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DO POLITICAL INSTITUTIONS INFLUENCE INTERNATIONAL TRADE?

MEASUREMENT OF INSTITUTIONS AND THE LONG-RUN

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Do Political Institutions Influence International Trade? Measurement of Institutions and the Long-Run Effects

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Abstract

The past literature presents ambiguous evidence about the bidirectional and causal influences between countries' institutional framework and their trading activity. In our analysis, we investigate the relationship between institutions and trade constructing a measure of institutions from the information given by the International Country Risk Guide and using a methodology that can control for omitted variables bias, endogeneity in the regressors, as well as cross-country heterogeneity. We examine the long-run effects of the political institutional framework on trade for a panel of 87 countries for the period from 1990 to 2007. We employ recent panel econometric methods for testing and estimating in the presence of non-stationarity, investigate panel causality and use methods that are robust to slope heterogeneity. Our results imply that an improved political institutional framework is a cause of increased trading activity.

Keywords: Political institutions, International trade, Panel co-integration, Cross-country heterogeneity

JEL-Code: F14, C10

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1. Introduction

The world has seen a great deal of economic and political transformation in the past. In Latin-America military dictatorships were disposed and a great wave of re-democratization took place in the 1980s, several new Eastern European countries were founded after the break-down of the Soviet-regime and market economy structures were introduced, and the upheaval in the Arab countries since December 2010 caused changes in institutional systems in (some of) these countries. However, reality teaches us that no clear association between the role that the political institutional framework (e.g. democratization) has played and the degree of globalization can be drawn: on the one hand, a higher degree of democratization fostered economic opening in the former Soviet republics, on the other hand, in cases like Bolivia or Peru just the opposite was true. China provides an illustrative example that economic and financial opening have not fostered democratization. In fact, China's successful economic development despite its adherence to an autocratic one-party-system seems to contradict previous academic arguments that democratic and good institutional structures are needed to generate economic development. Consequently, the link between institutions and globalization remains sort of a riddle and different studies in the former research literature find differential results.

Clarifying the direction and strength of the link between institutions and trade, however, remains an important task for both scientists as well as practitioners in politics. As several international organizations are supporting programs to improve countries' institutional structures, it is important to know how effective an improvement in institutional quality will be in order to enforce countries' trading activity.

When we speak of institutions, generally elements like the quality of contract enforcement, property rights, shareholder protection, investor protection or elements that refer to the political system like democracy, government stability, etc. are addressed (Levchenko 2007). However, there exists no clear single definition and the measurement of institutions varies across studies and depends on the specific context. In Levchenko's theoretical model good institutions lower transaction costs, they reduce uncertainty and provide a stable environment for productive interactions between economic agents and for allocating resources to their most efficient use.

In the literature, the effect of institutions on trade is also signed by great cross-country heterogeneity. Previous studies that used panel data in their analysis could at least control for some country-specific time-invariant effects due to the inclusion of country fixed effects. However, as changes in institutions might be correlated with other factors or correlate with socioeconomic conditions (Acemoglu et al. 2008, Brueckner and Ciccone 2011), simple panel estimation methodologies through capturing time and country fixed effects might be subject to omitted variables bias, still. Moreover, the past literature investigated not only the influence of institutions or democracy on trading activity but also the opposite link. It is fair to say that trade or globalization might have a direct impact on the institutional framework, as well. A two-sided relationship between democratization and trade is found for example by Eichengreen and Leblang (2008).

It is important to deal with non-stationarity issues in the analysis. As we will see in the following, clearly trading activity as well as the political institutional quality increased over time across countries, justifying analyses of non-stationarity. The recent literature has seen various applications of panel co-integration and causality methods which prove as a powerful tool to figure out the long-run relationship, for example between trade and income (Herzer 2013) or income inequality and growth (Herzer and Vollmer 2013). Panel co-integration estimators can cope with several severe econometric problems like omitted variables, cross-country heterogeneity and endogeneity of regressors (Pedroni 2007). A reduced system of a regression framework can be estimated since once we have found a co-integrating relationship, this relationship is robust against the inclusion of other factors that are supposed to exercise an influence on the dependent variable. In addition, in our analysis we will explicitly test for the bidirectional causality between institutions and trade.

In this paper, we analyze the relationship between political institutions and trade using data from the International Country Risk Guide and the World Development Indicators for a sample of 87 countries for the time period from 1990 to 2007. We find evidence for a significant, positive long-run relationship between political institutions and international trade and can also disentangle a one-sided causality that goes from institutions to trade.

The paper is organized as follows. Section 2 reviews the theoretical background and related literature. Section 3 introduces the empirical model, describes the data and methodological issues. Section 4 presents the empirical analysis and results. Section 5 concludes.

2. THEORETICAL BACKGROUND AND LITERATURE REVIEW

The economics literature is divided between studies that support a positive link between democracy and economic development, in particular economic growth (Acemoglu et al. 2014) and studies that support a non-significant or even negative link (Barro 1996, Gerring et al. 2005). Negative effects from democratization processes can occur if redistributions of income and land and interest group formation retard growth processes (Barro 1996). Positive effects can be expected to occur due to mutually reinforcing political and economic rights. Acemoglu et al. (2014), controlling especially for GDP dynamics, found that a country that switches from non-democracy to democracy achieves an about 20% higher GDP per capita in the next 30 years.

In the trade literature, studies found that institutions (Francois and Manchin 2013, Levchenko 2007, Anderson and Marcouiller 2002, De Groot et al. 2004, Ranjan and Lee 2003) or more specifically democracy (Yu 2010, 2007, Eichengreen and Leblang 2008) foster trading activity.[1] The literature addresses several reasons for this relationship. In the model of Yu (2010), democracy improves institutions. Better institutions involve stronger consumer rights, rule of law and property rights. This in turn will improve product quality and the reputation of a country's exports (Levchenko 2007), inducing decreasing trade costs. For an importer country, democratization influences trade costs via tariffs. The literature finds that democratization will lead to more liberal trade policies in less developed labor-intensive countries--where the political power is transferred from

[1] The economic literature has drawn attention to the individual countries' regime type for explaining trade instead of using information on congruence of the regime type for pairs of countries, a method which is usually applied in the political science literature (see e.g. Mansfield et al. 2000).

the elites to laborers, who benefit from pro-trade policies--, whereas in developed countries protectionism will be set up and fewer labor-intensive products are imported (O'Rourke and Taylor 2006, Milner and Kubota 2005).

Studying the relationship between institutions and trade is particularly interesting due to the two-sided effects that might exist. The literature's findings about the bidirectional relationship between institutions and trade remain rather inconclusive, however. Francois and Manchin (2013, page 167) argue that "institutional quality may also be driven by trade…", however, institutions "are more likely to have a more direct and immediate effect on the probability of trading and the amount traded than the other way around". The authors do not estimate the effect of trade on institutional quality, but rather use an instrumental variable strategy for estimating the one-sided effects on trade. Employing instrumental variables for trade, Eichengreen and Leblang (2008), Lopez-Cordova and Meissner (2005) and Rudra (2005) find a positive effect of trading activity for democratization, whereas Yu (2007), Li and Reuveny (2003) and Rigobon and Rodrik (2005) find a negative effect of trade liberalization on political liberalization. The literature provides explanations for both positive and negative effects of trade on institutional quality. On the one hand, free trade that results from a more efficient resource allocation will raise incomes, communication of ideas and therewith the demand for democracy (Lipset 1959, 1960). On the other hand, trade openness might sustain the status quo in a country (Yu 2010), because the land owners/ elites are the ones primarily receiving benefits from globalization (Acemoglu and Robinson 2006) and those have a keen interest for maintaining the current set-up of property rights and rule of law.

To the best of our knowledge, there exists one study so far in the literature that addresses non-stationarity issues and the long-run relationship between political institutions and trade. Using a panel data set of bilateral trade for 197 countries and the time period from 1976 to 2004, Nicolini and Paccagnini (2011) find no significant relationship between institutions (measured by the Freedom House index) and trade when running a causality test by Hurlin and Venet (2001) that controls for cross-sectional heterogeneity. This stands in contrast to previous studies that found a direct influence of the institutional quality on trade or a direct influence of trade on institutions. Nicolini and Paccagnini stress the fact that cross-country heterogeneity matters and it actually flaws results from conventional Granger causality tests which do not control for cross-country heterogeneity. This is a very important result, and as we will see in the following, our analysis confirms that cross-country heterogeneity is an important issue. Now, what always matters in empirical analyses, is which data set is chosen for the analysis, which countries and time span are covered, which variables are chosen and which methodology is used. Our study deviates from Nicolini and Paccagnini in three ways. First, we employ another data set and structure, namely monadic trade data for a more recent time period. We do thus not face the problem of the presence of zero trade flows that is apparent in bilateral trade data. When taking account of zero trade flows, actually PPML estimators would have to be used, however in the study of Nicolini and Paccagnini the Hurlin and Venet estimator controlling for cross-country heterogeneity could actually not be applied as the test is not specified for non-linear estimators. Second, we employ most recent dynamic panel econometric estimation methodologies capturing effects of cross-country heterogeneity. Lastly, we employ different measures for the institutional framework which are based on the International Country Risk Guide's political risk measure. We are not using measures like the Freedom House indices, as they have been subject to criticism in the past period concerning issues about systematical biases.

3. THE MODEL, DATA AND ECONOMETRIC METHODOLOGY

In order to investigate the long-run relationship between the political institutional framework and trade, we estimate the following bivariate model:

$$\ln(X_{it}) = \beta_{0i} + \beta_1 \text{ politicalinst}_{it} + \varepsilon_{it} \qquad (1)$$

where i=1,…,N denotes the cross-sectional unit (country), t=1,….,T denotes the time unit (year), and ε is the usual idiosyncratic error term. β_{0i} are country-specific fixed effects capturing country-specific omitted variables that are relatively stable over time. X is the trade variable, either exports or imports, measured as values of exports or imports in goods and services in constant 2000 US dollars. The trade data were taken from the World Development Indicators. The log of trade was taken such that a change in the explanatory variable would imply a percentage growth in trading activity. Our measure to capture the political institutional framework is *politicalinst*, which is a composite measure defined as the sum of the weighted components of the political risk measure of the International Country Risk Guide (ICRG). The index is based on the ICRG's rating of the following twelve components:

1. (12 points each) government stability, socioeconomic conditions, investment profile, internal conflict, external conflict
2. (6 points each) corruption, military in politics, religious tensions, law and order, ethnic tensions, democratic accountability, and
3. (4 points) bureaucratic quality.

Our measure for political institutions attains values between min 0 and max 100, a higher value representing a better political institutional framework. Our measure thus captures several aspects of institutions and is able to integrate several components that have been found to play an important role in the earlier literature, that is the political system, shareholder protection, property rights and contract enforcement, for example. Note that democracy (measured by the ICRG's variable "democratic accountability") remains rather stationary over time, instead, such that non-stationarity analyses cannot be run on the democracy variable (or other components) alone. For robustness checks we also computed an alternative measure of political institutions using principal component analysis based on subcomponents' covariance matrix. The index consists of the components law, bureaucracy and democratic accountability and its quality was assessed by a scree plot, see Table 1 and Figure 1 and the summary statistics of relevant variables in the Appendix.

Table 1

Principal Component Analysis generating an alternative measure of political institutions

Variable	Component 1
Law	0.5888
Bureaucracy	0.4675
Democratic accountability	0.6594

Note: This Table shows the component's factor loadings for the political risk measure's components law, bureaucracy and democratic accountability.

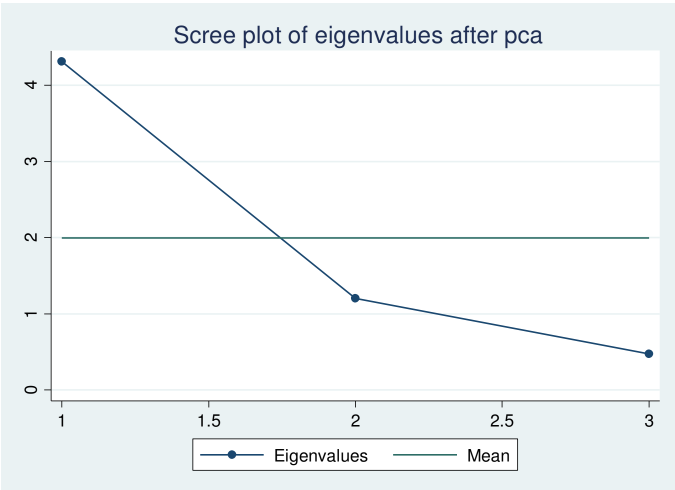

FIGURE 1

Scree plot of eigenvalues after principal component analysis

The political institutions index is a time series that is generally bounded, namely between 0 and 100. Recent advances in econometric methodology found that traditional unit root tests that are used in a context of bounded variables tend to over-reject the null hypothesis of a unit root (would indicate stationarity more often). Now, for our study this appears not to be a problem, as clearly the tests that we run indicate non-stationary of the variables.[2] Our data cover the period from 1990 to 2007 and we take up all countries into the sample for which data are available, leading to a total of 87 countries (see the list in the Appendix). Our data coverage was restricted by the fact that co-integration estimators need a balanced panel structure. Since we will exploit both the time and the cross-section dimension of our data when using panel co-integration estimation methods, the number of observations is sufficiently high to identify and estimate co-integration relationships. The summary statistics for our variables are displayed in the Appendix.

In our analysis, we will first test for the non-stationarity of the variables. The idea is to disentangle economic long-run relationships between variables that bear a stochastic trend over time and to differentiate these relationships from spurious regression results. In fact, the tendency to find a spurious relationship, only, is even higher in the case of panel data analysis than for pure time series analyses (Entorf 1997). The requirement for a relationship between non-stationary variables not to be spurious is that the two non-stationary variables co-integrate. If two non-stationary variables are found to be co-integrated, a long-run equilibrium relationship between these variables exists. Finding a co-integration relationship involves that no other important non-stationary variable has been omitted from the regression, otherwise no co-integration would be detected because omitted relevant non-stationary variables would enter the error term and produce non-stationary residuals (Everaert 2011). Moreover, the co-integration estimates are not biased by omitted stationary variables (Bonham and Cohen 2001), as the coefficients are super-consistent in the presence of temporal or contemporaneous correction between the error term and the regressors (Stock 1987). This justifies a reduced form model. Once a co-integration relationship is found, one can safely argue that it is invariant to the inclusion of further regressors (Lütkepohl 2007), the estimates will be unbiased. Furthermore, no endogeneity problems arise, because the co-integrating estimator is super-consistent (Engle and Granger 1987, Stock 1987, Herzer 2013), that is the parameter estimates are not only consistent, but converge to the true parameter values more quickly, with a rate T, than with a rate $\sqrt{T}$. Consequently, we will get more accurate estimates when using cointegration estimates than when using traditional methods. In a last step of our analysis, the direction of long-run causality will be investigated in order to figure out if an improved political institutional framework causes increased trading activity or if the former is an effect of the latter. Moreover, we will consider recent advances in econometric methodology to cope with cross-country heterogeneity.

4. Empirical Analysis

[2] Cavaliere and Xu (2014) further explain that conventional unit root tests behave according to standard asymptotic theory when the bounds are sufficiently far away from the maximal and minimal observations of a variable, which is also the case in our study.

4.1 PANEL UNIT ROOT AND CO-INTEGRATION TESTS

We use the panel unit root tests of Breitung (2000) and Pesaran (2007) to check for the presence of non-stationarity for the variables under study. A more detailed description of the methodology can be found in the Appendix. The unit-root test from Breitung belongs to the first-generation panel unit root tests and was found to have the highest power and smallest size distortions (Breitung 2000). However, a drawback of the Breitung test is that it assumes cross-sectional independence. As we can assume that institutional frameworks and trading activity vary considerably across countries, we also run estimations using the unit root test from Pesaran which is able to capture heterogeneity across countries. The test statistics as shown in Table 2 imply that the null hypothesis of a unit root cannot be rejected. The results from tests for the first differences display that the variables are all integrated of order 1 (non-stationarity).

TABLE 2

Panel unit root tests

Variables	Breitung (2000)	Pesaran (2007)
In levels		
ln(exports)	-0.32436	6.36
ln(imports)	2.0281	2.784
politicalinst	0.16042	0.604
First differences		
ln(exports)	-14.5139**	-5.537**
ln(imports)	-12.211**	-6.445**
politicalinst	-13.7199**	-6.916**

Note: For the tests maximally two lags were included to control for possible autocorrelation based on the Schwarz information criterion. Individual-specific intercepts and time trends were included in the regressions. ** indicates significance at the 5% level, * indicates significance at a 10% level.

In a next step, we investigated if a long-run equilibrium relationship (co-integration) between institutions and trade exists. We used the approach of Pedroni (2004), taking the panel and group Augmented Dickey Fuller (ADF) and Phillips-Perron (PP) tests. All tests reject the null hypothesis of no co-integration, implying that there exists a long-run relationship between the political institutional framework and trade. To control for cross-sectional dependence we also employed the cross-sectionally augmented IPS (CIPS) test (see the Appendix for a more detailed explanation). Again the results point to the existence of a co-integration relationship between institutions and trade.

TABLE 3

Panel co-integration tests

	Ln exports	Ln imports
Panel PP	-4.256131**	-5.963085**
Panel ADF	-6.892416**	-10.50642**
Group PP	-3.628639**	-5.044449**
Group ADF	-8.27847**	-9.648294**
CIPS	-10.9241**	-9.7389**

Note: Maximally two lags were included to control for possible autocorrelation based on the Schwarz information criterion. A deterministic trend and an intercept were included in the regressions. ** indicates significance at the 5% level, * indicates significance at a 10% level.

4.2 RESULTS FOR THE LONG-RUN RELATIONSHIP

We estimated the long-run effect of the political institutional framework on trade using dynamic ordinary least squares (DOLS) and fully modified ordinary least squares (FMOLS) estimators. These estimators are asymptotically efficient. In comparison, however, the DOLS-estimator outperforms the FMOLS-estimator (Kao and Chiang 2000). The estimators yield unbiased estimates for co-integrating variables, even under the presence of endogenous regressors. Both the DOLS and FMOLS estimators are super-consistent under co-integration and also robust to the omission of variables that are not part of the co-integration relationship (Herzer 2013).

The pooled panel DOLS estimator (Mark and Sul 2003) is given by:

$$\ln(X_{it}) = \beta_{0i} + \beta_1 \ \ politicalinst_{it} + \sum_{j=-k}^{k} \mathrm{B} \ \Delta \ politicalinst_{\ it-j} + \varepsilon_{it} \qquad (2)$$

where k lead and lag differences as well as the current difference of politicalinst are included in the regressions, accounting for possible serial correlation and endogeneity of the regressors. As such, this approach does not require the use of instrumental variables or exogeneity assumptions. So called group mean estimators (Pedroni 2001) allow the coefficients to vary across countries, whereas the FMOLS estimator adds a semi-parametric correction in addition to the OLS estimator to get rid of the bias due to endogeneity of regressors.

All the results in Table 4 imply that there exists a significant, positive long-run relationship between political institutions, imports and exports. The results appear to be robust as different estimators like the pooled or the mean group ones yield similar results.

TABLE 4

Long-run effects estimators

	DOLS pooled	DOLS group mean	FMOLS pooled	FMOLS group mean	CCE mean group
Exports equation	0.022704**	0.030563**	0.029897**	0.031785**	0.0043403**
Observations	1296	1296	1479	1479	1566
Imports equation	0.024497**	0.031539**	0.028518**	0.033314**	0.0092448**
Observations	1319	1296	1479	1479	1566

Note: DOLS estimation was run with max. two leads and two lags based on the Schwarz information criterion. ** denotes significance at the 5% level, * denotes significance at a 10% level.

Table 5

Cross-sectional dependence tests

CD test statistic for residuals from:	DOLS pooled	DOLS group mean	FMOLS pooled	FMOLS group mean	CCE mean group
Exports equation	94.03**	66.16**	145.59**	141.15**	0.12
Imports equation	97.26**	72.25**	143.09**	152.55**	1.17

Note: ** indicates significance at the 5% level, * denotes significance at the 10% level.

However, assuming that cross-country heterogeneity plays a role, we checked for the relevance of cross-sectional dependence for the estimators by employing the Pesaran (2004) test (see Eberhardt and Teal (2011) for a critical discussion). Cross-sectional dependence can be due to unobserved common factors. Not properly controlling for this issue might result in biased and inconsistent estimates of the coefficients and standard errors (Driscoll and Kraay 1998) and might obscure the causal relationship as has been shown by Nicolini and Paccagnini (2011). For every equation and version of the DOLS and FMOLS estimator the test indicated cross-sectional dependence, such that we favored to employ the CCE mean group estimator by Pesaran (2006) for further analyses. For this estimator cross-sectional averages of the dependent variable and the observed regressors are included as proxies for unobserved factors. This estimator has been shown to be consistent even when there is weak cross-sectional dependence of error terms or even correlation with the covariate (Pesaran 2006). The CD statistic is normally distributed under the null hypothesis of no cross-sectional dependence. The estimates as shown in Table 3 indicate a significant long-run relationship between institutions and trade and inference is not biased by cross-sectional dependence, as can be seen from the CD test results as shown in Table 5.

4.3 Panel Causality Tests

It might well be that the positive coefficient obtained for the relationship between political institutions and trade is not resulting from an impact of institutions on trade but from an impact of trade on institutions. This would justify research findings by Lopez-Cordova and Meissner (2005) and Yu (2005) who found that globalization fosters democratization. Consequently, we will also have to investigate the direction of causality. Finding a co-integration relationship implies long-run causality in at least one direction (Granger 1988).

We estimate the following panel vector error correction model:

$$\Delta \mathrm{X}_{\mathrm{it}} = \beta_{0,1} + \beta_{1,1i} Z_{it-1} + \sum_{j=1}^{k} (\gamma_{1,1\mathrm{i}} \Delta \mathrm{X}_{\mathrm{it-j}} + \gamma_{2,1\mathrm{i}} \Delta \mathrm{politicalinst}_{\mathrm{it-j}}) + \varepsilon_{1it} \quad (3)$$

$$\Delta \mathrm{politicalinst}_{\mathrm{it}} = \beta_{0,2} + \beta_{1,2i} Z_{it-1} + \sum_{j=1}^{k} (\gamma_{1,2\mathrm{i}} \Delta \mathrm{X}_{\mathrm{it-j}} + \gamma_{2,2\mathrm{i}} \Delta \mathrm{politicalinst}_{\mathrm{it-j}}) + \varepsilon_{2it} \quad (4)$$

where Z are the residuals of individual CCE mean group long-run estimations and the β_0 are fixed effects. The equations are estimated by CCE mean group methodology (see Sakyi et al. 2015). For a long-run relationship to hold, at least one of the adjustment coefficients must be different from zero, see the Granger representation theorem (Engle and Granger 1987). A significant adjustment coefficient β_1 indicates long-run Granger causality from the independent to the dependent variable (Granger 1988) and shows how the trade variable and the institutions variable respond to deviations from the long-run equilibrium. In the following, we test whether the coefficients β_1 are statistically different from zero, applying a likelihood ratio test that is distributed as χ^2.

The results in Table 6 show that in the long-run trade is caused by the political institutional framework but not vice versa.[3] The error-correction terms are significant and display long-run Granger causality and consequently long-run endogeneity (Hall and Milne 1994). The likelihood ratio test is rejected in case of the relationship from institutions to trade. The results appear not to be affected by cross-sectional dependence.

[3] This complies with results from Acemoglu et al. (2008) in the case of income and democracy who found that there is no causal effect of income on democracy, once they control for country fixed effects. Acemoglu et al. state that the effect of income for democracy can be found only for a longer horizon of time, about 500 years.

TABLE 6

Tests for causality

	Exports equation	Imports equation
Trade variable influenced by political institutions	-0.94587**	-1.17213**
$\chi^2(1)$	33.6	80.92
p-value	0.000	0.000
CD Stat	-0.74	0.4
Political institutions influenced by trade variable	3.1599	-1.17635
$\chi^2(1)$	0.42	0.05
p-value	0.5187	0.8272
CD-Stat	-0.22	-0.22

Note: This Table shows the error correction coefficients and denotes their significance. A Vector Error Correction Model specification was run. Two lags were included. The error term and the equation were estimated by the CCE mean group procedure. The results from a likelihood ratio test are shown as well as results from a test for cross-sectional dependence. ** indicates significance at a 5% level, * indicates significance at a 10% level.

4.4 ROBUSTNESS CHECKS

To check the robustness of our results, we already used different estimation methods as shown in Table 4, and will further investigate sample-selection bias and results that emerge from a different measurement of the explanatory variable. In order to control for sample-selection bias, we run regressions for those countries with higher or lower values of the political institutions variable than the countries' average over the period from 1990 to 2007. The results as shown in Table 7 reveal that the long-run relationship between political institutions and trade remains positive and significant for both the high or low institutions sample. We can see that cross-sectional dependence still remains a problem in these estimations.

Results from using a different measure for institutions again reveal that there exists a significant, positive long-run relationship between institutions and trade. The coefficients are higher than for the former composite measure of institutions. Our conclusion is that measurement of institutions and the choice of subcomponents to explain the institutional framework matter.

TABLE 7

Sub-sample estimates and different measure for institutions

	Political institutions values above country average	Political institutions values below country average	Alternative measure for political institutions
Exports equation	0.0034**	0.0024**	0.0461**
Observations	702	864	1566
CD Stat	-0.92	-2.3**	1.52
Imports equation	0.0095**	0.0066**	0.0325**
Observations	702	864	1566
CD Stat	1.72**	-2.54**	-0.39

Note: CCE mean group estimation was run. ** denotes significance at the 5% level, * denotes significance at a 10% level.

5. CONCLUSIONS

In this paper we examined the long-run relationship between the political institutional framework and trade using panel unit root and co-integration techniques for a panel of 87 countries for the time period from 1990 to 2007 making use of information on the institutional framework from the International Country Risk Guide. With the econometric methodology we were able to control for omitted variables, cross-country heterogeneity and endogeneity bias and could employ a reduced form model for countries' trading activity and the quality of political institutions. From our results, we can conclude that the political institutional framework has a positive long-run effect on trading activity. This relationship is robust to different estimation methods. In addition, our results confirmed a one-sided long-run causality from institutions to trade. An improved political institutional framework is a cause of increased trading activity.

These results bear important policy implications in terms of the effectiveness of international organizations' support for countries to improve their institutional framework. As our results show, improving institutions has a significant positive causal effect for trading activity.

Our analysis showed that cross-country heterogeneity is an important issue that should not be neglected in studies. Moreover, there are elements in the institutional framework that can be supposed bear stronger influence on trade or economic development than other ones. A more detailed analysis of individual components of institutions by using panel co-integration methodology, however, was prevented due to the lack of non-stationarity of variables.

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APPENDIX

Appendix A

Panel unit root and co-integration methodology

We employ several unit root tests designed especially for the use of panel data in order to test for non-stationarity of variables. The Breitung (2000) test belongs to the first-generation panel unit root tests which do not consider cross-sectional dependence issues. It is based on a transformed augmented Dickey Fuller (ADF) regression and the test statistic is:

$$B = (\frac{\hat{\sigma}^2}{nT^2}\sum_{i=1}^{n}\sum_{t=2}^{T-1}(y_{it-1}^{*})^2)^{-1/2}\frac{1}{\sqrt{n}T}\sum_{i=1}^{n}\sum_{t=2}^{T-1}(\Delta y_{it})^{*}y_{it-1}^{*}$$

where

$$(\Delta y_{it})^{*} = s_t(\Delta y_{it} - \frac{1}{T-t}(\Delta y_{it+1} + \cdots + \Delta y_{iT}))$$

and

$$y_{it-1}^{*} = y_{it-1} - y_{i0} - \frac{t-1}{T}(y_{iT} - y_{i0}).$$

An ADF regression of the ith cross-sectional unit is:

$$\Delta y_{it} = \alpha z_{it} + \beta_i y_{it-1} + \sum_{l=1}^{L} \gamma_{il}\Delta y_{it-l} + \varepsilon_{it}$$

where L is the order of lags and alpha denotes deterministic terms. The hypothesis that is tested is:

$H_0: \beta_i = 0$ for all i, which is the test for the existence of a unit root if H_0 is not rejected.

The Pesaran (2007) test, called CIPS test, is a cross-sectionally augmented IPS test. In the ADF regression cross-sectional averages of lagged levels and first differences are added:

$$\Delta y_{it} = \alpha z_{it} + \beta_i y_{it-1} + \sum_{l=1}^{L} \gamma_{il}\Delta y_{it-l} + \delta_i \bar{y}_{t-1} + \sum_{l=0}^{L}\eta_{il}\Delta\bar{y}_{t-l} + \varepsilon_{it}$$

with $\bar{y}_t$ as the cross-sectional mean. The test statistic is the average of individual IPS statistics:

$CIPS = \frac{1}{N}\sum_{i=1}^{N} t_i.$

In order to test for co-integration, we run the panel ADF and Philips-Perron (PP) and group ADF and PP statistics as proposed by Pedroni (1999, 2004). To consider cross-sectional dependence, we use the methodology as proposed by Holly et al. (2010). In the first step, we used the CCE procedure by Pesaran (2006) which allows for cross-sectional dependencies. The co-integrating regression for the ith cross-section is:

$$\ln(X_{it}) = \alpha_i + \beta_i \ln(politinst_{it}) + \gamma_i \overline{\ln(X_t)} + \delta_i \overline{\ln(politinst_t)} + \varepsilon_{it}$$

where unobserved factors are represented by the cross-section averages. In a second step cross-sectionally augmented IPS statistic for residuals from the CCE regressions were computed. This way unobserved common factors can be controlled for. A co-integration relationship exists when the hypothesis of a unit root in the residual regression can be rejected.

Appendix B

List of countries

Full sample

Algeria, Argentina, Australia, Austria, Bahamas, Bangladesh, Belgium, Bolivia, Botswana, Brazil, Brunei, Bulgaria, Cameroon, Canada, Chile, China, Colombia, Congo DR, Costa Rica, Cote d'Ivoire, Cuba, Cyprus, Denmark, Dominican Republic, Ecuador, Egypt, El Salvador, Ethiopia, Finland, France, Gabon, Gambia, Greece, Guatemala, Guinea, Honduras, Hungary, Iceland, India, Indonesia, Iran, Ireland, Italy, Japan, Jordan, Kenya, Luxembourg, Madagascar, Malaysia, Mali, Malta, Mexico, Morocco, Mozambique, Namibia, Netherlands, New Zealand, Nicaragua, Norway, Pakistan, Panama, Paraguay, Peru, Philippines, Poland, Portugal, Romania, Senegal, South Africa, South Korea, Spain, Sudan, Sweden, Switzerland, Syria, Tanzania, Thailand, Tunisia, Turkey, Uganda, United Kingdom, United States, Uruguay, Venezuela, Vietnam, Zambia, Zimbabwe

Low income and middle-low income countries

Bangladesh, Bolivia, Cameroon, Congo DR, Cote d'Ivoire, Egypt, El Salvador, Ethiopia, Gambia, Guatemala, Guinea, Honduras, India, Indonesia, Kenya, Madagascar, Mali, Morocco, Mozambique, Nicaragua, Pakistan, Philippines, Senegal, Sudan, Syria, Tanzania, Uganda, Vietnam, Zambia, Zimbabwe

High income and middle-high income countries

Algeria, Argentina, Australia, Austria, Bahamas, Belgium, Botswana, Brazil, Brunei, Bulgaria, Canada, Chile, China, Colombia, Costa Rica, Cuba, Cyprus, Denmark, Dominican Republic, Ecuador, Finland, France, Gabon, Greece, Hungary, Iceland, Iran, Ireland, Italy, Japan, Jordan, Luxembourg, Malaysia, Malta, Mexico, Namibia, Netherlands, New Zealand, Norway, Panama, Paraguay, Peru, Poland, Portugal, Romania, South Africa, South Korea, Spain, Sweden, Switzerland, Thailand, Tunisia, Turkey, United Kingdom, United States, Uruguay Venezuela

Countries with political institutions values above average

Argentina, Australia, Austria, Bahamas, Belgium, Botswana, Brunei, Bulgaria, Canada, Chile, Costa Rica, Cyprus, Denmark, Finland, France, Greece, Hungary, Iceland, Ireland, Italy, Japan, Luxembourg, Malaysia, Malta, Mexico, Namibia, Netherlands, New Zealand, Norway, Poland, Portugal, South Korea, Spain, Sweden, Switzerland, Tunisia, United Kingdom, United States, Uruguay

Countries with political institutions values below average

Algeria, Bangladesh, Bolivia, Brazil, Cameroon, China, Colombia, Congo DR, Cote d'Ivoire, Cuba, Dominican Republic, Ecuador, Egypt, El Salvador, Ethiopia, Gabon, Gambia, Guatemala, Guinea, Honduras, India, Indonesia, Iran, Jordan, Kenya, Madagascar, Mali, Morocco, Mozambique, Nicaragua, Pakistan, Panama, Paraguay, Peru, Philippines, Romania, Senegal, South Africa, Sudan, Syria, Tanzania, Thailand, Turkey, Uganda, Venezuela, Vietnam, Zambia, Zimbabwe

Appendix C

Table 8

Descriptive Statistics

Variables	Obs	Mean	Std. Dev.	Min	Max
Lnexports, overall	N=1566	23.23018	1.935405	18.75431	27.9887
between	n=87		1.904293	18.97259	27.53159
within	T=18		0.3985644	21.731	24.84474
Lnimports, overall	N=1566	23.30653	1.832426	19.02189	28.31792
between	n=87		1.800082	19.23561	27.7569
within	T=18		0.390752	21.86527	24.70418
Political institutions, overall	N=1566	68.07757	14.08944	14.24	96.08
between	n=87		12.71869	33.36611	92.38222
within	T=18		6.205188	34.69923	83.64368
Alternative political institutions measure, overall	N=1566	6.202585	2.076707	1.1269	9.3592
between	n=87		1.921796	1.719662	9.3592
within	T=18		0.8121157	2.538244	8.874432