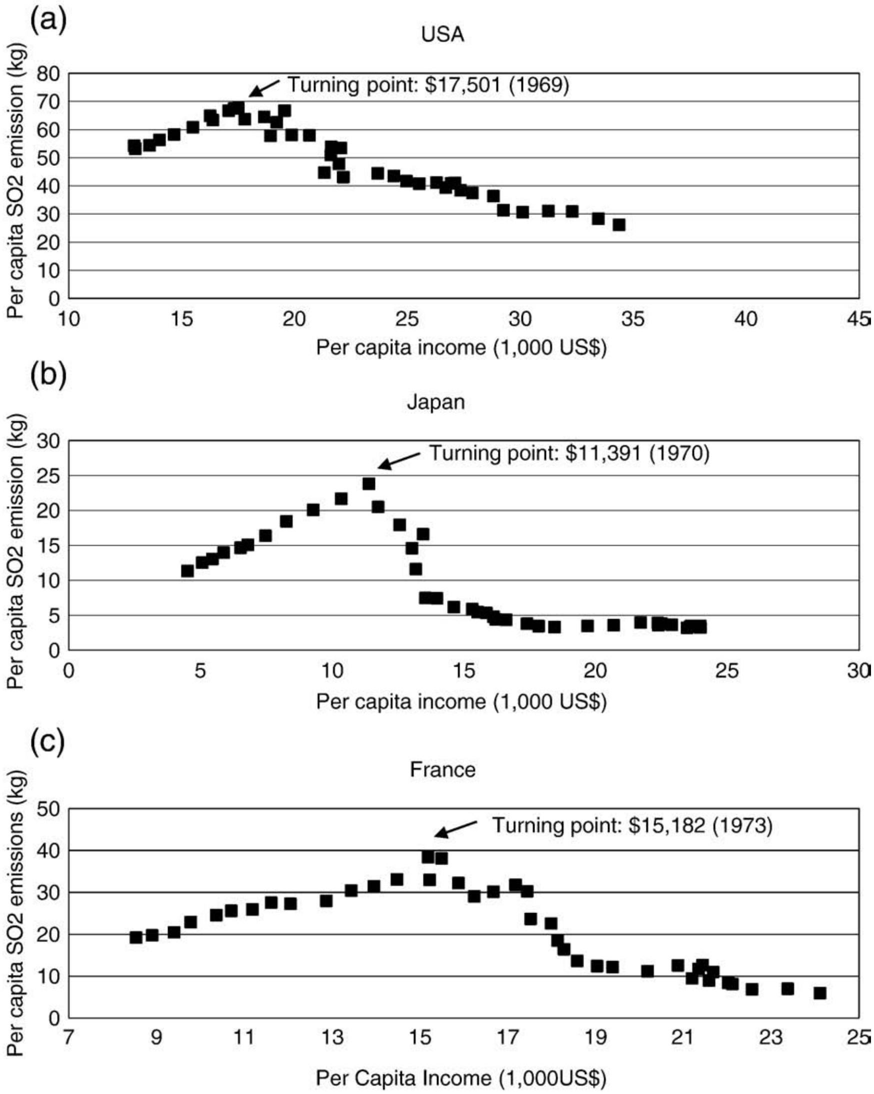

Fig. 2. Plots of per capita SO_2 emissions and per capita income for US, Japan and France.

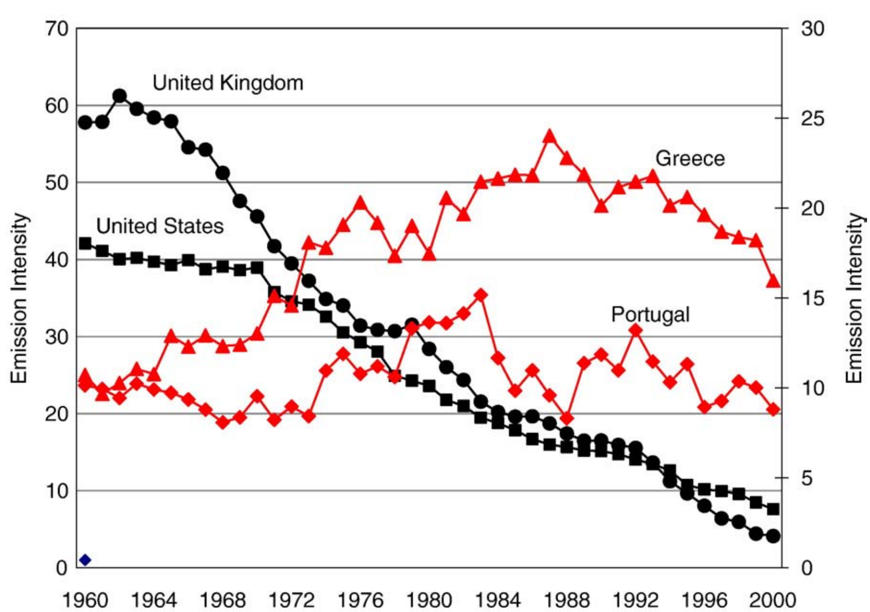

Fig. 3. Emissions intensities for developed countries.

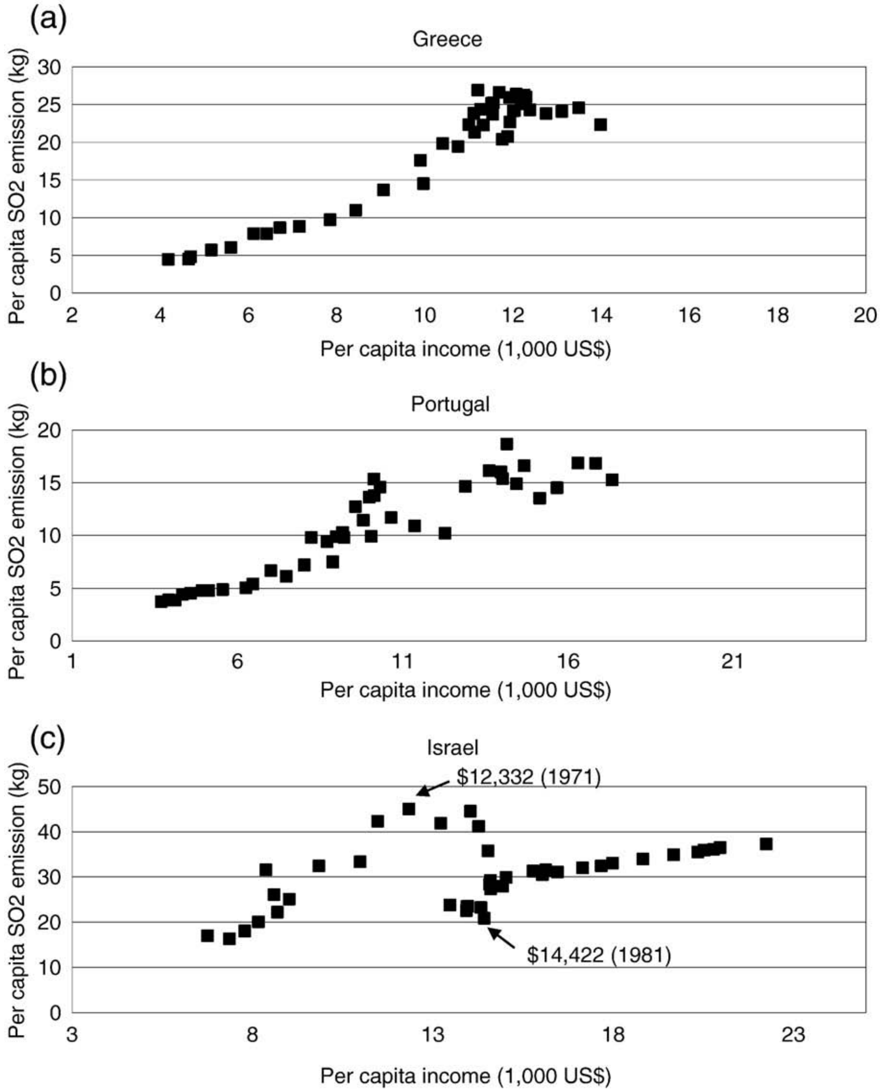

Fig. 4. Plots of per capita SO_2 emissions and per capita income for Greece, Portugal and Israel.

findings provide empirical evidence for the existence of the Environmental Kuznets Curve in the sense that as income of an economy grows, the emission level starts declining (rising) after (before) a threshold level of income has been crossed.

In addition, of the 17 developed countries in which all three variables are cointegrated, 16 countries show a negative long-run relationship between SO_2 emissions and openness, indicating that air pollution tends to decrease as a country's exposure to international markets increases. The finding supports the so-called *gains-from-trade hypothesis* for developed countries; output growth induced by trade liberalization increases incomes, and wealthier countries tend to be more willing and able to channel resources into environmental protection through the enforcement of environmental standards and the investment on cleaner production technologies, thereby improving environmental quality.

Of the 7 developing countries in which all three variables are cointegrated, on the other hand, 6 countries (Peru, Uruguay, Guatemala, Mexico, Sri Lanka and Turkey) show a positive long-run relationship between SO_2 emissions and income, indicating that economic growth worsens environmental quality (Table 3). For example, Peru and Mexico show that their emission levels tend to increase with higher income growth (Fig. 5); accordingly, these two countries show little improvement in the emissions intensities over the last 50 years (Fig. 6).[11] In addition, in these 6 countries, SO_2 emissions have a positive long-run relationship with openness, supporting the so-called *pollution haven hypothesis* for developing countries. More specifically, confronted with international competition, poor open economies have incentives to adopt excessively lax environmental standards in an effort to attract multinational corporations, particularly those engaged in highly polluting activities. As developed countries enforce strict environmental regulations, therefore, trade liberalization leads to more rapid growth of dirty industries in developing countries that export more pollution-intensive goods, thereby deteriorating environmental quality.

Notice that, in the case of China, SO_2 emissions have a negative long-run relationship with income and openness, suggesting that growth and trade liberalization seem to improve environmental quality. The finding may be peculiar in light of the fact that China has been the largest sulfur dioxide polluter over the last decade. However, emissions intensity can keep improving as the growth rate of per capita real income is faster than that of per capita SO_2 emissions.

[11] The remaining four countries (Uruguay, Guatemala, Sri Lanka and Turkey) also show the same patterns. To save space, however, the figures for those four countries are not shown here but will be available upon request.

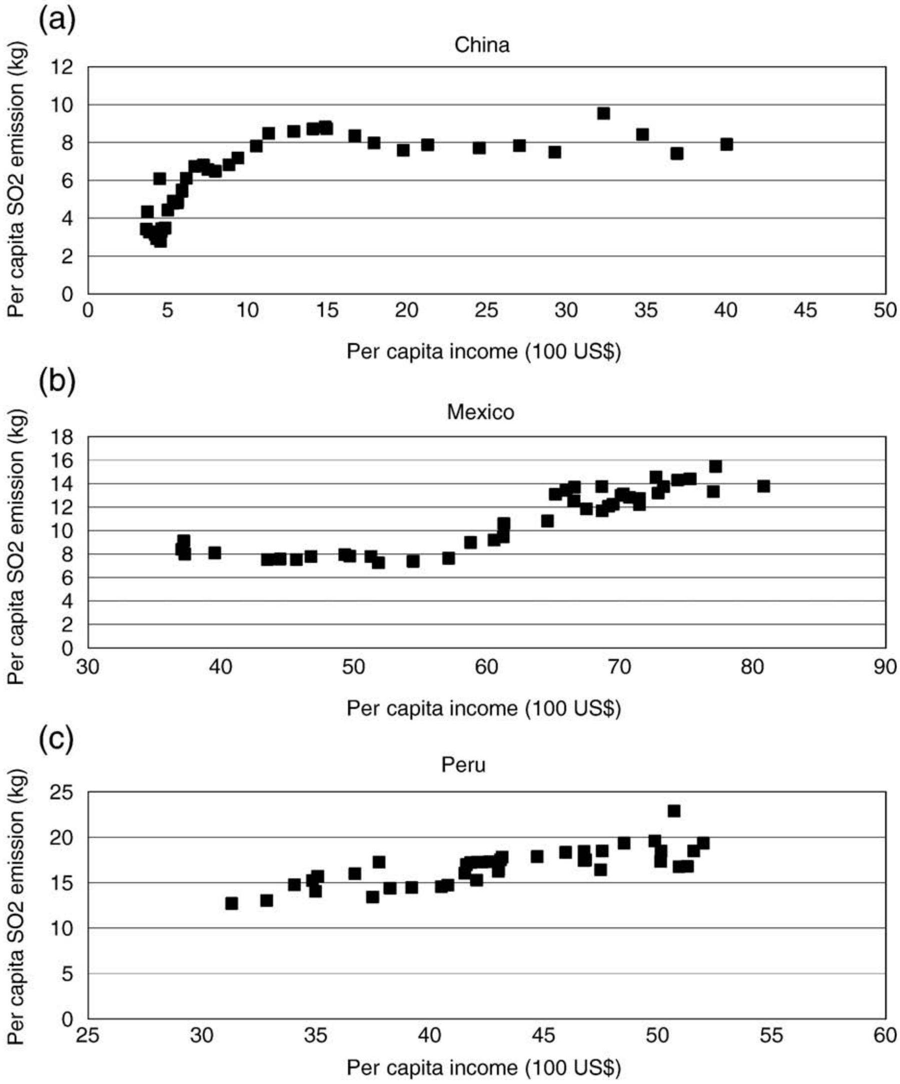

Fig. 5. Plots of per capita SO_2 emissions and per capita income for China, Mexico and Peru.

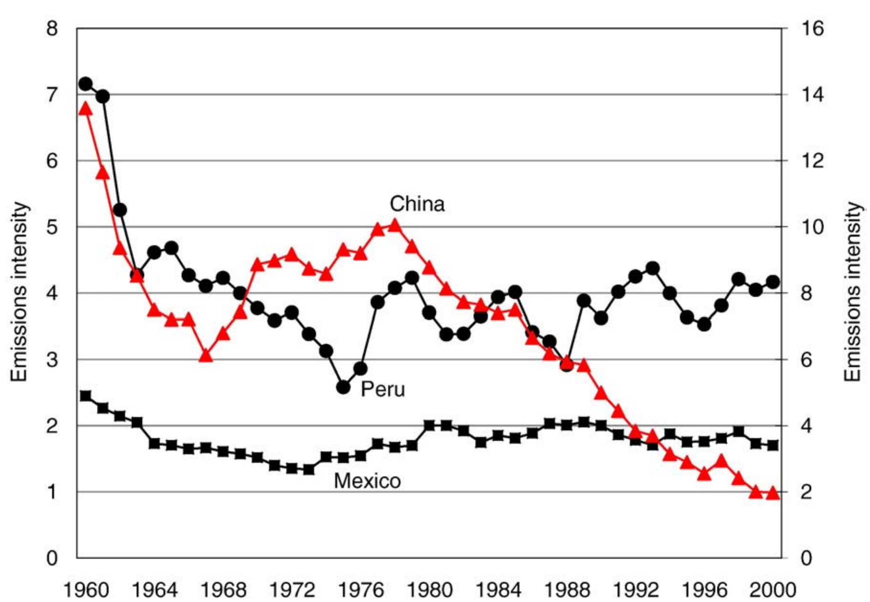

Fig. 6. Emissions intensities for developing countries.

Indeed, China has witnessed much faster growth in real income than SO_2 emissions since the late 1970s when the reforms began and opening-up started. For example, SO_2 emissions of China have increased by an average of 2.7% annually since 1978, while real income has grown by an average of 8.5% annually over this period. As a result, the Chinese economy seems to have shown substantial improvement in the emissions intensity for 20 years despite increasing SO_2 emission level (Fig. 6).

4.2. Identifying the causal effects

In order to identify the causal effects of trade and income on the environment, the long-run weak exogeneity test is conducted by restricting the speed-of-adjustment parameter ($\boldsymbol{\alpha}$) to zero in the model. This test examines the absence of long-run levels of feedback due to exogeneity (Johansen and Juselius, 1992). In other words, a weakly exogenous variable is a driving variable, which pushes the other variables adjusting to long-run equilibrium, but is not influenced by the other variables in the model. The results show that, of the 17 developed countries in which all three variables are cointegrated, the null hypothesis of weak exogeneity cannot be rejected for openness and/or income at the 5% level for 14 countries (Table 4), indicating that these two variables are weakly exogenous to the long-run relationships in the model. For the remaining 3 countries, on the other hand, the null hypothesis cannot be rejected for SO_2 emissions. These findings indicate that, for developed countries, openness and/or income are generally the driving variables in the system and significantly affect SO_2 emissions in the long-run, but are not influenced by SO_2 emissions. This implies that, since environmental quality is a normal good, economic growth induced by trade liberalization allows for the possibility that people in developed countries demand tougher environmental standards and cleaner production technologies, which can contribute to the improvement of environmental quality. This further suggests that, since economic growth in the industrialized countries tends to cause a shift in the structure of the economy from the more energy-intensive manufacturing sector toward the more environmentally-friendly service sector, those countries may be able to fulfill their aspirations for income growth and/or freer trade without environmental degradation.

Of the 7 developing countries in which all three variables are cointegrated, on the other hand, the null hypothesis of weak exogeneity cannot be rejected for SO_2 emissions at the 5% level for 6 countries — Sri Lanka, Turkey, Guatemala, Mexico, Peru and Uruguay. For China, however, the null hypothesis cannot be rejected at the 5% level for openness and income. These results indicate that, for developing countries (except for China), the SO_2 emissions are generally weakly exogenous to the long-run parameters in the system; thus, the emission does not adjust to deviations from any equilibrium state defined by the cointegration relation. This suggests that with relatively low environmental standards, developing countries tend to attract pollution-intensive industries as developed economies enforce strict environmental regulations, and developing countries are likely to be net exporters of pollution-intensive goods, which in turn deteriorates environmental quality as openness proceeds. This further implies that, if developing countries attempt to introduce tougher environmental standards, there will be slowdowns in the income growth rate and/or trade openness. As such, the developing countries may have to sacrifice some current income growth through trade openness if they decide to reduce permanently the emission level from what it is at present.

Table 4
Results of weak exogeneity tests for developing and developed countries.

Continent	Country	Weak exogeneity (H_0:$\alpha_i = 0$)		
		SO_2 emissions	Income	Openness
Developed countries				
Asia	Japan	14.52**	0.05	3.93*
	Korea	8.31**	0.64	7.53**
	Israel	0.91	16.29**	0.69
	Singapore	13.01**	4.50*	2.14
North America	USA	12.65**	1.35	15.88**
	Canada	6.95**	0.73	8.19**
Western Europe	Denmark	20.36**	2.95	16.02**
	Finland	7.75**	1.35	6.31*
	France	18.79**	14.35**	3.62
	Greece	11.22**	1.32	4.82*
	Ireland	1.02	15.50**	15.97**
	Italy	34.27**	5.19*	1.92
	Netherlands	14.83**	7.21**	0.54
	Portugal	12.52**	8.78**	0.01
	Spain	1.48	0.77	11.34**
	Sweden	3.86*	9.91**	2.64
	UK	14.06**	0.21	12.55**
Developing countries				
Asia	China	10.60**	0.91	1.48
	Sri Lanka	3.06	4.45*	0.09
	Turkey	2.85	2.45	20.83**
Central America	Guatemala	3.23	3.86*	0.61
	Mexico	0.01	0.19	42.45**
South America	Peru	0.59	5.00*	9.27**
	Uruguay	2.34	10.89**	7.34**

Note: ** and * denote rejection of the null hypothesis of weak exogeneity at the 1% and 5% levels, respectively. Values are the likelihood ratio (LR) test statistic based on the χ^2 distribution.

5. Conclusions

The environmental consequence of trade liberalization has been vigorously debated during the last decade. In this paper, we explore the dynamic effect of trade liberalization on the environment for both developed and developing countries over the last five decades. The primary contribution of this paper is to directly address the issue of potential endogeneity problems and the causal mechanism of trade, income and environmental quality (measured by SO_2 emissions) in the framework of individual country-specific data and the cointegrated VAR model. We use Johansen's maximum likelihood procedure to estimate the coefficients of the cointegrated VAR.

The empirical results generally indicate a negative long-run relationship between SO_2 emissions and income for developed countries and a positive long-run relationship between them for developing countries; that is, an increase in the level of income results in an improvement (deterioration) of environmental quality for developed countries (developing countries), supporting the existence of the Environmental Kuznets Curve. We also find that, while trade liberalization appears to increase environmental quality in developed economies, it has a detrimental effect on environmental quality in most developing countries, supporting *pollution haven hypothesis*. The results further show that for developed countries, the causality seems to run from trade and/or income to SO_2 emissions, indicating that a change in the level of trade openness and income causes a consequence change in environmental quality. For most developing countries, on the other hand, the causality is found to run from SO_2 emissions to trade and/or income, suggesting that a change in environmental quality causes a consequence change in income and openness; however, the opposite causal relationship is found to hold for China.

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