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Globalization and Corruption, revisited

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Abstract

This paper presents new empirical evidence on the determinants of corruption, focussing on the role of globalization and inequality. The estimates for a panel of 102 countries over the period 1995-2005 point to three main results: i) Detection technologies, reflected in a high level of development, human capital, and political rights reduce corruption, whereas natural resource rents increase corruption. ii) Globalization (in terms of both trade and financial openness) has a negative effect on corruption, which is more pronounced in developing countries. iii) Inequality increases corruption, and once the role of inequality is accounted for, the impact of globalization on corruption is halved. In line with recent theory, this suggests that globalization – besides reducing corruption through enhanced competition – affects corruption also by reducing inequality.

Keywords: Globalization, inequality, corruption
JEL Codes: F1, F3, F4, O1

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"*It is difficult to overstate the economic and social significance of corruption.*"
(Ades and Di Tella, 1999, p.982)

I Introduction

Both the determinants of corruption (rent-seeking behavior) and its welfare consequences have been subject to extensive theoretical and empirical research (e.g., Krueger, 1974; Murphy et al., 1993; Mauro, 1995; Bliss and Di Tella, 1997; Das and DiRienzo, 2009; Bhattacharyya and Hodler, 2010). It is widely agreed in the literature that significant economic rents (in particular through natural resource exports) increase corruption, whereas 'detection technologies', reflected in a high level of economic development, good institutions, competitive elections, political rights and press freedom, as well as an educated population decrease corruption.

At a general level, Svensson (2005) describes corruption as the outcome of a country's underlying legal, economic, cultural and political institutions. These factors – which are widely perceived as determinants of democracy and institutional quality – directly and indirectly affect the level of corruption. As a consequence, the analysis of the determinants of corruption is closely related to the literature on democracy and institutional quality.

Previous studies have focussed on the role of globalization as a determinant of institutional quality and corruption. One strand of this literature emphasizes the pro-competitive effect of trade. Ades and Di Tella (1999) find that an increase in trade openness leads to a decline in corruption and argue that import competition reduces the rents for domestic firms and thus the rewards from corrupt activities. A related argument can be made with respect to financial openness, since increased FDI and market entry of foreign firms are typically conducive to domestic competition. There is also evidence for a two-way relationship between competition (trade) and corruption: Emerson (2006) shows that corruption significantly decreases the number of firms and thus the level of competition within a country.

Recent theoretical work by Acemoglu and Robinson (2005) highlights the role of inequality as another channel through which globalization (in terms of both trade and financial openness) affects institutional quality. Building on the Heckscher-Ohlin model, they argue that when a labor abundant developing country opens up to trade or capital inflows, inequality is reduced since wages will increase relative to the return to capital. With a lower level of inequality, democracy becomes less redistributive and thus more likely. According to this reasoning, globalization is expected to affect democracy (and thereby also corruption) in developing countries both via trade and financial openness and mainly through its effect on inequality.

While the detrimental role of corruption for economic development is undisputed and there is also strong evidence on the link between trade and corruption, the empirical literature on the role of financial openness is sparse. One notable exception, using democracy rather than corruption as the dependent variable, is Eichengreen and Leblang (2008), who find evidence for a positive effect of financial openness. This lack of studies is surprising

in light of the surge in financial openness over the last two decades that has outpaced the increase in trade openness (Lane and Milesi-Ferretti (2007)). Moreover, there is hardly any evidence on the transmission channels through which globalization affects corruption. In particular, to the best of our knowledge, there is no empirical study that has focussed on the interplay of globalization (in terms of both trade and financial openness) and inequality in reducing corruption.

Against this background, the present paper makes the following contributions to the empirical literature on the globalization-corruption nexus. First, we use an encompassing model that relates corruption to globalization in terms of both trade and financial openness, thereby paying particular attention to the potential endogeneity of our globalization measures. Second, we explicitly consider the role of inequality as a determinant of corruption and its interplay with globalization, allowing us to provide some evidence on the transmission channels through which globalization affects corruption. Third, unlike previous studies that are mainly based on cross-sections of countries, we use a (unbalanced) panel of (up to) 102 countries over the period 1995-2005, allowing us to substantially increase the number of observations, exploit the time-variation in the data, and to control for region-specific effects.

In line with recent studies we find that a higher level of development, more education and political rights, and a lack of natural resource rents are negatively associated with corruption. Globalization, in terms of both trade and financial openness, appears to play an important role in reducing corruption, particularly in developing countries. Accounting for the effect of inequality, which itself increases corruption, the effect of globalization is halved, which provides empirical support for the theoretical argument by Acemoglu and Robinson (2005) that globalization affects democracy (and thereby, institutional quality and corruption) through its negative effect on inequality.

The remainder of this paper is organized as follows: Section II gives a brief overview of the literature on globalization, inequality, and corruption. Section III sets up the empirical model, describes the data used, and outlines our identification strategy. Section IV presents the estimation results. The final section V summarizes the results and concludes.

II Globalization, Inequality, and Corruption: A Brief Survey of the Literature

The literature on the links and interactions between globalization, institutional quality, and corruption is extensive. In the following we provide a very brief survey with a focus on the effects of trade and financial openness, the role of inequality, and the link between democracy and corruption. For a more comprehensive survey of the literature on globalization and institutional quality, see Eichengreen and Leblang (2008); for a detailed review of the globalization-inequality nexus, see Goldberg and Pavcnik (2007).

1. Globalization, Competition, and Corruption

Ades and Di Tella (1999) suggest that an increase in trade openness leads to a decline in corruption. Since firms can reap some of the rents by exchanging control rights for bribes, there will be a positive relationship between corruption and rents. In their empirical analysis, using a cross-section of 52 (31) countries over the period 1980-1983 (1989-1990) for the Business International Corporation's corruption survey (World Competitiveness Report corruption index), they relate corruption to competition (proxied by the share of imports in GDP and the distance to world's major exporters) and the share of fuel and mineral exports in total exports. They find that imports and proximity to world's major exporters significantly reduce corruption, whereas fuel and mineral exports increase corruption. Ades and Di Tella (1999) interpret this result as strong evidence that competition from foreign firms leads to a decrease in the rents of domestic firms and consequently to a decline in corruption. Leite and Weidmann (1999) also find a negative effect of trade openness on corruption using a cross-section of 72 countries and the ICRG corruption index in 1995. A positive effect of natural resource rents on corruption is also identified in the recent study by Bhattacharyya and Hodler (2010) for a sample of 124 countries over the period 1980-2004.

Pro-competitive effects, reducing the scope for corruption due to a decrease in bureaucratic power, do not only stem from international trade. Increased competition might also be a consequence of domestic market deregulation and policies facilitating firms' engagement in international activity (Svensson, 2005). Hence, trade might capture only part of the effects of competition on corruption. Das and DiRienzo (2009) cast doubt on the linear effect of globalization on corruption. Using a cross-section of 113 countries for the period 2008 and the KOF index as a broader measure of globalization (which includes economic, political and social integration and Transparency International's corruption perception index), they find that only countries with moderate levels of globalization can capture the positive effects of further integration. At lower levels of globalization, an increased participation in the global market may even increase the scope for corruption due to a poor regulative environment. Lalountas, Manolas and Vavouras (2011), voicing similar critique in an analysis for a cross-section of 127 countries in 2006 and using the same set of variables as Das and DiRienzo (2009), find that globalization significantly reduces corruption in middle- and high income countries, once nonlinearities and the endogeneity of GDP per capita are accounted for, while globalization has no effect in low-income countries.

2. Globalization, Inequality, and Democracy

Corruption can also be interpreted as a more general indicator of institutional quality, reflecting a country's underlying legal, economic, cultural and political institutions (Svensson, 2005). In the following, we review some of the literature on the relationship between globalization and democracy, which we consider relevant in identifying potential determinants of corruption. At the heart of our empirical analysis is the recent theoretical argument by Acemoglu and Robinson (2005) emphasizing the role of inequality. They

show that the emergence and survival of democracy depend on the distribution of income and thus on factor prices. Thus globalization in terms of international trade and financial openness will affect democracy and institutional quality through its effect on factor prices. In a Heckscher-Ohlin framework with labor abundant (and capital scarce) developing countries as a representative case, opening up to trade leads to an increase in the wage-rental ratio (Stolper-Samuelson and Heckscher-Ohlin theorem). This reduces the income gap between factors and thereby also the risk of political conflict, since voters of lower income groups have less demand for highly redistributive policies. As a consequence, democracy becomes less threatening to upper income groups and the elite. A similar argument applies to financial openness, which increases the elasticity of capital supply, leading to an inflow of capital in labor abundant countries. This lowers the relative return to capital, thereby reducing the income gap between factor earnings, and supports democratic development.

Of course, if labor is scarce and capital abundant, these effects are reversed.[1] If the developing country is land abundant, opening up to trade increases the income of land owners who typically represent a small elite, inequality increases and democracy becomes less likely. As a consequence, the bottom line effect of globalization on democracy (and corruption) remains an empirical question (Acemoglu and Robinson, 2005, p.282).

López-Córdova and Meissner (2008), analyzing the determinants of democracy, provide support for the close relation between international trade, natural resources and institutional outcomes based on a sample of three time periods spanning 130 years from 1870-1919, 1917-1939 and 1960-2000. Using the same data set, empirical evidence on the positive effect of financial integration on democracy is given by Eichengreen and Leblang (2008). However, none of these studies have considered trade and financial openness simultaneously or explicitly taken into account the role of inequality as a transmission channel.

In a recent paper Levchenko (2011) demonstrates that the consequences of globalization for institutional quality may vary across pairs of countries. He considers the role of institutions as a determinant of comparative advantage in trade. Institutions generate rents for some parties within the economy. When two countries with similar technology open up to trade, institutional quality will improve in both countries as rents disappear in the country with inferior institutions. However, if technological differences between trading partners are big enough, specialization patterns will not change and trade might not reduce rents and hence provide no incentive for institutional improvement. Using a sample of 139 countries for a cross-section over the period 1996-2000 Levchenko (2011) finds supportive evidence that trade in institutionally intensive goods improves institutional quality.

[1] However, regarding the role of financial openness, Acemoglu and Robinson (2005) argue that a capital flight might lead to a cut in capital taxes (as predicted by median voter models), which might stop and redirect capital flows, ultimately leading to lower capital returns and more democracy.

3. Democracy and Corruption

As evident from this brief survey, the links and interactions between globalization, inequality, institutional quality and corruption are manyfold and intertwined. This paper does not aim at incorporating all these interrelationships into the empirical analysis but rather puts its emphasis on some aspects and relationships that have not been considered in the literature so far, namely the role of financial openness besides trade openness, the role of inequality, and its interplay with globalization.

Since the theoretical motivation for our empirical analysis partly relies on the effect of globalization on democracy, some words on the link between democracy and corruption are in order here. Most theoretical models assume a negative relationship, with more democracy reducing corruption. In that case, a positive effect of globalization on democracy as hypothesized in Acemoglu and Robinson (2005) would translate directly into a negative effect on corruption. However, it should be added that this clear cut relationship between democracy and corruption is not undisputed and the empirical evidence is mixed. Rock (2009), for example, argues that there is an inverted U-shaped relationship between democracy and corruption. However, the turning point occurs rather early in the life of new democracies (after some 10 years). Hence, in a sample where recently established democracies are the exception rather than the rule, we would expect the negative effect of democracy on corruption to dominate, such that the theoretical rationale underlying the globalization-democracy nexus adds to the motivation of the empirical test of the effect of globalization on corruption.

III Empirical Model and Data

1. Basic Model Specification

The starting point for our baseline empirical model is the specification in Ades and Di Tella (1999), who relate corruption (CI) to trade openness (TO) and a set of control variables, namely GDP per capita (Y/L), human capital (HC), political rights (PR), and resource rents (RR). Our specification departs from theirs in three respects: First, we use a panel rather than a cross-section of countries, which allows us to increase the number of observations and to control for region-specific fixed effects. While most of the variation in our dependent variable is across countries rather than over time, using a panel (with year specific fixed effects) has the further advantage to avoid 'snapshot-results' specific to a single time period.[2]

Second, following the theoretical reasoning in section II, we include financial openness (FO) besides trade openness (TO) to obtain a more comprehensive measure of globalization and to asses the respective roles of its two major dimensions. Third, we will consider

[2]The issue of the sensitivity of cross-section regressions with respect to the choice of the reference year has been prominently raised by Baier and Bergstrand (2007). Estimating the same cross-sectional gravity model on the trade effects of free trade agreements (FTAs) for alternative time periods, they find that "the FTA dummy's coefficient estimates are highly unstable from year to year" (Baier and Bergstrand, 2007, p.76).

specifications including a measure of income inequality (IQ), which allows us to provide an assessment of its role as a determinant of corruption, its interplay with globalization, and the transmission channels through which globalization affects corruption. Hence, our most comprehensive empirical model reads:

$$CI_{i,t} = \beta_0 + \beta_1 \ln(Y/L)_{i,t} + \beta_2 HC_{i,t} + \beta_3 PR_{i,t} + \beta_4 RR_{i,t} + \beta_5 TO_{i,t} + \beta_6 FO_{i,t} + \beta_7 IQ_{i,t} + \eta_r + \nu_t + u_{i,t}. \tag{1}$$

The dependent variable CI is a country's score on the corruption perception index from Transparency International, which defines corruption as "the abuse of entrusted power for private gain" in the public and private sector with a special focus on politicians, civil servants and public officials.[3] The index ranges from zero to ten, where low values correspond to high levels of corruption. For the ease of interpretation we rescale the index, taking the difference to the maximum possible value (10), such that higher values of CI are associated with higher levels of corruption.

GDP per capita, human capital, and political rights are used as overall indicators of economic development and as a proxy for 'detection technologies', which reduce rent seeking behavior (Ades and Di Tella, 1999). Real GDP per capita (Y/L) in 2005 international dollars is from the Penn World Tables 7.0. Human capital (HC), measured as mean years of schooling for the working age population, is from the Barro and Lee (2011) data set. As an indicator of political rights and civil liberty, we use the Freedom House index, which summarizes the extent of competitive elections, political rights, press freedom, freedom of expression and belief, associational and organizational rights, rule of law, and personal autonomy and individual rights. It is defined over a range from 0 to 7 with higher values being associated with more political freedom and civil liberties. Data on the variable PR is taken from Norris (2009).

Rents generated by natural resources (RR) are included as a measure of exogenously created rents for domestic firms. As with other rents, we expect that resource rents increase corruption (Leite and Weidmann, 1999; Bhattacharyya and Hodler, 2010). Data is taken from the the World Bank's WDI database.

Trade openness (TO) is defined as ratio of imports plus exports to GDP (at current prices) and taken from the Penn World Tables 7.0, financial openness (FO) is defined as ratio of total foreign assets plus liabilities to GDP and taken from the updated and extended version of the data set by Lane and Milesi-Ferretti (2007).

Inequality (IQ) is measured by the Gini index. We use the net Gini index, which takes redistributional measures into account and thus corresponds more closely to our theoretical motivation than its gross counterpart. The index is increasing in inequality and ranges from 0 to 100. For ease of interpretation, the Gini index is divided by 10 and thus defined over the interval from 0 to 10. Data is taken from Solt (2012).

While the time series properties of our variables rule out the use of country-specific

[3]See the Transparency International Homepage, `http://cpi.transparency.org/cpi2011`, 27.01.2012.

fixed effects, we include a set of 10 regional dummy variables η_r in order to control for unobserved, time-invariant determinants of corruption.[4] In the robustness analysis, two further summary indicators of geography will be included, following Rodrik and Rodriguez (2000): distance from the equator (D^E), taken from the Mayer and Zignago (2011) database, and the percentage of a country's land area that is in the tropics (L^T), which is from Doppelhofer and Weeks (2009) and has been extended through online search by the authors. Finally, ν_t are year-specific fixed effects and $u_{i,t}$ is an idiosyncratic error term.

The time period of our panel (t) ranges from 1995-2005, a choice which is mainly determined by the availability of data on the corruption index by Transparency International. It was first collected for 40 countries in 1995 and the coverage has increased to 178 countries in 2005. Taking availability of the data on the other variables used in our empirical analysis into account, we end up with an unbalanced panel, ranging from 32 up to 82 countries over the period 1995-2005 with a total of 695 observations. A detailed overview of the countries in our sample and the period coverage is given in Table A1 in the appendix.

2. Instrumental Variables and Identification

As emphasized in the literature, trade and financial openness are likely to be endogenous in equation (1), since the causality may also run from corruption to globalization (Ades and Di Tella, 1999; Emerson, 2006; Eichengreen and Leblang, 2008; Levchenko, 2011). We use instrumental variable (IV) estimation to obtain consistent estimates and to identify the causal effect of globalization on corruption.

As an instrument for trade openness, we follow the standard approach introduced by Frankel and Romer (1999) and use (the country-specific sum of the) predicted values from a bilateral gravity model that includes geographical variables only. In particular, we regress, for each year, bilateral trade (the log of imports plus exports as share of GDP) on country size (in terms of population and area), distance, a common border dummy, and interactions between all variables and the common border dummy. Then we generate predicted values for (the level of) bilateral trade flows, including all partner countries for which data on the explanatory variables is available, and sum them up to country-specific predicted trade shares. In order to obtain a time-variant instrument for trade, we use an approach similar to Feyrer (2009) and estimate the geographical bilateral gravity model (and generate predicted values) separately for each year of our sample period (1995-2005), thereby allowing all parameters to vary over time. This country- and year-specific 'geographical trade share' or proximity measure, referred to as $\widehat{TO}_{i,t}$, is then used as instrument for trade openness in the IV estimation of equation (1)."[5]

[4] In particular, dummy variables are used for the following regions: Sub-Saharan Africa, South Asia, East Asia, South-East Asia, Pacific Islands, Middle East and North Africa, Latin America, Caribbean and Non-Iberian America, Central and Eastern Europe, Western Europe. The data source is Norris (2009).

[5] Bilateral trade data are from the Correlates of War database (Barbieri and Keshk, 2012), geographical variables from the Mayer and Zignago (2011) database. As a first indication of instrument quality, the correlation between the actual bilateral trade data and the predicted values in our gravity model with some $14,000$ observations (on average per year) amounts to 0.63.

The gravity model has been shown to perform well not only for trade but also for financial flows (Portes and Rey, 2005; Guerin, 2006), where distance assumes the interpretation of informational frictions rather than trade costs. Likewise, Badinger (2009) shows that the performance of the geographical gravity instrument based on bilateral trade is similar for (aggregate) trade and financial openness. As a consequence, we will use $\widehat{TO}$ as an instrument for financial openness (FO) in equation (1) as well.
Obviously, when both openness measures are included in equation (1), additional instruments are required to separately identify the effects of trade and financial openness. This task is severely aggravated by the high correlation between TO and FO, which amounts to 0.40 in our sample.[6]

Following the identification strategy of Badinger (2009) we employ (de jure, policy related) measures of trade and financial integration as additional instruments. In particular, we use the trade and financial freedom indices of the Heritage Foundation, which have been employed, e.g., by Das and DiRienzo (2009).[7] The Heritage Foundation trade freedom index, referred to as F^{TO}, is a composite measure of the trade-weighted average tariff rate and non-tariff barriers (NTB) on trade in goods and services, taken from various sources such as the World Bank, the World Trade Organization (WTO) and the Office of the U.S. Trade Representative.

The corresponding financial freedom index F^{FO} focuses on the extent of government regulation of financial services, the degree of state intervention in banks and other financial firms, financial and capital market development, government influence on credit allocation and openness to foreign competition. It is based, among others, on data from the Economist Intelligence Unit, the International Monetary Fund, and the OECD. Both indices are scaled between zero and 100 with higher values indicating fewer restrictions on trade and financial markets.[8] Of course, the use of policy based variables as instruments raises some concerns, since they could be correlated with other domestic policies affecting corruption. Hence, it will be important to carefully test for instrument validity in the empirical analysis below.

3. Descriptive Statistics

Before turning to the estimation results, Table 1 provides some descriptives statistics on the main variables used. There is considerable variation in our dependent variable mainly over the cross-section. The lowest corruption values are observed for Canada, Denmark, Finland, Norway, and Switzerland, whereas examples for countries at the upper end of

[6] Using Granger causality tests, Aizenman and Noy (2009) show that trade and financial openness are intricately intertwined. In general, constructing convincing instruments that explain the exogenous variation in trade openness but not financial openness (and vice versa) is a challenging task, given that the underlying drivers of these two dimensions of openness are largely equivalent as evident from the good performance of the gravity model for both variables.

[7] While there are alternative de jure measures of trade and financial openness, the ones by the Heritage Foundation are time variant and have the largest coverage for our sample of countries and years.

[8] For a detailed description of the Heritage Foundation's methodology and data sources, see `http://www.heritage.org/index/book/methodology`.

the distribution are Bangladesh, Bolivia, and Tajikistan. Inequality is most pronounced in Chile, Honduras, and Panama, the most equal income distribution in terms of the Gini index is shown by Sweden, the Netherlands, and Denmark. Highest values for the political rights variable are obtained by the U.S. and also in many European countries, low values are observed for China, Vietnam, and Syria.

Notice that the correlations shown in the lower panel of Table 1 are suggestive and correspond to the theoretical expectations. GDP per capita, human capital, political rights and the openness measures are negatively correlated with corruption. Resource rents show a positive association with corruption. Inequality is positively correlated with corruption and shows a negative correlation with the two openness measures, which are themselves strongly correlated with each other. Finally, the (trade-based) gravity instrument shows a strong correlation with the two openness measures, which is slightly higher for trade openness.

Table 1: *Summary Statistics*

a) Descriptives	Mean	Median	Std.Dev.	Max.	Min.
CI	5.173	6.000	2.411	9.600	0.000
$\ln(Y/L)$	9.063	9.101	1.119	10.793	6.080
HC	8.454	8.805	2.360	12.869	1.024
PR	5.402	6.000	1.622	7.000	1.000
RR	0.043	0.016	0.082	0.733	0.000
TO	0.788	0.680	0.448	3.777	0.132
FO	2.030	1.399	2.121	19.851	0.397
IQ	3.680	3.510	0.867	6.178	2.171
$\widehat{TO}$	1.759	1.444	1.065	5.591	0.320
F^{TO}	68.472	72.800	13.550	85.000	0.000
F^{FO}	58.144	50.000	18.562	90.000	10.000

b) Correlations	CI	$\ln(Y/L)$	HC	PR	RR	TO	FO	IQ	$\widehat{TO}$	F^{TO}
$\ln(Y/L)$	-0.798									
HC	-0.572	0.746								
PR	-0.637	0.690	0.586							
RR	0.282	-0.163	-0.131	-0.420						
TO	-0.167	0.172	0.252	0.024	-0.062					
FO	-0.504	0.419	0.273	0.321	-0.147	0.396				
IQ	0.544	-0.550	-0.515	-0.387	0.176	-0.199	-0.282			
$\widehat{TO}$	-0.264	0.338	0.422	0.325	-0.224	0.528	0.347	-0.611		
F^{TO}	-0.544	0.640	0.581	0.576	-0.205	0.269	0.337	-0.307	0.354	
F^{FO}	-0.539	0.485	0.440	0.528	-0.405	0.148	0.413	-0.210	0.258	0.455

All descriptive statistics and correlations are based on a sample of 695 observations for 102 countries over the period 1995-2005. See the appendix for details.

IV Estimation Results

1. Separate Results for Trade Openness and Financial Openness

We start with the estimation results for model (1), including one of our two openness measures at a time. The first column in Table 2 reports the least squares estimates for the specification, where only the controls reflecting detection technologies (along with region and time dummies) are included. All variables enter significantly and show the expected sign. A higher level of development, reflected in higher per capita income, a better educated population, and more political participation and rights have a negative effect on (i.e., reduce) corruption. Revenues from resource exports significantly increase corruption in each specification. These results are in line with previous findings by Ades and Di Tella (1999), Leite and Weidmann (1999), Emerson (2006), and Bhattacharyya and Hodler (2010). The regional dummy variables, which are not reported in Table 2 for the sake of brevity, are jointly significant at 1 percent and will thus be included in all specifications.

To assess the economic significance of the coefficients, note that the average of the corruption index is 5.17. Hence, an increase in GDP per capita by 10 percent lowers corruption by 0.11 index points or 2.14 percent in relative terms (evaluated at the sample mean). One more year of schooling lowers corruption by 0.20 index points (or 2.40 percent of its mean). An increase in the political rights index by one standard deviation (i.e., 30 percent of its mean) reduces corruption by 0.19 index points (or 3.19 percent), the implied average elasticity is -0.12. Finally, natural resource rents have a strong positive effect. An increase of resource rents (measured as share of GDP) by one percentage point increases corruption by 0.03 index points (or 0.49 percent). The magnitude of these effects is very similar in the alternative specifications that will be considered in the following.

Table 2: *Corruption and Trade (Financial) Openness: Estimation Results for Model (1)*

	Dependent variable CI						
	(1) LS	(2a) LS	2b IV	(2c) IV	(3a) LS	(3b) IV	(3c) IV
$\ln(Y/L)$	-1.106***	-0.978***	-0.968***	-0.871***	-1.078***	-1.008***	-0.883***
	(0.122)	(0.080)	(0.085)	(0.103)	(0.105)	(0.079)	(0.098)
HC	-0.203***	-0.164***	-0.160***	-0.186***	-0.175***	-0.103**	-0.158***
	(0.036)	(0.033)	(0.033)	(0.033)	(0.035)	(0.046)	(0.039)
PR	-0.115**	-0.161***	-0.165***	-0.117***	-0.133***	-0.179***	-0.120***
	(0.051)	(0.039)	(0.040)	(0.043)	(0.046)	(0.042)	(0.044)
RR	2.537***	1.782***	1.720***	2.131***	2.254***	1.529***	2.068***
	(0.553)	(0.445)	(0.478)	(0.486)	(0.514)	(0.550)	(0.518)
TO		-0.860***	-0.931***	-0.524**			
		(0.124)	(0.211)	(0.204)			
FO					-0.124***	-0.443***	-0.232**
					(0.043)	(0.106)	(0.091)
IQ				0.626***			0.682***
				(0.098)			(0.100)
Hausman[1]			(0.681)	(0.065)		(0.002)	(0.265)
Instr. Quality[2]			279.26	302.76		39.81	49.36
Adj.R^2	0.792	0.807	0.807	0.816	0.799	0.751	0.808
SEE	1.100	1.060	1.060	1.033	1.081	1.204	1.057
Obs.	695	695	695	695	695	695	695

All models are based on an unbalanced panel of 102 countries over the period 1995-2005 (695 observations) and include region- and time-specific fixed effects. *, **, *** indicate significance at the 10, 5, and 1 percent level. Heteroskedasticity-robust standard errors in parentheses.

[1] Robust Hausman test for the exogeneity of TO and FO respectively.

[2] Stock and Yogo test for weak instruments: F-statistic on excluding $\widehat{TO}$ from first stage regression.

Column (2a) shows the results for model (1) with the trade openness measure (TO) included. In line with Ades and Di Tella (1999) we find a strong negative effect on corruption. Addressing endogeneity concerns and reestimating the model by IV, using the geographical trade share ($\widehat{TO}$) as an instrument for TO, its effect remains statistically significant and even increases in magnitude. The corresponding results in column (2b) suggest that an increase in trade openness by 10 percentage points lowers corruption by 0.09 index points (or 1.80 percent); the implied elasticity, evaluated at the sample mean amounts to -0.14.

Instrument quality is fine with an F-statistic of 279.28 on excluding $\widehat{TO}$ from the first stage regression; this by far exceeds the critical value (of 16.38) for the null hypothesis that instrument quality is below the highest quality level in terms of the Stock and Yogo (2005) weak instruments test. Also notice that the Hausman test clearly rejects the null of exogeneity, reinforcing empirically the importance to account for the endogeneity of trade as suggested by economic theory.

In a subsequent step we include inequality (IQ) as an additional regressor. If trade affects corruption mainly through inequality, we would expect the coefficient of TO to become insignificant or at least to be strongly reduced. As can be seen from the IV estimates in column (2c), the coefficient of TO is roughly halved to -0.52 but remains

statistically significant at 5 percent. This suggests that inequality is in fact one important yet not the only channel through which trade affects corruption. Regarding the role of inequality itself, its effect on corruption is sizeable: an increase by one standard deviation increases corruption by 0.54 index points (or 10.5 percent); the implied elasticity amounts to 0.45.

We next turn to the regressions including financial openness. As with trade, the Hausman tests reject the exogeneity of the variable FO. The quality of the (trade-based) gravity instrument $\widehat{TO}$ turns out weaker for FO, but the F-statistic of 39.81 is still far above the critical value for the highest instrument quality level judged by the Stock and Yogo (2005) weak instruments test. Comparing the least squares estimates in column (3a) with the IV estimates in column (3b), we observe an increase in the magnitude of the coefficient of FO from -0.12 to -0.44. The implied elasticity of corruption with respect to financial openness amounts to -0.17. Once we include inequality as an additional variable (column (3c)), the coefficient of FO goes down to -0.23, which is approximately half of its value in column (3b). Again, this indicates that inequality is an important transmission channel through which financial openness affects corruption.

Summing up our separate results so far, the exogeneity for both variables TO and FO is clearly rejected. Their effect on corruption is quantitatively similar with an elasticity around -0.15, which is halved when inequality is controlled for. A natural questions arising from the analysis so far is whether there is a separate role for each openness measure. In that case, the estimated effect of trade openness (TO) in columns (2b) and (2c) would be biased due to the omission of financial openess (FO), which is strongly correlated with TO. The same reasoning applies to the estimates including financial openness only in columns (3b) and (3c), where TO is omitted.

Before turning to the analysis including both openness measures, we provide an indirect test of whether the two openness measures both play an independent role. First, assume that financial openness (FO) were irrelevant in the model including trade openness (TO) in column (2c). In that case, financial openness would qualify as a (high quality) instrument for trade openness. Under the maintained assumption that the instrument $\widehat{TO}$ is valid, this can be checked by adding FO as an instrument for TO in the specification in column (2c) and testing for overidentifying restrictions (OID). It turns out that the OID test rejects the null of valid instrument with a p-value of 0.005 and the coefficient of TO increases in magnitude from -0.52 to -0.85. Analogously, adding TO as an instrument for FO to the model in column (3c), the coefficient of FO increases in magnitude from -0.23 to -0.37 and the null hypothesis of valid instruments is rejected by the OID test with a p-value of 0.08. We interpret these results as strong evidence that FO (TO) is an omitted variable in column (2c) (in column (3c)). In other words, both openness measures play an genuine role, i.e., they matter simultaneously and omitting one of the two variables will lead to biased estimates of the effect of the other.

2. Results for Models Including Both Trade and Financial Openness

The first two columns in Table 3 show the least squares estimates, when trade and financial openness are included at the same time. As before the Hausman test rejects the null of exogeneity of TO and FO such that we focus on the IV estimates in the following. Column (1b) reports the TSLS estimates, using $\widehat{TO}$ and the two policy measures F^{TO} and F^{FO} as instruments. It is worth emphasizing that the coefficients of TO and FO turn out to be very close with values of -0.34 and -0.29 respectively, and they are indistinguishable in statistical terms (The p-value of the F-test for parameter equality amounts to 0.88.) It should be noted, however, that the variable TO becomes insignificant, which is not too surprising in light of the high correlation with FO (and among the instruments), which inflates the standard errors. Adding further instruments does not appear to be a promising route. Even in the specification in column (1b), the OID test rejects the null of valid instruments at 5 percent, raising concerns about the use of the two policy based instruments F^{TO} and F^{FO} for identification.

Notwithstanding these qualifications, the overall results suggest that the effects of TO and FO are both qualitatively and quantitatively similar. Moreover, including inequality roughly halved the coefficient of both TO and FO (compare Table 1). Hence, we proceed with a restricted model in the following, including a joint openness measure $(TO + FO)$, which also allows us to dispense from using the policy based instruments F^{TO} and F^{FO} and stick with the geography based gravity instrument $\widehat{TO}$.

Column (1c) shows the results where the parameters of TO and FO are restricted to equality. Its coefficient turns out the be -0.30, which is smaller in magnitude than the corresponding estimates in Table 2, where TO (FO) apparently also captures the effect of FO (TO). The implied average elasticity with respect to the joint openness measure amounts to -0.16. Adding inequality in column (1d), we observe again a reduction in the effect of globalization by around one half. Hence, the qualitative conclusions from the separate regressions are unchanged, though the magnitude of the effect of each variable is smaller now. The model in column (1d) is our preferred specification, and we will further explore its robustness in the following.

Considering alternative functional forms, we first investigate whether there is a more intricate interplay between globalization and inequality by including an interaction term between $(TO + FO)$ and IQ. This hypothesizes that the effect of globalization is larger for higher levels of inequality. The results in column (2) provide no support for this specification.

Column (3) includes additional control variables, addressing the critique by Rodrik and Rodriguez (2000), who argue that the regression results by Frankel and Romer (1999) might be sensitive against including geographical variables. In particular, two summary indicators of geography are added to the specification in column (1d): distance from equator (D^E), and the share of a country's land area that is in the tropics (L^T).[9] As can be seen from the results in column (3), our key results are not affected; in fact, the coefficient of our measure of globalization becomes larger in magnitude and its p-value becomes

[9]For reasons of data availability, the sample is slightly reduced to 688 observations.

Table 3: *Corruption and Globalization: Estimation Results for Model (1)*

	Dependent variable CI											
	(1) LS	(1b) TSLS	(1c) IV	(1d) IV	(2) IV	(3) IV	(4a) IV	(4b) IV	(5a) IV	(5b) IV	(6a) IV	(6b) TSLS
$\ln(Y/L)$	-0.984***	-0.991***	-0.995***	-0.880***	-0.879***	-0.881***	-1.115***	-0.952***	-0.948***	-0.848***	-0.922***	-0.858***
	(0.079)	(0.080)	(0.078)	(0.099)	(0.098)	(0.094)	(0.085)	(0.104)	(0.070)	(0.089)	(0.092)	(0.076)
HC	-0.157***	-0.122***	-0.122***	-0.167***	-0.166***	-0.120***	-0.130***	-0.167***	-0.139***	-0.165***	-0.184***	-0.145***
	(0.033)	(0.037)	(0.040)	(0.037)	(0.038)	(0.036)	(0.039)	(0.039)	(0.031)	(0.029)	(0.035)	(0.033)
PR	-0.163***	-0.175***	-0.174***	-0.119***	-0.118***	-0.071*	-0.115***	-0.088**	-0.128***	-0.082**	-0.128***	-0.152***
	(0.039)	(0.039)	(0.040)	(0.043)	(0.045)	(0.041)	(0.040)	(0.045)	(0.036)	(0.040)	(0.042)	(0.038)
RR	1.767***	1.581***	1.591***	2.087***	2.090***	2.223***	2.268***	2.433***	2.142***	2.539***	2.257***	1.801***
	(0.445)	(0.488)	(0.508)	(0.503)	(0.505)	(0.507)	(0.460)	(0.478)	(0.431)	(0.442)	(0.486)	(0.444)
TO	-0.742***	-0.343										
	(0.134)	(0.300)										
FO	-0.052	-0.288***										
	(0.039)	(0.098)										
$(TO+FO)$			-0.300***	-0.161**	-0.127	-0.259***			-0.335***	-0.192***	-0.218***	-0.339***
			(0.068)	(0.062)	(0.315)	(0.058)			(0.071)	(0.062)	(0.073)	(0.061)
IQ				0.665***	0.696***	0.555***		0.633***		0.617***	0.691***	0.708***
				(0.099)	(0.269)	(0.094)		(0.096)		(0.085)	(0.098)	(0.099)
$(TO+FO)\times IQ$					-0.011							
					(0.097)							
D^E						-0.058***						
						(0.006)						
L^T						-0.706**						
						(0.280)						
$\ln(TO+FO)$							-0.882***	-0.518**				
							(0.219)	(0.227)				
$(TO+FO)\times D$											0.174***	0.137**
											(0.057)	(0.055)
Hausman[1]		(0.010)	(0.010)	(0.695)	(0.620)	(0.695)	(0.404)	(0.011)	(0.058)	(0.615)	(0.664)	(0.021)
Instr. Quality[2]		24.719	73.345	87.300	48.812	87.300	167.253	176.332	108.780	127.522	42.074	34.038
OID		0.044				0.						(0.014)
Adj.R^2	0.807	0.786	0.785	0.815	0.815	0.828	0.818	0.824	0.856	0.873	0.820	0.815
SEE	1.058	1.114	1.119	1.038	1.038	1.005	1.028	1.011	0.903	0.849	1.023	1.036
Obs.	695	695	695	695	695	688	695	695	659	659	695	695

All models are based on an unbalanced panel of 102 countries over the period 1995-2005 (695 observations) and include region- and time-specific fixed effects. Column (3) includes as additional geographical control variables: distance to equator (D^E), and the share of land area that is in the tropics (L^T). In columns (6a) and (6b), D represents a dummy for OECD countries. *, **, *** indicate significance at the 10, 5, and 1 percent level. Heteroskedasticity-robust standard errors in parentheses.

[1] Robust Hausman test for the exogeneity of TO and FO respectively.

[2] Stock and Yogo test for weak instruments: F-statistic on excluding $\widehat{TO}$ from first stage regression; Cragg Donald statistic in columns (1b), (3), and (6b).

smaller. Moreover, the p-value of the Hausman test for exogeneity is virtually unchanged, supporting the exclusion restrictions underlying the IV approach for our empirical model.[10]

Columns (4a) and (4b) consider a specification where the openness measure is used in logarithmic form, allowing for decreasing effects of globalization. Theory is silent on whether to use to the level or the log, specification tests are inconclusive across the specifications, and the adjusted R^2 is very close in both models. Therefore, it is reassuring that our results are qualitatively and quantitatively insensitive with respect to the choice of levels or logs. As can bee seen from column (4a), the coefficient of the log of the joint openness measure ($TO + FO$) is highly significant and points to an elasticity of -0.17, which is very close to the average elasticity from the estimates in levels (-0.16). Including inequality in column (4b) reduces the effect of globalization by 40 percent.

To ensure that our results are not driven by outliers, we reestimate the preferred specifications in columns (1c) and (1d), excluding all observations where the standardized residuals exceed an absolute value of 2. Results in columns (5a) and (5b) show that the results are essentially unchanged. Finally, we explore the subsample stability with respect to the cross-section dimension. This is potentially important, since the the theoretical motivation of our empirical model is stronger for developing countries. Given the relatively small cross-section dimension, we test for differences by including an interaction term of our joint openness measure ($TO + FO$) with a dummy for OECD membership (D).[11]

Using $\widehat{TO}$ (and its interaction with the OECD dummy) as instruments, we find that the coefficient of the interaction term, representing the difference of the effect for the OECD group, is positive and significant (column (6a)). The magnitude suggests that the effect of globalization on corruption is close to zero for the group of OECD countries, which is also confirmed by an F-test on the sum of the parameters for ($TO+FO$) and ($TO+FO)\times D$. If the policy based variables F^{TO} and F^{FO} are added as further instruments (column (6b), the effect of globalization on corruption is reduced in magnitude for OECD countries but not eliminated. Hence, we carefully argue that the effect is more pronounced in developing countries. Given the small subsample, the results should not be overstressed but taken together with the theoretical rationale underlying the globalization-corruption nexus, with the relatively high level of trade and financial openness and relatively low level of corruption in OECD countries, a smaller effect for this subgroup of countries appears to be a very plausible result.

Our main results establish a negative effect of globalization on corruption. Although the number of questions that can be answered with a reduced form empirical model such as equation (1) is limited, we argue that our results indirectly suggest that globalization reduces inequality. At least it is very difficult to think of any convincing transmission channels and interrelationships that could reconcile our key results (the negative effect of

[10] We have also estimated the other specifications in Tables 1 and 2 with the two additional control variables included and obtained qualitatively identical results: Globalization (inequality) has a negative (positive) effect on corruption, and the effect of trade is roughly halved, once inequality is controlled for.

[11] We do not include a separate constant for OECD countries, which would be highly collinear with the regional dummies.

globalization on corruption, the positive effect of inequality, and the reduction in the effect of globalization once inequality is controlled for) with a positive (or insignificant) effect of globalization on inequality.

V Conclusions

Using an unbalanced panel of 102 countries over the period 1995-2005, this paper investigates the determinants of corruption, thereby paying particular attention to the effects of globalization and its interplay with inequality, which has been suggested as a key transmission channel in theoretical work by Acemoglu and Robinson (2005). As measures of globalization, both trade and financial openness are considered; to establish their causal effect on corruption, we exploit the exogeneity of countries' geographical characteristics as determinants of openness.

In line with previous studies we find that 'detection technologies', approximated by GDP per capita, education and political rights, significantly reduce corruption while natural resource rents increase corruption. Globalization significantly reduces corruption with a more pronounced effect in developing countries. We find that both measures of globalization play an independent role in reducing corruption, such that omitting financial (trade) openness leads to an upward bias of the estimated effect of trade (financial) openness. However, the quantitative effects of trade and financial openness on globalization are very similar, which allows focussing on a joint measure of globalization.

We find that inequality increases corruption. Moreover, the effect of both trade and financial openness (and the joint globalization measure) are halved, once the effect of inequality is controlled for. This confirms the role of inequality as a transmission channel through which globalization affects corruption. The remaining and equally important effect is most likely to be due to the pro-competitive effect of globalization, which has been emphasized in previous studies (Ades and Di Tella, 1999; Emerson, 2006).

Overall, our results also provide indirect evidence that globalization has a negative effect on inequality. However, the transmission channels through which globalization affects corruption (inequality, and thereby democracy and institutional quality) are manyfold and intertwined, with causality potentially running into both directions. Moreover, changes in the institutional arrangements required by supranational trade associations (Das and DiRienzo, 2009), trade patterns and countries' technological and institutional position relative to its trading partners (Levchenko, 2011) can be expected to have both direct and indirect effects on corruption as well as on globalization. Against this background, considering and disentangling these interrelationship in a simultaneous system of equations and developing a comprehensive identification strategy, appear to be a fruitful and challenging avenue for future research.

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Appendix

Table A1: *Country and Year Coverage*

Albania	(1999-2005)	Latvia	(1998-2005)
Argentina	(1995-2005)	Lithuania	(1999-2005)
Armenia	(1999-2005)	Malawi	(1998-2004)
Australia	(1995-2004)	Malaysia	(1995-2004)
Austria	(1995-2005)	Malta	(2004-2005)
Bangladesh	(1996-2005)	Mauritius	(1999-2001)
Belgium	(1996-2005)	Mexico	(1995-2005)
Bolivia	(1996-2005)	Moldova	(1999-2005)
Brazil	(1995-2005)	Mongolia	(1999-2005)
Bulgaria	(1998-2005)	Morocco	(1998-1999)
Cameroon	(1996-2001)	Mozambique	(1999-2003)
Canada	(1995-2000)	Nepal	(2004-2004)
Chile	(1995-2003)	Netherlands	(1996-2005)
China	(1995-2005)	New Zealand	(1996-2004)
Colombia	(1995-2003)	Nicaragua	(1998-2005)
Congo, Republic of	(2005-2005)	Niger	(2005-2005)
Costa Rica	(1997-2005)	Norway	(1996-2005)
Cote D'Ivoire	(2001-2002)	Pakistan	(1995-2005)
Croatia	(1999-2005)	Panama	(2001-2004)
Cyprus	(2003-2005)	Paraguay	(1998-2005)
Czech Republic	(1996-2005)	Peru	(1998-2005)
Denmark	(1996-2005)	Philippines	(1995-2003)
Dominican Republic	(2001-2005)	Poland	(1996-2005)
Ecuador	(1996-2005)	Portugal	(1995-2005)
Egypt	(1996-2005)	Romania	(1997-2005)
El Salvador	(1998-2005)	Russian Federation	(1996-2005)
Estonia	(1998-2005)	Senegal	(1998-2005)
Finland	(1996-2005)	Sierra Leone	(2003-2003)
France	(1995-2005)	Singapore	(1995-2000)
Gabon	(2005-2005)	Slovakia	(1998-2005)
Gambia	(2003-2003)	Slovenia	(1999-2005)
Germany	(1995-2005)	South Africa	(1995-2000)
Ghana	(1998-1999)	Spain	(1995-2005)
Greece	(1995-2005)	Sri Lanka	(2002-2002)
Guatemala	(1998-2002)	Sweden	(1995-2005)
Honduras	(1998-2005)	Switzerland	(1996-2002)
Hungary	(1995-2005)	Syria	(2004-2004)
Iceland	(2004-2005)	Tajikistan	(2003-2004)
India	(1995-2005)	Tanzania	(1998-2000)
Indonesia	(1995-2005)	Thailand	(1995-2004)
Iran	(2003-2005)	Tunisia	(1998-2000)
Ireland	(1995-2005)	Turkey	(1995-2005)
Israel	(1996-2001)	Uganda	(1996-2005)
Italy	(1995-2005)	Ukraine	(1998-2005)
Jamaica	(1998-2004)	United Kingdom	(1995-2005)
Japan	(1995-1998)	United States	(1995-2004)
Jordan	(1996-2003)	Uruguay	(1997-2005)
Kazakhstan	(1999-2005)	Venezuela	(1995-2005)
Kenya	(1996-2005)	Vietnam	(1998-2004)
Korea, Rep.	(1995-2003)	Yemen	(2003-2005)
Kyrgyzstan	(1999-2005)	Zambia	(1998-2004)

Number of cross-sections for each year: 1995 (32), 1996 (50), 1997 (38), 1998 (73), 1999 (82), 2000 (70), 2001 (72), 2002 (72), 2003 (73), 2004 (67), 2005 (66)