Trade Openness and Economic Growth: A Panel Causality Analysis

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Abstract

This paper examines the short-term and long-run dynamics between per capita GDP growth and openness for 158 countries over the period 1970-2009. We use panel cointegration tests and panel error-correction models (ECM) in combination with GMM estimation to explore the causal relationship between these two variables. We approach the problem of a potential endogeneity between openness and growth by including only growth rates and lagged values of the independent variable. Additionally, we apply Difference GMM and System GMM estimation. These estimators also address the issue of a possible correlation between the lagged endogenous variable and the error term. The results suggest a long-run relationship between openness and economic growth with a short-run adjustment to the deviation from the equilibrium for both directions of dependency. The long-run coefficients indicate a positive significant causality from openness to growth and vice versa, indicating that international integration is a beneficial strategy for growth in the long term. By contrast the short-run coefficient shows a negative short-run adjustment, suggesting that openness can be painful for an economy undergoing short-term adjustments. In addition to the entire panel we subdivide the data into income-related subpanels. While the long-run effect remains predominantly positive and significant, the short-run adjustment becomes positive when the income level increases. This result suggests that different trade structures in low-income and high-income countries have different effects on economic growth.

Keywords:	openness, trade, growth, development, panel cointegration,
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1 Introduction

The relationship between openness and economic growth has long been a subject of much interest and controversy in international trade literature. With regard to a theoretical relationship between openness and growth most of the studies provide support for the proposition that openness effects growth positively. Romer (1993), Grossman and Helpman (1991) and Barro and Sala-i-Martin (1995) among others, argue that countries that are more open have a greater ability to catch up to leading technologies of the rest of the world. Chang, Kaltani, Loayza (2005) point out that openness promotes the efficient allocation of resources through comparative advantage, allows the dissemination of knowledge and technological progress, and encourages competition in domestic and international markets. However there exists also the opposed position. For example Krugman (1994) and Rodrik and Rodríguez (2001) argue that the effect of openness on growth is doubtful. Furthermore, if we include the gains from trade debate we look at a long lasting debate discussing conditions and circumstances when openness and trade may be favourable and may improve the economic performance or not. These controversial theoretical findings also appear in the empirical literature. Numerous econometric studies have tried to identify the relationship and the causal direction between openness and economic growth. These studies can be divided into three groups. First, conventional regression analyses trying to capture the effect of openness by regressing it on per capita growth. There are several studies in this vein, with the vast majority concluding that openness to trade is a significant explanatory variable for economic growth (see e.g. some more recent contributions like Dollar (1992), Edwards (1998), Harrison (1996), Barro and Lee (1994), Easterly and Levine (2001), Dollar and Kraay (2002), Irwin and Tervio (2000), Islam (1995), Sala-i-Martin (1997)). However, conventional regression methods analyse only one direction of a possible bidirectional relationship and are unable to uncover the causation of trade openness and GDP growth. The second group of studies uses Granger causality based tests on the openness and economic growth variables. Here the results show a more mixed picture. Jung and Marshall (1985) employ the Granger causality test and find unidirectional causality from exports to growth for four out of 37 countries for the period 1951-1981. Chow (1987) analyses eight industrialized countries and finds bidirectional causality in six cases and one case with unidirectional causality from exports to growth. Hsiao (1987) uses Granger causality tests for Asian countries. The results show only unidirectional causality from growth to exports for the case of Hong Kong. Ahmad and Kwan (1991) investigate 47 African countries and find no causality between exports and growth. However, Bahmani-Oskooee (1991) applies Granger causality tests for 20 countries and finds both positive and negative causality effects for both directions. Although these kinds of studies try to capture a bidirectional relationship, Engle and Granger (1987) have shown that when the series are cointegrated a standard Granger-causality test is misspecified. The third group of studies picks up the problem of biased results in the event of cointegrated series and uses the concept of cointegration and error-correction to explore the short-run and long-run dynamics between openness and growth. These analyses are based on time series data and investigate the causalities at country level. Islam (1998) uses an error-correction model for each of 15 Asian countries[1] for the period 1967-1991 and finds export to growth causality for 10 out of 15 countries. Liu et al. (2002) investigate China over the period 1981-1997 and find bidirectional causality between FDI, exports and growth. Bouoiyour (2003) applies the concept of cointegration and error correction to the relationship between trade and economic growth in Morocco over the period 1960-2000. The results show a lack of long-run causality. In the short run higher imports and exports cause higher GDP. Awokuse (2007) examines the impact of export and import expansion on growth in three transition economies on country level. The results show bidirectional causality between exports and growth for Bulgaria, the Czech Republic exhibits unidirectional causality from exports and imports on growth, and for Poland only the import-led growth hypothesis can be supported. A weak point of these studies is the absence of a general examination of the causality between openness and economic growth on cross-country level. All these studies investigate the relationship on a certain country level for time series data. A general panel error correction model has not been applied yet.

Therefore, this study aims to address this problem and to re-examine the issue of causal links between trade openness and growth using an error correction model for a panel of 158 countries over the period 1970-2009. This methodological framework allows us to test for bidirectional causality relations from openness to GDP and vice versa. Furthermore, this method enables us to distinguish between short-run and long-term effects between trade and growth. In particular we shed light on the question of whether benefits of trade or fears of negative effects of trade address different time horizons. That is, finding a long-term positive causality from trade to growth would provide evidence on long-term benefits of international integration. Ba contrast, the presence of negative causal effects in the short term would be an indicator of the pain of adjustment an economy has to sustain if long-term benefit is the target. We also break down our data set into five subpanels following the World Bank income classification to allow us to investigate income-related effect differences. The paper proceeds as follows. Section 2 presents the empirical investigation including the data, the methodology and the results of the panel unit root tests, the panel cointegration tests, and the error-correction model. Section 3 concludes.

2 Empirical evidence

This section investigates the causal relationship between trade openness and economic growth. In a first step we use recently developed panel unit root and cointegration tests. Then we apply panel-based error-correction models to

[1] The countries included in the analysis are Bangladesh, Fiji, Hong Kong, India, Indonesia, Japan, Malaysia, Nepal, Pakistan, Papua New Guinea, Philippines, Singapore, South Korea, Sri Lanka, and Thailand.

explore the bidirectional short-run and long-run dynamics between these two variables. In our analysis we try to capture the problem of possible endogeneity of the openness variable. Our starting point is to test for causality of openness on growth and vice versa. However, a loop of causality between the independent and dependent variables of a model leads to endogeneity problems in a simple regression model. Previous studies of causality between openness and growth ignore this potentially interdependence which leads to a correlation between the endogenous variable and the error term. These kinds of studies disregard a potential endogeneity of openness and growth and so produce biased and inconsistent parameter estimates. Current literature identifies this problem and provides two common suggestions to deal with it. The first is the use of lagged values of the exogenous variables. Second, the endogeneity problem can also be addressed appropriately by using instrumental variables (IV) techniques. So if a dependent variable is potentially endogenous, it is intuitively appealing to look for a proxy that does not suffer from the same problem. For example Frankel and Romer (1999) use geographic attributes as instruments to identify the effects of trade on income. This approach is also adopted by Irvin and Terviö (2002) and Noguera and Siscart 2005). We abstract from these geographical determinants of trade by focusing on changes in openness. Our model includes only growth rates and lagged values of the independent variable. Additionally, we approach this problem by using GMM estimation. It also addresses the issue of a possible correlation between the lagged endogenous variable and the error term. We use the Difference GMM and System GMM estimators developed by Arellano and Bond (1991) and Blundell and Bond (1998). These estimators deal effectively with the endogeneity problem by using a set of instruments for the endogenous variables. The former uses lagged levels as instruments for the equation in differences; in addition to that the latter uses lagged differences as instruments for the additional equations in levels. Furthermore, concerning the substance we explore both directions of action between trade and growth using a simultaneous bivariate model.

2.1 Data

Regarding openness there are several variables that can be used to measure the degree of openness. They can be broadly divided into two categories: measures of trade volumes and measures of trade restrictions. The most common measure in the first group is trade share, which is the sum of exports plus imports divided by GDP. The second category includes measures of trade barriers that include average tariff rates, export taxes, total taxes on international trade, and indices of non-tariff barriers. To perform a broad panel analysis of a large number of countries and over a long period we select a measure that is widely available, namely trade share. We use a balanced panel data set containing 158 countries over the period 1970-2009. Our analysis is based on two variables from the Penn World Table 7.0. provided by Heston et al. (2011):

Openness: $openness_{i,t}$ Trade measured by the sum of exports and imports as a percentage of GDP at 2005 constant prices. The variable corresponds to the *openk* variable from the Penn World Table.

GDP per capita: $GDP_{i,t}$ GDP per capita is PPP converted GDP per capita (Laspeyres) at 2005 constant prices in international dollar per person. The variable corresponds to the *rgdpl* variable from the Penn World Table.

In addition to the entire panel, we segment the data set into five subpanels according to per capita income. We use the World Bank country classification that distinguishes between low-income economies ($1005 or less), lower-middle-income economies ($1006 to $3975), upper-middle-income economies ($3976 to $12275), high-income economies ($12276 or more) and high-income OECD members. Table 1 provides the descriptive statistics of the panel data set and the subpanels.[2]

Table 1: Descriptive Statistics

panel	1	2-LI	3-LMI	4-UMI	5-HI	6-OECD
observations	6320	1200	1800	1640	640	1040
openness						
mean	75.52	50.38	80.88	76.13	133.08	58.85
std. dev.	50.23	29.46	39.34	44.76	74.25	44.91
GDP per capita						
mean	8797.21	793.34	2734.28	6557.28	21164.17	24447.75
std. dev.	11083.70	349.01	1721.35	3451.02	14772.27	10193.08

2.2 Estimation

Methodology: To explore the short-run and long-run dynamics between GDP growth and changes in openness we follow Yasar et al. (2006) and apply a generalized one-step error-correction model (ECM) in combination with panel data and GMM estimation. We prefer dynamic panel estimators for various reasons. GMM estimation circumvents the bias associated with including a lagged dependent variable as a regressor and enables us to calculate consistent and efficient estimates. Additionally, by combining the time series dimension with the cross-sectional dimension, the panel data provides a richer set of information to exploit the relationship between the dependent and independent variables, reduces collinearity among the explanatory variables, increases the degrees of freedom, and gives more variability and efficiency. More specifically, our point of departure is a bivariate autoregressive-distributed lag model

[2] Appendix 1 provides a detailed list of all included countries and the income segmentation.

$$y_{i,t} = \alpha_0 + \sum_{j=1}^{2} \alpha_j y_{i,t-j} + \sum_{j=0}^{2} \delta_j x_{i,t-j} + f_i + u_{i,t} \quad (1)$$

$$x_{i,t} = \beta_0 + \sum_{j=1}^{2} \beta_j x_{i,t-j} + \sum_{j=0}^{2} \gamma_j y_{i,t-j} + \eta_i + \nu_{i,t} \quad (2)$$

where index i=1...N refers to the country and t=1...T to the period. This method allows us to include specific effects for each country (f_i and η_i). This individual effect may correlate with the included explanatory variables, hence omitting the individual effect would become part of the error term, which would lead to a bias in the estimates. The disturbances $u_{i,t}$ and $\nu_{i,t}$ are assumed to be independently distributed across countries with a zero mean. They may display heteroskedasticity across time and countries, though. Following Granger (1969) there is Granger causality from x to y if past values of x improve the prediction of y given the past values of y. With respect to the model x Granger causes y if not all δ_j are zero. By the same token Granger causality from y to x occurs if not all γ_j are equal to zero. However, Engle and Granger (1987) have shown that, if the series x and y are cointegrated, the standard Granger causality test is misspecified. In this case an ECM should be used instead. In a first step we have to apply a unit root and a cointegration test. On the basis of the results we determine whether to use the Granger causality framework or an ECM model to test causality.

Panel unit root test: The Granger causality test requires the variables to be stationary. We check their stationarity using two common panel unit root tests, the IPS test by Im, et al. (2003) and the Fisher-type test by Maddala and Wu (1999) and Choi (2001).

Formally, the test equation of both tests is

$$\triangle y_{i,t} = \mu_i + \beta_i y_{i,t-1} + \varepsilon_{it}, \quad (3)$$

with the null hypothesis that each cross-section series in the panel has a unit root and the alternative hypothesis that at least one cross-section in the panel is stationary. Additionally, the formulation allows β_i to differ across cross-sections so that both tests allow for heterogeneity.

$$H_0 : \beta_i = 0 \quad for\ all\ i \quad (4)$$

$$H_1 : \beta_i < 0, \quad i = 1, 2, ..., N_1, \quad \beta_i = 0, \quad i = N_1 + 1, N_2 + 1, ...N. \quad (5)$$

Table 2: Panel unit root test

variable	deterministic	IPS	Fisher-type	first diff. IPS	first diff. Fisher-type
openness	constant	1.0752	319.487	-40.5784***	2176.00***
	const. + trend	-0.2000	329.420	-34.8509***	1744.49***
GDP	constant	5.1565	250.553	-34.1635***	1787.10***
	const. + trend	0.7044	334.263	-30.0096***	1472.30***

Notes:
* Rejects the null of a unit root at the 10% level.
** Rejects the null of a unit root at the 5% level.
*** Rejects the null of a unit root at the 1% level.

The IPS test is a t-bar statistic based on the augmented Dickey-Fuller statistic (Dickey and Fuller 1979). The test statistic is computed by the sample mean of the individual unit root tests for each of the N cross-section units. The main idea of the Fisher-type unit root test is to combine p-values from the unit root tests applied to each of the N cross-section units in the panel. While both IPS and the Fisher-type test combine information based on individual unit root tests, the crucial difference between the two is that the IPS test combines the test statistics while the Fisher-type test combines the significance levels of the individual tests. Table 2 presents the results of the tests for both variables in levels and in first differences. The results indicate that for both variables the level data is non-stationary, however the test statistics of the differenced variables are highly significant and show stationarity regardless of whether the trend is included in the test or not.[3] Hence, the following analysis is based on the differenced data, namely GDP growth and changes in openness.

Panel cointegration test: Since the panel unit root tests presented above indicate that the variables are integrated of order one I(1), we test for cointegration using the panel cointegration test developed by Pedroni (1999, 2004). This test allows for heterogeneity in the panel by permitting heterogenous slope coefficients, fixed effects and individual specific deterministic trends. The test contains seven cointegration statistics, the first four based on pooling the residuals along the "within-dimension" which assume a common value for the unit root coefficient, and the subsequent three based on polling the residuals along the "between dimension" which allow for different values of the unit root coefficient. The common idea of both classes is to first estimate the hypothesized cointegration relationship separately for each group member of the panel and then pool the resulting residuals when constructing the test for the null of no

[3] Additionally, we test also the homogeneous alternative of the H_1 hypothesis that assumes that the autoregressive parameter is identical for all cross-section units. The Levin, Lin and Chu (2002) and Breitung (2000) unit root tests also indicate that the variables are integrated of order on I(1).

cointegration. Table 3 presents the results. In all cases the null of no cointegration is rejected at the 1% significance level, indicating that the variables exhibit a cointegration relationship.

Table 3: Panel cointegration test

	openness - GDP
Panel υ-test	4.3802***
Panel ρ-test	-5.7259***
Panel *pp*-test	-6.3860***
Panel *adf*-test	-5.9561***
Group ρ-test	-3.2054***
Group *pp*-test	-5.9503***
Group *adf*-test	-6.4598***

Notes:
Based on individual intercept and automatic lag length selection based on AIC.
* Rejects the null of no cointegation at the 10% level.
** Rejects the null of no cointegation at the 5% level.
*** Rejects the null of no cointegation at the 1% level.

Error correction model: Engle and Granger (1987) have shown that when the series x and y are cointegrated a standard Granger-causality test as presented in the equations (1) and (2) is misspecified, because it does not allow for the distinction between the short-run and the long run-effect. At this point a error correction model (ECM) should be used instead. It is a linear transformation of the ADL models above and provides a link between the short-run and the long-run effect (Banerjee et al. 1993, 1998).

$$\begin{aligned}\Delta y_{i,t} &= (\alpha_1 - 1)\Delta y_{i,t-1} + \delta_0 \Delta x_{i,t} + (\delta_0 + \delta_1)\Delta x_{i,t-1} \\ &\quad + \lambda(y_{i,t-2} - \phi x_{i,t-2}) + f_i + u_{i,t} \qquad (6) \\ \Delta x_{i,t} &= (\beta_1 - 1)\Delta x_{i,t-1} + \gamma_0 \Delta y_{i,t} + (\gamma_0 + \gamma_1)\Delta y_{i,t-1} \\ &\quad + \kappa(x_{i,t-2} - \phi y_{i,t-2}) + \eta_i + \nu_{i,t} \qquad (7)\end{aligned}$$

While the coefficients $(\alpha_1 - 1)$, δ_0 and $(\delta_0 + \delta_1)$ as well as $(\beta_1 - 1)$, γ_0 and $(\gamma_0 + \gamma_1)$ capture the short-run effects, the coefficients λ and κ of the error correction terms give the adjustment rates at which short-run dynamics converge to the long-run equilibrium relationship. If λ and κ are negative and significant a relationship between x and y exist in the long run. The standard error-correction procedure is a two-step method where in a first step the error correction term is obtained by saving residuals of separate estimation of the long-run equilibrium of x and y. In a second step the ECM with the included error correction term can be estimated. However, the two-stage error correction models have been criticized in the literature. Banerjee et al. (1998) argue that there can be a substantial small-sample bias compared to a single-equation error correction model where the long-run relation is restricted to being homogeneous. Accordingly, in this study we use a one-step procedure to indicate the short-run and long-run dynamics. The generalized one-step ECM is transformed as follows:

$$\begin{aligned}
\Delta y_{i,t} &= (\alpha_1 - 1)\Delta y_{i,t-1} + \delta_0 \Delta x_{i,t} + (\delta_0 + \delta_1)\Delta x_{i,t-1} \\
&\quad + \lambda(y_{i,t-2} - x_{i,t-2}) + \theta x_{i,t-2} + f_i + u_{i,t} \qquad (8) \\
\Delta x_{i,t} &= (\beta_1 - 1)\Delta x_{i,t-1} + \gamma_0 \Delta y_{i,t} + (\gamma_0 + \gamma_1)\Delta x_{i,t-1} \\
&\quad + \kappa(x_{i,t-2} - y_{i,t-2}) + \Psi x_{i,t-2} + \eta_i + \nu_{i,t} \qquad (9)
\end{aligned}$$

where the long-run multiplier is restricted to being homogeneous $\phi = 1$. Using this form of the error correction model allows us to calculate the true long-run relationship between x and y, which can be written as $1 - (\hat{\theta}/\hat{\lambda})$ and $1 - (\hat{\Psi}/\hat{\kappa})$. Hence, the one step ECM permits us to directly calculate the short-run and long-run elasticities between openness and growth. To avoid the problem of biased estimates through a possible correlation between the lagged endogenous variable and the error term we use the Difference GMM and System GMM estimators developed by Arellano and Bond (1991) and Blundell and Bond (1998). The former uses all lagged observations to instrument the lagged endogenous variable and circumvent a possible bias. The latter combines the regression in differences with the regression in levels in a system and uses additional instruments in levels. The moment conditions of the instruments of both estimators can be verified using the Sargan statistic that tests the validity of all instruments.

2.3 Results

The results of the corresponding error-correction regressions are summarized in table 4. They include the coefficients of the regressions, the summation of the short-run and long-run effects with the corresponding Wald test p-values, the Sargan tests and the M1 and M2 tests for the regressions. The first two columns explore the dynamics of openness on GDP growth and contain the results with reference to equation (8), while the third and fourth column investigate the other direction of causality and are consequently based on equation (9). The results of the income subpanels are presented in tables 5 and 6.

To verify GMM consistency, we have to make sure that the instruments are valid. We use the Sargan test of over-identifying restrictions to test the validity of the instrumental variables. The null hypothesis assumes that the orthogonality conditions of the instrumental variables are satisfied. In all cases the p-values show satisfactory results, indicating that the instruments used for the estimation are valid. We also consider the test of second-order serial correlation of the error term suggested by Arellano and Bond (1991). If the null hypothesis of no second-order serial correlation of the error term cannot be rejected, the GMM estimator is valid.

The coefficients of the error-correction term give the adjustment rate at which short-run dynamics converge to the long-run equilibrium relationship. With cover to our specifications it is the adjustment rate at which the gap between openness and growth is closed. Generally, all these coefficients are negative and highly significant as expected, so the results show that there exists a long-run relationship and provide evidence of a cointegration relationship between the variables.

The short-run effect can be divided into the effect of the lagged dependent variable and that of the independent variable. The short-time adjustment of the independent variable is measured by the effect of the contemporaneous and lagged change of the independent variable. The significance of the summarized short-run effects, which is simply the sum of the two coefficient values, is tested via a Wald test. The long-run coefficients indicate the long-run elasticities of the independent on the dependent variable. They are computed by subtracting the ratio of the coefficient of the scale effect (lag of independent variable) and the coefficient of the error-correction term from one; again, a Wald test proves the significance of the effect.

Table 4: Estimated error-correction model

model	Dependent Variables			
	$\Delta \ln GDP$		$\Delta \ln openn$	
	Diff.GMM	Sys.GMM	Diff.GMM	Sys.GMM
$\Delta \ln openn$	-0.0480***	-0.0609***		
	(0.0009)	(0.0009)		
$\Delta \ln openn_{t-1}$	0.0084***	-0.0051***	-0.1441***	-0.1199***
	(0.0006)	(0.0006)	(0.0029)	(0.0020)
$\ln openn_{t-2}$	-0.0044	0.0053		
	(0.0036)	(0.0010)		
$\ln openn_{t-2} - \ln GDP_{t-2}$			-0.1300***	-0.1082***
			(0.0031)	(0.0014)
$\Delta \ln GDP$			-0.1261***	-0.1355***
			(0.0015)	(0.0025)
$\Delta \ln GDP_{t-1}$	0.0097**	0.0357***	0.0917***	0.0744***
	(0.0042)	(0.0012)	(0.0015)	(0.0015)
$\ln GDP_{t-2}$			-0.0968***	-0.0927***
			(0.0033)	(0.0014)
$\ln GDP_{t-2} - \ln openn_{t-2}$	-0.0616***	-0.0345***		
	(0.0036)	(0.0006)		
Summation				
Short-run coefficient	-0.0396	-0.0660	-0.0343	-0.0612
Wald test (P-value)	0.0000	0.0000	0.0000	0.0000
Long run coefficient	0.9280	1.1550	0.2555	0.1434
Wald test (P-value)	0.0000	0.0000	0.0000	0.0000
Sargan test (P-value)	1.0000	1.0000	1.0000	1.0000
AR1 (P-value)	0.0000	0.0000	0.0000	0.0000
AR2 (P-value)	0.1353	0.1088	0.2083	0.2266
Observations	5846	6004	5846	6004

Notes:
(1) Estimation based on the Difference GMM and the System GMM estimator.
(2) Asymptotically robust standard errors reported in parentheses.
(3) Sargan test is based on the estimation with GMM standard errors.
(4) *, ** and *** denote significance at the 10%, 5% and 1% level.

- With regard to the first ECM presented in the first two columns of table 4 all coefficients except the lagged openness variable are significant at the 5% significance level at least. As expected the error-correction term is negative and significant, indicating that there is a long-run relationship between growth and openness. Furthermore, the significant error-correction

Table 5: Estimated error-correction model: long-run and short-run dynamics of changes in openness on growth for subpanels

model	Dependent Variables $\Delta \ln GDP$									
	low-income		lower-middle-income		upper-middle-income		high-income		OECD	
	Diff.GMM	Sys.GMM	Diff.GMM	Sys.GMM	Diff.GMM	Sys.GMM	Diff.GMM	Sys.GMM	Diff.GMM	Sys.GMM
$\Delta \ln GDP_{t-1}$	0.0240	0.0244	-0.1248***	-0.0673***	0.0622	-0.0153	-0.0631	0.2012***	0.2438***	0.2785***
	(0.1118)	(0.0450)	(0.0331)	(0.0086)	(0.0753)	(0.0177)	(0.2162)	(0.0522)	(0.0486)	(0.0661)
$\Delta \ln openn$	-0.1348***	-0.1338***	-0.0367***	-0.0728***	-0.0056	0.0018	0.1745*	0.1378***	0.2168***	0.2461***
	(0.124)	(0.0166)	(0.0096)	(0.0117)	(0.0094)	(0.0088)	(0.0915)	(0.0352)	(0.0128)	(0.0154)
$\Delta \ln openn_{t-1}$	-0.0112	-0.0022	-0.0039	-0.0164***	-0.0024	-0.0230***	0.1842	0.1457***	0.0747***	0.0709***
	(0.0168)	(0.0106)	(0.0035)	(0.0017)	(0.0089)	(0.0086)	(0.1512)	(0.0169)	(0.0159)	(0.0199)
$\ln GDP_{t-2} - \ln openn_{t-2}$	-0.0246	-0.0336	-0.1442***	-0.0840***	-0.0214	-0.0671***	-0.2155**	-0.0724*	-0.0651***	-0.0590
	((0.0993)	(0.0389)	(0.0341)	(0.0087)	(0.0599)	(0.0172)	(0.0863)	(0.0392)	(0.0246)	(0.0395)
$\ln openn_{t-2}$	-0.0044	-0.0197	-0.1048***	-0.0384***	0.0026	-0.0356***	0.0033	0.0392**	-0.0235**	-0.0204
	(0.1074)	(0.0437)	(0.0353)	(0.0082)	(0.0502)	(0.0134)	(0.2292)	(0.0171)	(0.0119)	(0.0175)
Summation										
Short-run coef.	-0.1461	-0.1361	-0.0407	-0.0892	-0.0080	-0.0212	0.3588	0.2834	0.2915	0.3170
Wald test (P-value)	0.000	0.000	0.000	0.000	0.643	0.173	0.136	0.000	0.000	0.000
Long run coefficient	0.8199	0.4133	0.2730	0.5432	1.1207	0.4688	1.0155	1.5418	0.6384	0.6546
Wald test (P-value)	0.821	0.493	0.000	0.000	0.677	0.000	0.340	0.001	0.000	0.000
Sargan test (P-value)	1.0000	1.000	1.000	1.000	1.000	1.000	1.000	1.000	1.000	1.000
AR1 (P-value)	0.0016	0.0015	0.0080	0.0081	0.0354	0.0513	0.1592	0.0123	0.0004	0.0004
AR2 (P-value)	0.4845	0.4738	0.0392	0.0296	0.3381	0.2030	0.1875	0.0774	0.2217	0.2961
Observations	1110	1140	1665	1710	1517	1558	592	608	962	988

Notes:
(1) Estimation based on the Difference GMM and the System GMM estimator.
(2) Asymptotically robust standard errors reported in parentheses.
(3) Sargan test is based on the estimation with GMM standard errors.
(4) *, ** and *** denote significance at the 10%, 5% and 1% level.

Table 6: Estimated error-correction model: long-run and short-run dynamics of growth on changes in openness for subpanels

model	Dependent Variables $\Delta \ln openn$									
	low-income		lower-middle-income		upper-middle-income		high-income		OECD	
	Diff.GMM	Sys.GMM	Diff.GMM	Sys.GMM	Diff.GMM	Sys.GMM	Diff.GMM	Sys.GMM	Diff.GMM	Sys.GMM
$\Delta \ln openn_{t-1}$	-0.2181	-0.3076***	-0.1576**	-0.1284***	-0.3318***	-0.1075***	-0.1905	-0.0334	-0.0801	0.0721*
	(0.1867)	(0.5808)	(0.0748)	(0.0225)	(0.0834)	(0.0087)	(0.1429)	(0.0600)	(0.0762)	(0.0381)
$\Delta \ln GDP$	-0.5135***	-0.5148***	-0.1964***	-0.1512***	0.0939***	0.0847***	0.2728***	0.3005***	0.5686***	0.5694***
	(0.0458)	(0.0425)	(0.0479)	(0.0225)	(0.0263)	(0.0258)	(0.0735)	(0.0641)	(0.0369)	(0.0283)
$\Delta \ln GDP_{t-1}$	0.0126	-0.0608	-0.0667	0.0321	0.1599***	0.1213***	-0.0279	-0.0631**	-0.0252	-0.1576***
	(0.1098)	(0.0567)	(0.0738)	(0.0222)	(0.0144)	(0.1497)	(0.0723)	(0.0266)	(0.0454)	(0.0354)
$\ln openn_{t-2} - \ln GDP_{t-2}$	-0.1605	-0.2704***	-0.1949***	-0.1341***	-0.3330***	-0.1193***	-0.2082	-0.0805	-0.1359**	-0.0144
	(0.1958)	(0.0606)	(0.0720)	(0.2155)	(0.0873)	(0.0082)	(0.1471)	(0.0617)	(0.0661)	(0.0254)
$\ln GDP_{t-2}$	-0.2912	-0.4410***	-0.3573***	-0.2087***	-0.2307***	-0.0810***	-0.1989	-0.0935	-0.0381*	-0.0132*
	(0.2717)	(0.1338)	(0.1574)	(0.0429)	(0.0710)	(0.0107)	(0.1417)	(0.0575)	(0.0218)	(0.0076)
Summation										
Short-run coef.	-0.5009	-0.5756	-0.2631	-0.1291	0.2538	0.2060	0.2448	0.2374	0.5434	0.4119
Wald test (P-value)	0.000	0.000	0.029	0.003	0.000	0.000	0.000	0.000	0.000	0.000
Long run coefficient	-0.8143	-0.6307	-0.8334	-0.5559	0.3070	0.3209	0.0443	-0.1606	0.7199	0.0796
Wald test (P-value)	0.448	0.085	0.141	0.069	0.000	0.000	0.616	0.607	0.000	0.959
Sargan test (P-value)	1.0000	1.000	1.000	1.000	1.000	1.000	1.000	1.000	1.000	1.000
AR1 (P-value)	0.0013	0.0001	0.0000	0.0000	0.0165	0.0037	0.1409	0.0990	0.0004	0.0000
AR2 (P-value)	0.4020	0.5130	0.2588	0.4941	0.2257	0.3471	0.2792	0.2721	0.9870	0.9010
Observations	1110	1140	1665	1710	1517	1558	592	608	962	988

Notes:
(1) Estimation based on the Difference GMM and the System GMM estimator.
(2) Asymptotically robust standard errors reported in parentheses.
(3) Sargan test is based on the estimation with GMM standard errors.
(4) *, ** and *** denote significance at the 10%, 5% and 1% level.

coefficient implies that when there are deviations from long-run equilibrium, short-run adjustments in openness will re-establish the long-run equilibrium. The absolute value of the term provides the speed of the short-run adjustment process, indicating that about six per cent (-0.0616) of the discrepancy in the case of the Difference GMM estimator and about three per cent (-0.0345) in the case of the System GMM estimator are corrected in each period. Aside from speed of adjustment the results indicate the magnitude of the short-run effect. It is measured by the sum of the contemporaneous and lagged dependent variable and indicates a negative significant causal effect from changes in openness on growth. The long-run effect is positive significant, indicating that in the long run changes in openness cause higher GDP growth. These short-term and long-run results show that the debate of free trade versus protectionism in the international trade literature should not be considered as two contradictory aspects. Rather, foreign competition seems to have a negative short-term effect on growth. At the firm level, in particular import-competing firms are disadvantaged and seek protection against openness. However, the results show that in the long run free trade policies prove beneficial to productivity and growth, which is consistent with recent literature that suggests that openness promotes economic development through various channels, such as technological progress, increasing key markets and rising competition.

- Concerning the other direction of the causal relationship presented in the third and fourth column of table 4, all coefficients as well as the short- and long-run effects are significant at the 1% level for both estimators. Again the error correction term is negative and significant. This intensifies the long-run relationship with a short-run adjustment to equilibrium that we already found in the first model. At about 13 and 11 per cent (-0.1300 for Difference GMM and -0.1082 for System GMM), respectively, the speed of adjustment is higher than in the reversed model. The short-run effect of GDP growth on openness is negative and at nearly the same level as the effect of the other direction, indicating that the short-run openness response to a temporary growth shock is of the same magnitude as the growth response of a temporary openness shock. The long-run effect is positive and significant, indicating that higher GDP growth causes greater changes in openness. The magnitude of this effect is lower than that of the effect of openness on growth. This implies that the long-run growth response to permanent shocks in openness tends to be greater than the openness response to permanent changes in growth; hence growth is more sensitive to openness than vice versa. Again, the results show that even though in the long run openness seems to be beneficial for growth, in the short term negative growth shocks may hit the economy and invoke a call for protectionism.

- Tables 5 and 6 present the results for the income subpanels, which are low-income economies, lower-middle-income economies, upper-middle-income economies, high-income economies and high-income OECD members. The former table shows the causality from openness to GDP, the latter the reverse direction. When GDP growth is the dependant variable the long-run effect remains positive when significant. However, we are unable to reveal a long-run causal effect for the low-income economies as well as for the difference estimator of the upper-middle-income and high-income economies. The short-run effect is significant except for the upper-middle-income economies and the difference estimator of the high-income economies. The results change depending on income. While the coefficients of the poorer subpanels are negative, the results of the high-income subpanels show a positive short-run effect. This suggests that not only the trade level but also the structure of trade should be taken into consideration. For example, Hausmann et al. (2007) and Rodrik (2006) suggest that the structure of export products matters to growth, while Lederman and Maloney (2003) show that trade has a different effect on growth depending on its structure in terms of natural resource abundance, export concentration, and intra-industry trade. This is consistent with our results that suggest that because of a different trade structure the effect of trade in developing countries is different from that of trade in the industrialized countries. A proposal for further research could be to analyse the causality of trade openness and growth as a function of trade structure. Concerning the other direction of causality when changes in openness are the dependent variable, both the short-run and the long-run effect exhibit a change in sign from negative for lower-income countries to positive for higher-income countries. Furthermore, the higher the average income the greater the effect of growth on openness. The results show that economic growth only effects trade openness positively above a certain income level (upper-middle-income), while in lower-income countries growth seems to impede openness.

In summary the overall results of the estimated ECMs for the entire panel suggest a bidirectional positive long-run causality between GDP growth and openness, indicating that openness promotes economic development and vice versa. The short-run results suggest negative effects pointing towards required painful adjustments which often elicit a call for protectionism. That said, deeper income subpanel analyses indicate that this result has needs to be differentiated by income groups. Whereas higher-income countries exhibit positive causalities for both directions for the short-run and long-run effect, lower-income countries have a negative short-run adjustment and a positive long-run affect from openness to growth but a negative long-run effect in the other direction.

3 Summary and conclusion

The present study examines the causal relationship between trade openness and economic growth. After reviewing recent empirical research regarding the link between openness and growth we use recent panel estimation methods to explore the causal relationship between these variables. In a first step we check for stationarity using two common panel unit root tests, the IPS test and the Fisher-type test. After that we apply a panel cointegration test on openness and growth. As the variables are cointegrated we use panel ECMs to explore the bilateral short-run and long-run dynamics between these variables.

We approach the problem of a potential endogeneity between openness and growth by including only growth rates and lagged values of the independent variable. Additionally, we use instrumental variables and the Difference GMM and System GMM estimators developed by Arellano and Bond (1991) and Blundell and Bond (1998) These estimators also address the issue of a possible correlation between the lagged endogenous variable and the error term. The results suggest that the long-run causality between trade openness and growth runs in both directions. This is in line with Harrison (1996) who argues that although more open trade policies do precede higher growth rates, it is also true that higher growth rates lead to more open trade regimes. The short-run adjustment for both directions is negative. However, additional analyses for income-grouped subpanels show that apart from the long-run effect of openness on growth which is persistently positive for all subpanels the effect changes in sign depending on income. While the lower-income subpanel shows a negative causality, the high-income countries exhibit a positive relationship between growth and openness. The desired growth-led openness and openness-led growth hypothesis can only be supported for industrialized countries. In developing countries only the long-run openness-led growth hypothesis holds, while growth seems to slow down openness in the long run.

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4 Appendix

Appendix 1

The countries included in the analysis are

low-income economies ($1005 or less GDP per capita):

Afghanistan, Bangladesh, Benin, Burkina Faso, Burundi, Cambodia, Central African Republic, Chad, Comoros, Dem. Rep Congo, Ethiopia, The Gambia, Guinea, Guinea-Bissau, Haiti, Kenya, Liberia, Madagascar, Malawi, Mali, Mozambique, Nepal, Niger, Rwanda, Sierra Leone, Somalia, Tanzania, Togo, Uganda, Zimbabwe

lower-middle-income economies ($1006 to $3975 GDP per capita):

Angola, Belize, Bhutan, Bolivia, Cameroon, Cape Verde, Republic of Congo, Cote d'Ivoire, Djibouti, Egypt, El Salvador, Fiji, Ghana, Guatemala, Guyana, Honduras, India, Indonesia, Iraq, Kiribati, Laos, Lesotho, Marshall Islands, Mauritania, Fed. Sts. Micronesia, Mongolia, Morocco, Nicaragua, Nigeria, Pakistan, Papua New Guinea, Paraguay, Philippines, Samoa, Sao Tome and Principe, Senegal, Solomon Islands, Sri Lanka, Sudan, Swaziland, Syria, Tonga, Vanuatu, Vietnam, Zambia

upper-middle-income economies ($3976 to $12275 GDP per capita):

Albania, Algeria, Antigua and Barbuda, Argentina, Botswana, Brazil, Bulgaria, Chile, China Version 1, China Version 2, Colombia, Costa Rica, Cuba, Dominica, Dominican Republic, Ecuador, Gabon, Grenada, Iran, Jamaica, Jordan, Lebanon, Malaysia, Maldives, Mauritius, Mexico, Namibia, Panama, Peru, Romania, Seychelles, South Africa, St. Kitts & Nevis, St. Lucia, St.Vincent & Grenadines, Suriname, Thailand, Tunisia, Turkey, Uruguay, Venezuela

high-income economies ($12276 or more GDP per capita):

Bahamas, Bahrain, Barbados, Bermuda, Brunei, Cyprus, Equatorial Guinea, Hong Kong, Israel, Macao, Malta, Oman, Puerto Rico, Singapore, Taiwan, Trinidad & Tobago

high-income OECD members:

Australia, Austria, Belgium, Canada, Denmark, Finland, France, Germany, Greece, Hungary, Iceland, Ireland, Italy, Japan, Republic of Korea, Luxembourg, Netherlands, New Zealand, Norway, Poland, Portugal, Spain, Sweden, Switzerland, United Kingdom, United States