The Environmental Consequences of Globalization:

A Country-Specific Time-Series Analysis

Jungho Baek
Research Assistant Professor
Department of Agribusiness and Applied Economics
North Dakota Sate University
Email: jungho.baek@ndsu.edu
Phone: (701) 231-7451
Fax: (701) 231-7400

Yong S. Cho
Professor
Department of Food and Resource Economics
Korea University
Email: yscho@korea.ac.kr
Phone: (822) 3290-3037

Won W. Koo
Professor
Department of Agribusiness and Applied Economics
North Dakota Sate University
Email: won.koo@ndsu.edu
Phone: (701) 231-7448
Fax: (701) 231-7400

Selected Paper prepared for presentation at the American Agricultural Economics Association Annual Meeting, Orlando, FL, July 27-29, 2008

The Environmental Consequences of Globalization: A Country-Specific Time-Series Analysis

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, and the opposite relationship holds for developing countries.

Keywords*:* Developed countries, Developing countries, Environmental quality, Globalization, Time-series analysis, Trade

INTRODUCTION

One of the most important debates in international trade policy over the last decade has been the environmental consequences of trade liberalization/globalization (Copeland and Taylor 1994 and 2004). Proponents of trade liberalization argue that, since 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. 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 methods do not change, then environmental quality deteriorates as trade increases the scale of economic activity. Moreover, developing economies tend to adopt looser standards of environmental regulations to attract more foreign investment. Trade liberalization thus may lead more growth of pollution-intensive industries in developing countries as developed countries enforce strict environmental regulations. As a result, free trade has 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.[1] For example, Lucas et al. (1992) investigate the influence of trade openness on the growth rate of toxic intensity of output. They find that a high degree of restrictive trade policies tends to increase pollution intensity in fast-growing economies. Gale and Mendez (1998) analyze the relationship between trade, growth and the environment, and find that an increase in income has a detrimental effect on environmental quality, but trade effect on pollution is

[1] 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.

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, and conclude that there is little evidence that 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 reduced-form models to examine the presence of significant statistical association of trade openness and income growth with environmental quality. Little attention has been paid to the causal effects of trade liberalization and income on the environment (Coondoo and Dinda 2002, Chintrakarn and Millimet 2006). More specifically, with the treatment of trade and income as being exogenous variables for their reduced-form models, past studies run the regression of measures of environmental quality/damage (e.g., sulfur dioxide and carbon dioxide emissions) on 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. Hence, this presumption neglects the possibility of endogeneity of trade and income in the model. In other words, since 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 increase 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' experience (e.g., 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 (e.g., 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 allows determining both the short- and long-run dynamic effects of selected variables and testing the endogeneity of them. The results of this procedure could thus be interpreted as revelation of potential impacts of shocks in an exogenous variable on every endogenous variable. Compared to reduced-form equations, 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. By far, however, no studies have attempted to directly address the potential endogeneity of income, trade and the environment with individual country-specific data and time-series models.[2]

In this paper, therefore, we use the Johansen cointegration analysis to examine the dynamic effect of trade liberalization on the environment using 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 and 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 between 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 draw some conclusions.

A THEORETICAL FRAMEWORK

[2] 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 (e.g., Coondoo and Dinda 2002 and 2006, Perman and Stern 2003) 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.

Following Copeland (2005), a simple model involving demand and supply of emissions is presented in Figure 1 to examine the effects of trade liberalization on income and the environment. In this model, a country is assumed to export pollution-intensive goods (dirty goods) and a pollution tax (τ) is used as a proxy for the stringency of environmental policy. The demand for emissions (D) is a derived demand, reflecting emission of pollution as a side effect of production; a country produces more pollution as a pollution tax (costs of environmental damage) is low. The supply of emissions (S) represents the country's willingness to allow emissions as reflected by the pollution policy.

Consider a country that has a fixed pollution tax. The supply curve is then S_0. Initially equilibrium values for pollution tax (τ_0) and pollution level (Z_0) are determined by the intersection between the demand (D_0) and supply curves (S_0). In this case, trade liberalization leads to an increase in exports of pollution-intensive goods and results in a shift in the demand for emissions to D_1, thereby increasing in emissions to Z_1. On the other hand, consider a situation where the government tightens up environmental policy as pollution increases. The supply curve is then represented by S_1. In this case, a trade-induced outward shift of demand for emissions leads to pollution at Z_2. As a result, the endogenous policy response dampens the increase in pollution from Z_1 to Z_2.

It should be emphasized that since the motivation for trade liberalization is usually to increase real income in a country, income effects play a key role in the analyses of trade effects on the environment (Copeland 2005). Figure 1 also can be used to illustrate how income effects influence the predicted effects of trade and the environment. Assume that the supply curve of emissions is income-responsive. In this case, since environmental quality is a normal good, trade-induced income growth causes people to increase their demand for a clean environment and results in a shift in the supply of emissions to S_2 if the government responses to people's preference, which leads to a decrease in pollution (Z_3) from trade liberalization despite the country having a comparative advantage in the dirty goods. The magnitudes by which the supply curve shifts back depend on income and substitution effects.

EMPRICIAL METHODOLOGY

This study examines the dynamic relationship between income, trade liberalization and environmental quality for each of 50 developing and developed countries. There are two emission variables that have been widely used in the literature: sulfur dioxide (SO_2) and carbon dioxide (CO_2). Of these, SO_2 represents the measure of local air pollution, whereas CO_2 represents a global pollutant (externality), which individual countries are unable to regulate without international cooperation (Dinda 2004, Frankel and Rose 2005). It is thus more appropriate to use SO_2 as a proxy for the measure of environmental quality in our individual country-specific analysis.

Development of Empirical Time-Series Models

To examine dynamic interrelationship between trade, income and environmental quality (SO_2), the cointegrated vector autoregression (CVAR) model developed by Johansen is applied (Johansen 1995). The Johansen method uses a statistical model involving up to k lags as follows:

(1) $y_t = \mu + A_1 y_{t-1} + ... + A_k y_{t-k} + u_t$

where y_t is a (3×1) vector of endogenous variables ─ in this analysis, for example, $y_t =$ $[Openness_t, Income_t, Emission_t]$; A_k is an (3×3) matrix of parameters; μ is a vector of constant; and u_t is a vector of normally and independently distributed error terms, or white noise. Equation (1) is in reduced form with each variable 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.

It should be emphasized 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 and F-statistics. To avoid this problem, non-stationary variable should be differentiated 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 y_t are non-stationary, a test for cointegration is identical to a test of long-run equilibrium. Following Johansen (1995), equation (1) can be reformulated into a vector error-correction (VEC) form to impose the cointegration constraint as follows:

(2) $\Delta y_t = \mu + \Gamma_1 \Delta y_{t-1} + ... + \Gamma_{k-1} \Delta y_{t-k+1} + \Pi y_{t-k} + u_t$

where Δ is the difference operator; $\Gamma_1,\ldots,\Gamma_{k-1}$ are the coefficient matrices of short-term dynamics; and $\Pi = -(I - \Pi_1 + \ldots + \Pi_k)$ are the matrix of long-run coefficients. If the coefficient matrix Π has reduced rank — i.e., there are $r \leq (n-1)$ cointegration vectors present, then the Π can be decomposed into a matrix of loading vectors, α, and a matrix of cointegrating vectors, β, such as $\Pi = \alpha\beta'$. For three endogenous non-stationary variables in our analysis, for example, $\beta' y_{t-k}$ in equation (2) represents up to two linearly independent cointegrating relations in the system. The number of cointegration vectors, the rank of Π, in the model is determined by the likelihood ratio test (Johansen 1995).

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, β, and loadings, α (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 α_{ij}, $j = 1,\ldots,r$, indicating that the cointegration vectors in β do not enter the equation determining Δy_{it}. This means that, when estimating the parameters of the model (Γ_i, Π, α, β), there is no loss of information from not modelling the determinants of Δy_{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).

Data

We have 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 and 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 GDP in the model, per capita SO_2 emissions for individual countries (measured in kg) are calculated using their population sizes. The per capita 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 GDP) is used as a proxy for trade openness and is obtained from the Penn World Table.

It should be pointed out that 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 been an impediment to 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 (2005) 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 per capita: (1) 25 developing economies, \$876- \$10,725; and (2) 25 developed economies, \$10,726 or more.

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 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. 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). 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. 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.[3]

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 24 countries at the 5% significance level, whereas no cointegration is found for 26 countries (Tables 1-2). 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.[4] 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. The results thus, by and large, support for the hypothesis that cointegration between SO_2 emissions, income and openness is pervasive across developed countries. 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 $r=0$. This

[3] The results of unit roots and diagnostic tests are not reported here for brevity.
[4] Interested readers can contact the authors for more details of cointegration test results.
.

finding indicates that the three variables have no inherent co-movement tendency over the long-run across developing countries.

When determining the existence of cointegration relationship, the cointegration vectors (β_j) estimated from equation (2) 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 (β_1) of the three eigenvectors is most highly correlated with the stationary part of the process Δy_t when corrected for the lagged values of the differences. Thus, β_1 represents the cointegration vector determined by the CVAR model (Johansen 1995). After normalizing the coefficient of SO_2 emissions, for example, the long-run equilibrium relation (β_1) between the three variables in the United States can be represented as the following reduced form; $Emmision_t = -0.98 Income_t - 0.11 Openness_t$. In this equation, a negative coefficient of income on sulfur emissions suggests that environmental quality improves as the U.S. income increases. A negative coefficient of openness on SO_2, on the other hand, implies that trade liberalization tends to reduce SO_2 emissions in the United States. Note 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). For example, a 1% increase in the U.S. real income 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.

Analyzing Long-Run Relationship

As noted earlier, the cointegration vector, β_1, estimated from equation (2) 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 1). The results show that, of the 17 developed countries in which all three variables are cointegrated, 12 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 5 countries (Israel, Singapore, Greece, Portugal and Spain), 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 directly associated with changes in emissions intensity. More specifically, emissions intensity is defined as the ratio of sulfur dioxide emissions to a measure of economic output (per capita income). Deterioration (improvement) of emissions intensity implies that SO_2 emissions tend to increase (decrease) as income grows, which in turn indicates a positive (negative) relationship between SO_2 emissions and income.[5] In fact, the emissions intensities of the 12 economies that show a negative emission-income relationship have significantly improved over the last 50 years. In contrast, the emissions intensities of the 5 economies that show a positive emission-income relationship have improved little (Israel, Singapore and Spain) or even have deteriorated (Greece and Portugal) over the last 50 years (Figure 2). In addition, of the 17 developed countries in which the three variables are cointegrated, 15 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.

[5] It should be noted that SO_2 emissions can keep increasing unless emissions intensity improves faster than the economy grows. In this case, SO_2 emissions could have a positive relationship with income despite improvement of emissions intensity.

These results support for the so-called *gains-from-trade* hypothesis for developed countries; a rise in income growth through trade gradually tends to increase cleaner techniques of production, thereby improving environmental quality.

On the other hand, of the 7 developing countries in which all three variables are cointegrated, 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 2). In addition, in these 6 countries, SO_2 emissions have a positive long-run relationship with openness, supporting for the so-called *race-to-the bottom* hypothesis for developing countries; confronted with international competition, poor open economies have incentives to adopt excessively lax environmental standards in an effort to attract multinational corporations and export pollution-intensive goods. As a result, trade is the cause of environmental degradation in developing countries. For China, on the other hand, SO_2 emissions have a negative long-run relationship with income and openness, suggesting that growth and trade liberalization improve environmental quality. In fact, unlike other developing countries, the emissions intensity of the Chinese economy has substantially improved since 1978 (Figure 3). From these findings, therefore, it seems reasonable for us to conclude that, among developing countries, only China has led to both continued economic growths through more open trade and a cleaner environment.

Identifying the Causal Effects

In order to identify the casual effects of trade and income on the environment, the long-run weak exogeneity test is conducted by restricting parameter in speed-of-adjustment

(α) 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, the null hypothesis of weak exogeneity cannot be rejected for openness and/or income at the 5% level for 14 countries (Table 3), 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 trade liberalization and income growth may cause people in developed countries to increase their demand for a cleaner environment, thereby enforcing strict environmental regulations. This further suggests that the developed countries tend to restrain their aspirations for income growth and/or freer trade in order to control environmental degradation.

Of the 7 developing countries, on the other hand, the null hypothesis of weak exogeneity cannot be rejected for SO_2 emissions at the 5% level for 5 countries. For the remaining two countries (Peru and China), on the other hand, the null hypothesis cannot be rejected at the 5% level for openness and income, respectively. These results indicate that, for developing countries, 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 trade liberalization tends to create an incentive for pollution-intensive industries (so called dirty industries) to relocate in developing countries with lower environmental standards as developed countries adopt tighter environmental protection, thereby deteriorating environmental quality. This further implies that that, if developing countries attempt to control the emission rate, there will be a corresponding reduction in the income growth rate and/or trade volume. As such, the developing countries may have to accept a reduction of their current income levels and/or degree of trade openness if they have to reduce permanently the emission level from what it is at present.

CONCLUSIONS

In this paper, we examine the long-run effect of trade liberalization on the environment for both developing and developed countries over the last half-century. For this purpose, the effects of the trade openness and per capita income on per capita SO_2 emissions are investigated using the Johansen multivariate cointegration analysis. It is generally found 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. On the other hand, we 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. We also find that for developed countries the causality seems to run from trade and/or income to SO_2 emissions. For developing countries, on the other hand, the causality is found to run in the opposite direction from SO_2 emissions to trade and/or income. These results imply that for developed economies SO_2 emissions are the adjusting parts, while trade/income are the determining parts of the long-run relationship, and the opposite relationship holds for developing countries.

Table 1. Results of Johansen cointegration tests and long-run relationship between SO_2 emissions, income and openness for developed countries

	Country	*Cointegration*	*Income*	*Openness*
Asia	Japan	**Yes**	−	−
	Korea	**Yes**	−	−
	Israel	**Yes**	+	−
	Singapore	**Yes**	+	−
North America	USA	**Yes**	−	−
	Canada	**Yes**	−	−
Western Europe	Austria	No		
	Belgium	No		
	Denmark	**Yes**	−	−
	Finland	**Yes**	−	−
	France	**Yes**	−	−
	Greece	**Yes**	+	−
	Iceland	**Yes**	−	−
	Ireland	**Yes**	−	+
	Italy	**Yes**	−	−
	Luxembourg	No		
	Netherlands	**Yes**	−	−
	Norway	No		
	Portugal	**Yes**	+	−
	Spain	**Yes**	+	+
	Sweden	**Yes**	−	−
	Switzerland	No		
	UK	**Yes**	−	−
Oceania	Australia	No		
	New Zealand	No		

Note: − and + denote negative and positive signs, respectively. The long-run equilibrium relation (β_1) is normalized to SO_2 emissions; for example, a negative(positive) sign for income (openness) presents a negative (positive) relationship between SO_2 emissions and income (openness).

Table 2. Results of Johansen cointegration tests and long-run relationship between SO_2 emissions, income and openness for developing countries

	Country	*Cointegration*	*Income*	*Openness*
Asia	China	**Yes**	−	−
	India	No		
	Indonesia	No		
	Jordan	No		
	Philippines	No		
	Sri Lanka	**Yes**	+	+
	Thailand	No		
	Turkey	**Yes**	+	+
Central America	Costa Rica	No		
	El Salvador	No		
	Guatemala	**Yes**	+	+
	Honduras	No		
	Mexico	**Yes**	+	+
	Nicaragua	No		
	Panama	No		
South America	Argentina	No		
	Bolivia	No		
	Brazil	No		
	Chile	No		
	Columbia	No		
	Ecuador	No		
	Paraguay	No		
	Peru	**Yes**	+	+
	Uruguay	**Yes**	+	+
	Venezuela	No		

Note: − and + denote negative and positive signs, respectively. The long-run equilibrium relation (β_1) is normalized to SO_2 emissions; for example, a negative(positive) sign for income (openness) presents a negative (positive) relationship between SO_2 emissions and income (openness).

Table 3. Results of weak exogeneity tests for developing and developed countries

Continent	Country	Weak exogeneity ($H_0 : \alpha_i = 0$)		
Developed countries				
		$\ln(Emission_t)$	$\ln(Income_t)$	$Openness_t$
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				
Continent	Country	Weak exogeneity ($H_0 : \alpha_i = 0$)		
		$\ln(Emission_t)$	$\ln(Income_t)$	$Openness_t$
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	9.27**	5.00*	0.59
	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. ln represents natural logarithm. Values are the likelihood ratio (LR) test statistic based on the χ^2 distribution.

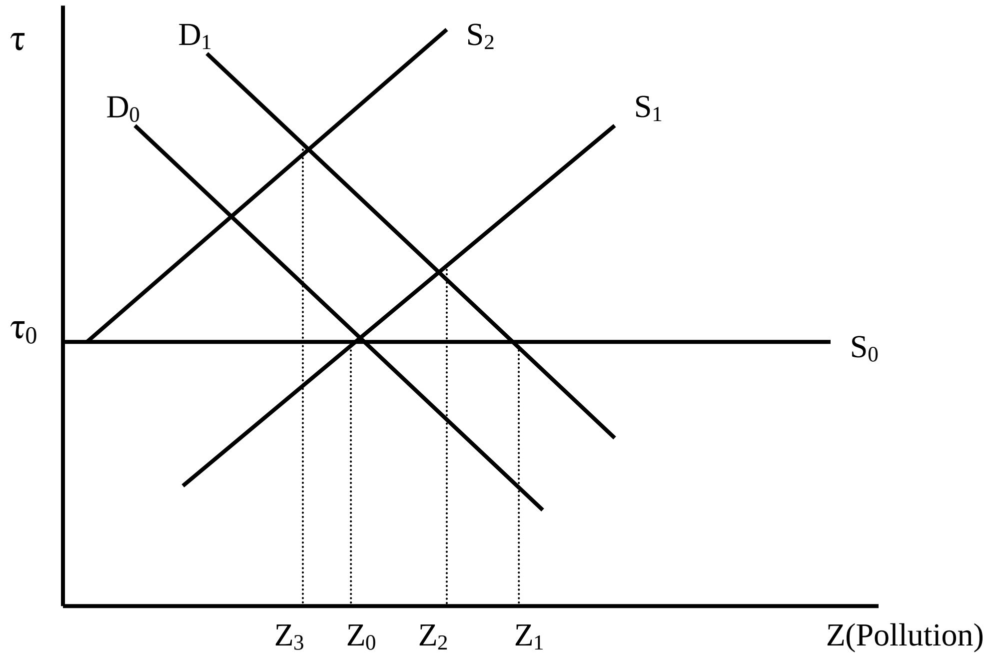

Figure 1. Effect of Trade Openness on the Environment

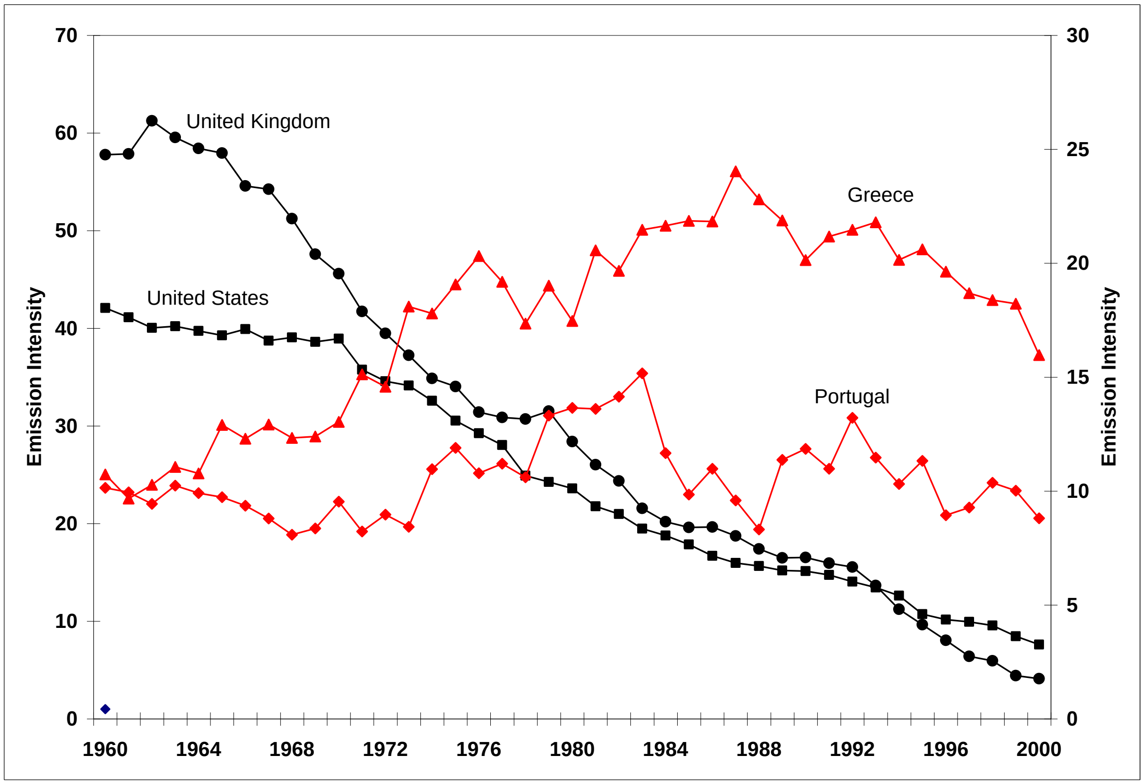

Figure 2. Emissions intensities for developed countries

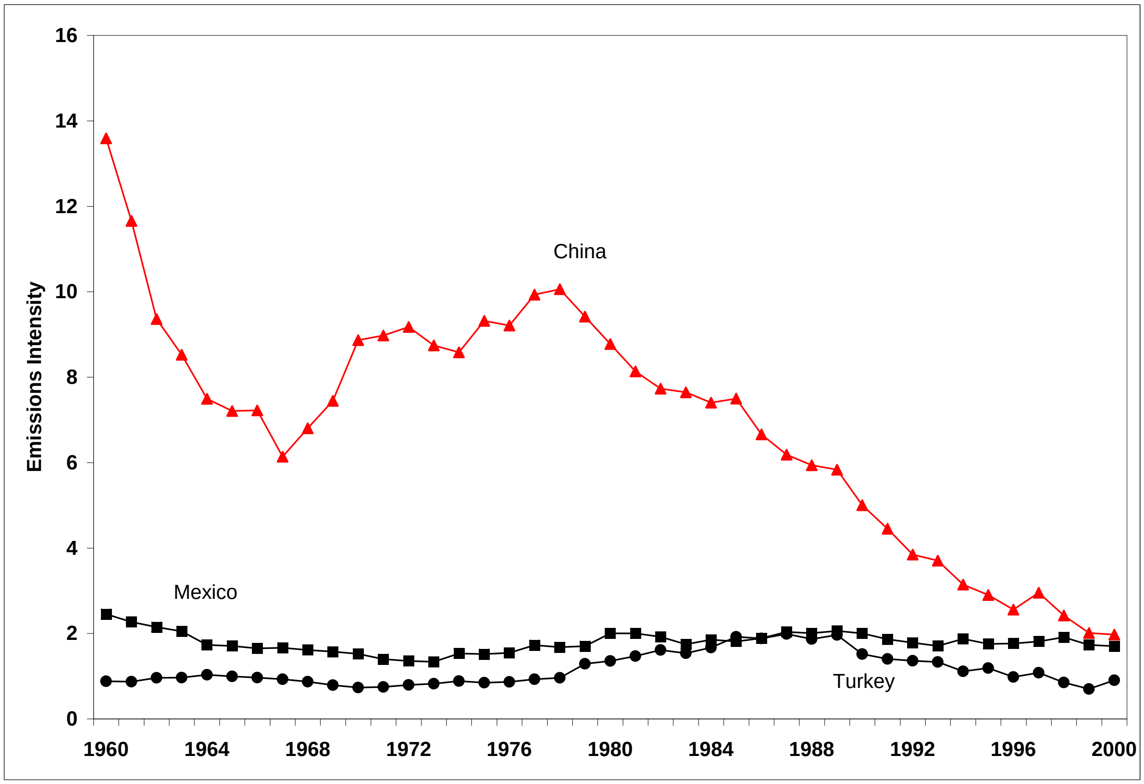

Figure 3. Emissions intensities for developing countries

REFERENCES

ASL and Associates. 1997. Sulfur emissions by country and year. Report No: DE96014790. U.S. Department of Energy, Washington, DC.

Chintrakarn, P., and Millimet, D.L. 2006. The environmental consequences of trade: evidence from subnational trade flows. Journal of Environmental Economics and Management 52(1): 430-453.

Coondoo, D., and Dinda, S. 2002. Causality between income and emission: a country group-specific econometric analysis. Ecological Economics 40(3): 351-367.

Copeland, B.R., and Taylor, M.S. 1994. North-South trade and the environment. Quarterly Journal of Economics 109(3): 755-787.

Copeland, B.R., and Taylor, M.S. 2004. Trade, growth, and the environment. Journal of Economic Literature 42(1): 7-71.

Copeland, B.R. 2005. Policy endogeneity and the effects of trade on the environment. Agricultural and Resource Economics Review 34(1): 1-15.

Dean, J. 2002. Does trade liberalization harm the environment? A new test. Canadian Journal of Economics 35(4): 819-842.

Dinda, S. 2004. Environmental Kuznets curve hypothesis: a survey. Ecological Economics 49(4): 431-455.

Dinda, S., and Coondoo, D. 2006. Income and emission: a panel data-based cointegration analysis. Ecological Economics 57(2): 167-181.

Doornik, J., and Hendry, D. 1994. Interactive econometric modeling of dynamic System (PcFiml 8.0). London: International Thomson Publishing.

Doornik, J., and Hendry, D. 2001. Empirical econometric modeling (PcGive 10). Timberlake Consultants Ltd, London, UK.

Elliot, G., Rothenberg, T.J., and Stock, J.H. 1996. Efficient tests for an autoregressive unit root. Econometrica 64(4): 813-836.

Engle, R., and Granger, C.W.J. 1987. Cointegration and error-correction: representation, estimation and testing. Econometrica 55: 251-276.

Frankel, J.A., and Rose, A.K. 2005. Is trade good or bad for the environment? Sorting out the causality. Review of Economics and Statistics 87(1): 85-91.

Gale, L.R., and Mendez, J.A. 1998. The empirical relationship between trade, growth and the environment. International Review of Economics and Finance 7(1): 53-61.

Gonzalo, J. 1994. Five alternative methods of estimating long-run equilibrium relationships. Journal of Econometrics 60(1-2): 203-233.

Grossman, G.M., and Krueger, A.B. 1991. Environmental impacts of the North American Free Trade Agreement. NBER. Working paper 3914.

Grossman, G.M., and Krueger, A.B. 1993. Environmental impacts of the North American Free Trade Agreement, in Peter Garber (Ed.), The U.S.-Mexico Free Trade Agreement. MIT Press, Cambridge.

Harris, R., and Sollis, R. 2003. Applied time series modeling and forecasting. Chichester, W. Sussex: John Wiley and Sons.

Johansen, S. 1995. Likelihood-based inference in cointegrated vector autoregressive models. Oxford: Oxford University Press.

Johansen, S., and Juselius, K. 1992. Testing structural hypotheses in a multivariate cointegration analysis of the PPP and the UIP for UK. Journal of Econometrics 53(1-3): 211-244.

Lefohn, A.S., Husar, J.D., and Husar, R.B. 1999. Estimating historical anthropogenic global sulfur emission patterns for the period 1850-1990. Atmospheric Environment 33(21): 3435-3444.

Lucas, R.E.B., David, W., and Hemamala, H. 1992. Economic development, environmental regulation and the international migration of toxic industrial pollution: 1960-1988, in Patrick Low (Ed.), *International trade and the environment*. World Bank Discussion Paper No. 159. Washington DC: World Bank, 1992, pp. 67-86.

Maddala, G.S., and Kim, I.M. 1998. Unit roots, cointegration, and structural change. Cambridge: Cambridge University Press.

Perman, R., and Stern, D.I. 2003. Evidence from panel unit root and cointegration tests that the environmental Kuznets Curve does not exist. Australian Journal of Agricultural and Resource Economics 47(3): 325-347.

Porter, M., and van der Linde, C. 1995. Toward a new conception of the environment-competitiveness relationship. Journal of Economic Perspectives 9(4): 97-118.

Sims, C. 1980. Macroeconomics and reality. Econometrica 48(1): 1-48.

Stern, D.I. 2005. Global sulfur emissions from 1850-2000. Chemosphere 58(2): 163-175.

Stern, D.I. 2006. Reversal of the trend in global anthropogenic sulfur emissions. Global Environmental Change 16(2): 207-220.

Wooldridge, J.M. 2000. Introductory econometrics: a modern approach. Mason, Ohio: South-Western College Publishing.