Does external trade promote financial development?*

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Abstract

Several recent papers have argued that trade and financial development may be linked, either for political economy reasons, or because openness influences the demand for external finance. In this paper we use the cross-country and time-series variation in openness to study the relationship between trade and finance in more detail. Our results suggest that increases in goods market openness are typically followed by sustained increases in financial depth.

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1 Introduction

Many of the world's economies are becoming more open to international trade. Not surprisingly, this has led to renewed interest in the consequences of openness for competition, technology transfer, and productivity. In this paper, we ask whether openness promotes financial development, using cross-section methods to examine the long-run relationship, and recently developed panel data methods to examine Granger-causality. Our panel data estimates suggest that, for lower-income countries, increases in goods market openness are typically followed by sustained increases in financial development.

Trade and finance have been connected in the literature in at least two different ways, which can be loosely characterized as supply-side and demand-side. In an important paper, Rajan and Zingales (2003) emphasize the supply-side role of interest groups, and especially the vested interests of incumbent industrialists and financial intermediaries. Incumbents, worried by the threat of entry, have strong incentives to resist financial development. These incentives are weakened if a country becomes more open to foreign competition or to international flows of capital. In this view, goods market openness can improve the supply of external finance, because it aligns the interests of the economically powerful more closely with financial development.[1]

In the long run, financial depth and the quality of financial intermediation are equilibrium outcomes that will vary with the demand for external finance, as well as supply-side forces (Do and Levchenko 2004). Svaleryd and Vlachos (2002) emphasize the role of risk diversification. To the extent that openness is associated with greater risks, such as increased exposure to external demand shocks or foreign competition, it will create new demands for external finance. Firms will need credit in order to overcome short-term cashflow problems and adverse shocks. In this view, the effects of trade on finance are likely to work primarily through the demand side.

Do and Levchenko (2004) show how the demand for finance can interact with the quality of financial intermediation. In their model, opening to trade leads some countries to specialize in production sectors that make intensive use of external finance. As the financial sector expands, liquidity increases, and this reduces the risks faced by entrepreneurs using external finance; the change in the demand for finance has improved the conditions under which credit is made available. More generally, increased openness may promote or reduce financial development, depending on whether financially intensive sectors expand or contract.

These long-run effects may be complemented by other, more short-lived considerations. Although our empirical work examines the long-run relationship between trade and finance, it acknowledges that short-run effects will also play a role. For example,

[1] The Rajan and Zingales view assigns importance to financial openness as well as goods market openness. We discuss the role of financial openness in more detail in section 2 below.

when developing countries liberalize trade on a major scale, such as Mexico under the NAFTA, restructuring and investment are likely to increase the demand for external finance. If trade liberalizations are followed by investment and lending booms, there could be a strong association between openness and finance in the short run, whether or not there is an association in the long run.[2] Giavazzi and Tabellini (2004) find that trade liberalizations are indeed often followed by higher investment.

When examining the effects of external trade on financial depth, disentangling cause and effect will not be straightforward, partly because finance may influence trade as well as vice versa. Our empirical work will address this problem in two ways. First, we follow Rajan and Zingales (2003) and Stulz and Williamson (2003) in using the measure of "natural openness" due to Frankel and Romer (1999). We find evidence that openness and finance are strongly associated for higher-income countries, but not for lower-income. Perhaps contrary to some of the ideas put forward in Rajan and Zingales, we find that openness has a much stronger association with bank-based finance than with stock market development.

Our second approach, which forms the heart of the paper, is to use panel data methods. These allow us to exploit the substantial time series variation in goods market openness. For example, in a subset of countries, openness has risen quite sharply. To show this we use a standard measure of openness, namely imports plus exports, relative to GDP, all in current prices. For a sample of 82 countries, the median extent of openness rose from 43% in 1960-64 to 66% in 1995-99. There are 59 countries in which openness rose by at least ten percentage points over this period, and 34 in which it rose by at least 25 percentage points. To show this pattern in more detail, figure 1 is a smoothed (kernel density) plot of the distributions for 1960-64 and 1995-99. The rightward shift in the distribution of openness over the forty-year period is clear.[3]

Panel data methods allow us to examine the consequences of increased openness for financial development, and to see whether timing patterns are consistent with a causal effect. In a forty-year data set for 88 countries, we examine whether changes in openness precede (Granger-cause) changes in financial depth. From the estimated models, we find strong effects of openness on financial development in the whole sample, and for lower-income countries, but not for higher income countries. There is some evidence that these effects persist into the long run, and do not simply reflect temporary

[2]One mechanism here may be information asymmetries. To the extent that performance in export markets is a useful and observable index of a company's productivity, increased openness can reduce information asymmetries among different banks. Dell'Ariccia and Marquez (2005) present a formal model in which reduced information asymmetries can generate a lending boom.

[3]Note that this shift is not universal, and not all countries are more open now than twenty or thirty years ago. See Dollar and Kraay (2004) for more details on trends in openness. Their analysis confirms that the time series variation in the trade share can be substantial, helping to motivate our emphasis on a panel data approach.

booms in bank lending.

It is important to note that, for lower-income countries, the panel data results indicate stronger positive effects of trade than we find in the cross-section. Increases in openness are systematically associated with subsequent increases in financial depth. It is possible that the (conflicting) cross-section estimates are contaminated by the presence of individual effects that are correlated with openness. For the higher-income countries, the situation is reversed: the panel data results are weaker than the cross-section. The imprecise panel data estimates may arise partly because of the small cross-section dimension of the higher-income panel.

The remainder of the paper is organized as follows. Section 2 discusses the literature on financial development in more detail. Section 3 describes the data sources, and the measures of financial development that we adopt. Section 4 focuses on the cross-section methods and results. In section 5, we discuss the methods and tests used in the panel data analysis. Section 6 reports the panel data estimates, and can be seen as the heart of the paper. In section 7 we briefly summarize our conclusions.

2 Trade and finance

In this section, we sketch the theoretical background to the paper, and expand on the arguments in the introduction. We discuss the links between finance and growth, the role of trade and other determinants of financial development, and the possibility that finance, in turn, can influence the extent and structure of trade.

Conventionally, trade is seen as a way to benefit from specialization and scale economies. But if external trade promotes financial development, this offers a more complex route by which trade may raise productivity and living standards. Beginning with King and Levine (1993), the study of finance and growth has flourished, and the evidence that financial sector development plays a role in growth is increasingly persuasive; Rajan and Zingales (1998) is one influential contribution. The literature on finance and growth is reviewed by Levine (1997, 2005).

There has been less work on the determinants of financial development, and our work contributes to this emerging line of research. The central question is why entrepreneurs and firms appear to have easier access to external finance in some countries than others. As discussed in the introduction, Rajan and Zingales (2003) argue that such differences arise because of political economy considerations.[4] Incumbent industrial firms, and perhaps domestic financial intermediaries, will wish to block entry to their sectors: they have a direct interest in maintaining an underdeveloped financial sec-

[4]For a general discussion of the political economy of financial development, see Pagano and Volpin (2001a).

tor. These incentives may weakened by openness, however. A combination of foreign competition through external trade, and openness to capital flows, will tend to align the interests of incumbents more closely with financial development. Although Rajan and Zingales give most emphasis to arms-length finance (especially stock markets) some of the arguments can be applied to bank-based external finance as well.

In support of their arguments, Rajan and Zingales document long-term swings in financial depth, using data for 1913 and decades since. Some countries have seen long-term declines in the importance of external finance relative to GDP, before a resurgence in the 1980s and 1990s. An attractive feature of the Rajan and Zingales theory is that it can explain these reversals in political economy terms. Openness to trade was relatively low in the wake of the Great Depression, and even the advanced industrial economies maintained capital controls for several decades after 1945, motivating the familiar point that the economy was less globalized in mid-century than at the beginning and end of the twentieth century. These reductions in openness, perhaps helping to create a political constituency opposed to financial sector development, could explain the U-shaped path that has been taken by the importance of external finance relative to GDP.[5]

Other mechanisms that could link trade and finance do not rely as strongly on political economy, and give more emphasis to the demand side. As discussed in the introduction, Svaleryd and Vlachos (2002) argue that openness may be associated with greater risks, including exposure to external shocks and foreign competition. This will encourage the development of financial markets that can be used to diversify such risks, and that allow firms to overcome short-term cashflow problems or adverse shocks.[6] Do and Levchenko (2004) emphasize the interaction between openness, the relative size of financially intensive sectors, and the quality of financial intermediation.

Openness and financial development may also be linked in simpler ways. The cross-country study of Levine and Renelt (1992) identified a robust correlation between openness and the share of investment in GDP, and if trading economies are also high investment economies, this could promote financial development. Finally, some of the ideas in Acemoglu and Zilibotti (1999) could also be used to link trade and finance: in their framework, if greater openness makes relative performance evaluation easier, this would encourage market-based financial intermediation rather than direct monitoring.

Our central focus is goods market openness, but previous research also points to

[5]A more ambitious argument is that political forces lead either to a 'corporatist' political settlement, with low investor protection and high employment protection, or a non-corporatist (Anglo-Saxon) outcome with flexible labour markets and greater financial depth (Pagano and Volpin 2001b). Commentary in the media often assumes, perhaps wrongly, that increased trade and globalization has made a corporatist settlement harder to sustain, providing another way to link trade and finance.

[6]The evidence for this should not be overstated. Using stock market data for emerging markets, Li et al. (2004) find that goods market openness is associated with greater market-wide variation, but not greater firm-specific variation.

the importance of financial openness: for example, the extent of restrictions on equity ownership by foreign investors. Alessandria and Qian (2005) use a formal model to argue that capital account liberalization has ambiguous effects on the efficiency of domestic financial intermediaries, while Chinn and Ito (2005) and Law and Demetriades (2005) consider the empirical evidence. There are at least three good reasons for complementing these studies with a consideration of goods market openness. First, goods market openness will be more relevant for the many developing countries where domestic equity markets are absent or small in size. Second, there is often richer time series variation in goods market openness than in measures of capital controls.[7] Third, goods market openness and financial openness are not independent. The empirical work of Aizenman and Noy (2003), Aizenman (2004) and Chinn and Ito (2005) suggests that capital account liberalization is often preceded by goods market openness, perhaps because trade integration makes restrictions on capital flows harder to sustain.

As background to our work, we now consider other determinants of financial depth. Several contributions have examined the effects of the legal, regulatory and macroeconomic environment on the functioning of the banking sector and equity markets. Most prominently, La Porta et al. (1998) argue that the origins of the legal code are important for financial development, because legal systems differ in their treatment of creditors and shareholders, and in contract enforcement. La Porta et al. argue that the English common law tradition protects the rights of minority shareholders and creditors, while the French civil code is associated with less efficient contract enforcement, weaker investor protection and possibly higher corruption. Countries with German or Scandinavian legal origins are said to have intermediate levels of investor protection and contract enforcement.[8]

Whether or not the origins of the legal code matter, government regulation clearly plays a strong role. The starting point of the arguments in Rajan and Zingales (2003) is that government regulation is needed to ensure effective contract enforcement, and transparency in accounting and disclosure. Regulations concerning information disclosure, accountings standards, permissible practice of banks and deposit insurance do appear to have material effects on financial development (Mayer and Sussman 2001).

Less central to the current paper, the macroeconomic environment may also be relevant. Huybens and Smith (1999) examine theoretically, and Boyd, Levine and Smith (2001) empirically, the effects of inflation on financial depth. They conclude

[7]The availability of long spans of data on goods market openness lends itself to conventional panel data methods. In contrast, given the discrete nature of many capital account liberalizations, event studies may be the best approach in that context, as in Henry (2003).

[8]As with any innovative line of research, this is not without its critics. Looking at historical data, Rajan and Zingales (2003) find that common law legal codes have not always been associated with greater financial development. They argue that law could play a stronger role in filtering the effects of interest groups and incentives than in influencing the overall level of financial development.

that economies with higher inflation rates are likely to have smaller, less active, and less efficient banks and equity markets.

Finally, it is worth pointing out that financial development can, in turn, influence the structure and extent of trade. Two papers by Beck (2002, 2003) have examined this issue in detail. Drawing on the arguments of Kletzer and Bardhan (1987), Beck (2002) develops a model in which countries with better-developed financial markets will tend to have a comparative advantage in manufacturing.[9] Using a 30-year panel with 65 countries, he shows that financial depth is associated with a higher level of manufacturing exports and a higher trade balance in manufactured goods. In a companion paper, Beck (2003) shows that economies with greater financial depth have higher manufacturing export shares and higher trade balances in industries that rely more on external finance. These results help to motivate our Granger-causality approach: since finance may influence trade, as well as vice versa, evidence on timing becomes especially important.

3 The data

This section describes the data and variables that we will use in our analysis. Section 3.1 describes the data on openness and GDP. In section 3.2, we will outline some of the widely used measures of financial development. Our empirical work will combine the standard measures into aggregate indices of financial development, an approach that we describe in section 3.3. Definitions of the main variables we use, and information on data sources, can be found in Table 1.

3.1 Basic data

Our cross-section sample relates to the period 1990-2001. We have excluded countries with populations of less than 500,000 in 1990, using population data from the Penn World Table (PWT).[10] We also exclude transition economies from the sample.

To measure initial GDP in our regressions, we use real chain-weighted GDP per capita (denoted RGDPCH in PWT 6.1) averaged over 1988-90 to lessen the effect of temporary measurement errors and departures from trend. The data on openness are also from PWT 6.1. Openness is measured as the sum of exports and imports divided by GDP, in either current prices (*OPENC*) or constant international prices (*OPENK*), averaged over 1988-1990. Following Svaleryd and Vlachos (2002), we always exclude Hong Kong and Singapore, two city-states for which openness is unusually high, re-

[9] Recent theoretical work by Ju and Wei (2005) and Wynne (2005) also develops connections between financial development and trade specialization.

[10] The exact source is version 6.1 due to Heston et al. (2002).

flecting transit trade and production that involves a high import content (for example, electronics). We also exclude other countries for which measured openness exceeds 150%.

In order to capture the geographic predisposition to trade, we use the measure of natural openness due to Frankel and Romer (1999) and call this *CTRADE*, for constructed trade share. Information on the country of legal origin, and classifications by income levels and export specialization, are obtained from the Global Development Network (GDN) Growth Database. To keep sample sizes reasonably large, our empirical work aggregates the GDN income classes into two larger groups: lower-income countries made up of low-income and lower-middle income countries, and higher-income countries made up of high-income and upper-middle income countries.

3.2 Standard measures of financial development

A number of measures of financial development have been used in recent work. Our basic data on financial development are from the Financial Structure Database introduced by Beck et al. (2000). We average standard indicators over 1990-2001, omitting any observations for which fewer than three years of data available. We now summarize in more detail the available indicators, drawing on Demirguc-Kunt and Levine (1999) in particular. We start with indicators of bank-based financial depth, which play the dominant role in our empirical work, and then turn to indicators of stock market development.

For measuring overall financial development, an especially popular measure is the ratio of liquid liabilities to GDP, based on the liquid liabilities of the financial system (currency plus demand and interest-bearing liabilities of bank and nonbank financial intermediaries). This measure has been used by McKinnon (1973) and King and Levine (1993) among others; we denote this measure *LLY,* as is standard in the literature. Other widely used measures include the ratio to GDP of credit issued to the private sector by banks and other financial intermediaries (denoted *PRIVO*) and the ratio of commercial bank assets to the sum of commercial bank assets and central bank assets (denoted *BTOT*). *LLY*, *PRIVO* and *BTOT* are highly correlated. They will be used to construct the main indicator of financial depth in our later panel data analysis, given the availability of long spans of data for these measures.

The Beck et al. (2000) database also includes two measures of the efficiency of financial intermediation. The variable *OVC* is the ratio of overhead costs to total bank assets. In the short run, high overhead costs may be related to investments by competitive banks in improving financial services, but over longer time spans, high overhead costs are likely to reflect inefficiency and a lack of competition. A second measure, the Net Interest Margin or *NIM*, equals the difference between bank interest income and interest expenses, divided by total assets. Again, high values for this variable tend to suggest a lack of competition among banks.

Recently, some studies have taken a wider view of financial development. Levine and Zervos (1998) discuss the independent effects of banks and stock markets on economic growth, rather than focusing on simply the extent of intermediation. Some measures of stock market development, such as stock market turnover, can be seen as indices of financial sector efficiency or sophistication rather than simply financial depth. As well as a standard measure, market capitalization relative to GDP (*MCAP*), Levine and Zervos (1998) use Total Value Traded (*TVT*) as an indicator of stock market activity. This is the ratio of trades in domestic shares (on domestic exchanges) to GDP, and can be used to gauge market liquidity relative to the size of the economy. Finally, the Turnover Ratio (*TOR*) is the ratio of trades in domestic shares to market capitalization. High values for *TOR* indicate a more active equity market, which may be associated with a relatively efficient allocation of capital.

3.3 New measures of financial development

The indicators described in the above section are standard. Since they are all intended as proxies for an underlying, latent variable - financial development - there may be significant advantages in combining them. This could help to alleviate measurement errors and outlier problems that might be associated with the use of a single indicator. We take one of the simplest possible approaches, namely to use principal components analysis. The method takes N specific indicators and yields new indices (the principal components) $P_1, P_2, ...P_N$ that are mutually uncorrelated and so capture different dimensions of the data. In our work we use solely the first principal component. Formally, this is defined by a vector of weights $a = (a_1, a_2, ..., a_N)'$ on the (standardized) indicators $X = (X_1, X_2, ...X_N)'$ such that $a'X$ has the maximum variance for any possible weights, subject to the constraint that $a'a = 1$.

We use this method to aggregate different sets of components into five new measures of financial depth. Their structure can be seen from Table 2, together with the weights on each component.[11] Our first aggregate measure is designed to capture overall financial development, and is denoted *FD*. This is based on the complete set of eight components, namely *LLY*, *PRIVO, BTOT, OVC, NIM, MCAP, TVT* and *TOR*. The first principal component accounts for 63% of the variation in these eight indicators. Here and subsequently, all the weights have the expected signs: positive for all variables except *OVC* and *NIM*, given that high values for these latter two variables indicate inefficiency in the financial sector.

[11] All components are measured in natural logarithms, which helps to reduce outlier problems in this particular application, and are then standardized to have mean zero and unit standard deviation.

Our second measure, *FDSIZE*, is effectively the average of *LLY* and *MCAP*, and provides a summary of the combined importance of bank-based and equity-based finance, relative to GDP. In contrast, *FDEFF* is designed to capture financial efficiency, and is based on *OVC, NIM, TVT* and *TOR*. A fourth measure, *FDBANK*, captures the extent of bank-based intermediation. This uses *LLY, PRIVO, BTOT, OVC* and *NIM*, and accounts for 71% of the variation in these five indicators. *FDSTOCK* captures equity market development, and accounts for 86% of the variation in *MCAP, TVT* and *TOR*. Finally, a measure of financial depth, *FDEPTH*, uses only *LLY, PRIVO* and *BTOT*, and accounts for 74% of the variation in these indicators.

By construction, all the new indices have a mean of zero. Panel A of Table 3 provides some other descriptive statistics for the new indicators and their components. In panel C of Table 3, we present the correlations among the new measures. As expected, they are highly correlated with one another. It is interesting to note that by far the lowest correlation is between *FDBANK* and *FDSTOCK* (0.63), reflecting the tendency for intermediation to be either bank-based or equity-based. Another interesting aspect of Table 3 is that the correlations between openness and the new indicators, shown in Panel C, are noticeably higher than when the original proxies are used (Panel B). This is consistent with the idea that aggregation of the original measures has reduced measurement error. The correlations with openness are still low, but these simple associations do not control for the level of development and other influences on financial depth.

We now take a look at scatter plots of the trade-finance relationship. Figure 2 plots two measures of financial development, *LLY* and *FD*, against openness. Openness and *LLY* are measured in natural logarithms, which helps to reduce outlier problems in this case. In figure 2, there is some evidence that financial development and openness are positively related, although outliers such as Japan (coded JPN in the figure) and El Salvador (SLV) may obscure the relationship to some extent. In figure 3, we present partial scatter plots between the same sets of variables, after conditioning on initial GDP per capita, and legal origin dummies (for English, French and German legal origin). Again, there is some evidence of a positive association between financial development and openness, whether using *LLY* or the additional information contained in the new indicator *FD*.

4 Cross-section: methods and results

4.1 Cross-section methods

Our basic framework will be regressions of the form:

$$Finance = \alpha + \beta \log(OPENC) + f(controls) + \varepsilon \tag{1}$$

Here $Finance$ is the level of financial development, based on either individual financial indicators, or our new aggregate indices, measured over 1990-2001. The explanatory variables are the natural logarithm of openness, and some controls. As in Rajan and Zingales (2003) we condition on the log of real GDP per capita, to control for the demand for finance. Our other explanatory variables are dummies for country of legal origin, as in La Porta et al. (1998), Beck et al. (2003) and Berkowitz et al. (2003).

Figures 2 and 3 suggested that influential outliers could play an important role in any cross-section analysis. To increase robustness in this dimension, we estimate our models in two stages. We first apply median or least absolute deviation regression, in which the parameters are chosen to minimise the sum of the absolute values of the residuals, rather than the sum of their squares as in OLS. We then exclude any observations for which the LAD residual is more than two standard deviations from the mean residual, before estimating the model by OLS or GMM. This procedure is not perfect, but should help to exclude the worst outliers, including some that would not be identified by more conventional OLS diagnostics.

Once outliers have been excluded, we run straightforward regressions using OLS. Since there are many reasons that openness could be correlated with the error term in such a regression - including reverse causality, measurement error and omitted variables - we supplement OLS with IV procedures. We treat openness as an endogenous explanatory variable, and use the Frankel-Romer (1999) trade share as an instrument.[12] Their work models openness using the size of the domestic population (given that large countries trade a lower fraction of GDP internationally) and the proximity of large external markets. Our maintained assumption will be that this measure of "natural openness" affects financial development through external trade, but is uncorrelated with the error term in the structural equation. This rules out, among other things, any influences of geography on financial development that act through routes other than openness.

For the Frankel-Romer measure of natural openness to be a good instrument, it must be correlated with openness, after conditioning on other (exogenous) variables. We investigate instrument relevance using a F-test on the excluded instrument in the first-stage regression of 2SLS (the reduced form regression of openness on all the exogenous variables, including natural openness). These tests suggest that natural openness is not a weak instrument, consistent with the findings in Frankel and Romer (1999).

We use 2SLS to construct diagnostic tests, including a test for heteroskedasticity based on Pagan and Hall (1983). Compared to OLS, the use of IV procedures is associated with a loss of efficiency if openness is uncorrelated with the error term. We therefore implement two tests for endogeneity. The first is a standard Wu-Hausman test

[12] This follows similar analyses in Do and Levchenko (2004), Rajan and Zingales (2003) and Stulz and Williamson (2003), among others.

based on auxiliary regressions in the 2SLS case.[13] The second is a difference-Sargan type test, sometimes called the C statistic in the GMM context. Not surprisingly, the results are close to those of the auxiliary regression approach. Both tests indicate that openness is endogenous in many of the models we consider.

4.2 Cross-section results

This section presents cross-country regression results. To anticipate some of our main findings, partitioning the sample into a higher-income group and a lower-income group will be an important step. Openness is associated with financial development in higher-income countries, but the evidence for this is much weaker in the lower-income sample. We find this when using the individual finance measures and also when using the new aggregate indicators.

We present our first set of results in Table 4. We report the point estimates and heteroskedasticity-robust p-values only for openness; the estimates for initial GDP and the legal origin dummies are not reported, to save space. The relationship between openness and finance is weak for the sample as a whole. In the higher-income group of countries, however, Panel A of Table 5 shows that openness and bank-based finance are strongly associated. This result emerges whether using OLS or GMM, and also when using individual finance measures like *LLY* and *PRIVO* (results not reported).

In Panel B of Table 5, we consider the lower-income group. Here, the GMM results suggest that, if anything, the more open countries are less financially developed, although these results are fragile. The R^2 for the OLS regressions is much lower in the lower-income sample than in the higher-income sample, suggesting that these regressions omit important determinants of financial depth. One possibility is that, even conditional on the level of income, countries that specialize in primary commodities are likely to have weak financial development, relative to countries where manufacturing activity has a greater role. This effect may be especially pronounced for economies dominated by point-source resources like oil or diamonds, since oil and mineral extraction is often carried out by the state or multinationals, implying less demand for external finance at any given level of income. We have explored this effect using export classifications from the GDN database, and a dummy variable for exporters of point-source resources, based on the work of Isham et al. (2002). Financial development is significantly lower for primary commodity exporters (results available on request) but controlling for this effect, or splitting the sample by export classification, does not strengthen the estimated effect of openness.

We have considered the robustness of these findings in a number of other dimensions. One simple test is to replace the trade share in current prices (*OPENC*) with the

[13] See Davidson and MacKinnon (1993, p. 237-242) for details of this approach.

trade share evaluated at prices that are constant across countries (*OPENK*). This makes little difference to the results, tending to weaken the effect of openness only slightly. We have also experimented with aggregate measures of financial development in which the individual components are not in logarithms. Again, the results are very similar to those shown in Tables 4 and 5.

In summary, these regressions indicate that external trade is associated with financial development in richer countries. These regressions have relatively high explanatory power. The trade effect is robust to treating openness as endogenous, and emerges for many different indicators of financial development. Perhaps contrary to some of the arguments in Rajan and Zingales (2003), the trade effect is stronger for bank-based finance than it is for equity-based finance. For poorer countries, the picture is a great deal more ambiguous. We find scant evidence that trade and finance are positively associated. The GMM estimates indicate that the association may even be negative, but there is also evidence that important determinants of financial development have been omitted.

5 Panel data methods

Cross-section methods are simple and easy to interpret, but have some important weaknesses. Relationships may be artificially created or obscured by unobserved heterogeneity and outliers. The use of panel data can overcome these problems, and has other advantages. We can look at relationships over time, to see whether increases in openness are followed by increases in financial depth, and to distinguish between short-run and long-run effects. In principle, we can also use the time series variation to obtain more precise estimates of the parameters of interest.

For our panel data analysis, we will use 88 countries. The data are five-year averages over the period 1960-99, giving a maximum of 8 cross-sections per country, although the panel is unbalanced. As in the empirical growth literature, averaging the data over five-year intervals means that the results are less likely to be driven by co-movements at very short horizons. Averaging is also likely to lessen the impact of measurement error, and simplifies the specification of the dynamics of the model. We will see that, even using five-year averages, we need an AR(2) rather than an AR(1) model to capture the dynamics adequately. This implies that a model based on annual data would be likely to require many parameters.[14]

To measure financial development for as many countries as possible, we drop the indicators for which data are incomplete, and construct a measure *FDEPTH*. As before,

[14]For a complementary approach that does make use of annual data, but for a smaller number of countries, see Law and Demetriades (2005).

this is the first principal component of *LLY*, *PRIVO*, and *BTOT,* but now the loadings are based on the pooled cross-section time series data for the individual measures. As before, our openness data are from version 6.1 of the Penn World Table.

We will carry out a panel data version of a Granger-causality test. Given that causality may run in either direction, we cannot treat openness as strictly exogenous. Instead, we estimate partial adjustment models that allow feedback, using sequential moment conditions to identify the model. We provide a brief introduction to this approach, which is covered in more detail in Arellano (2003, chapter 8).[15] We start with the simple AR(1) case, and then consider the AR(2) model that seems more appropriate for this application.

The most common approach in the empirical literature would be to specify an AR(1) model of the form:

$$\begin{aligned} y_{it} &= \alpha_1 y_{it-1} + \beta_1 x_{it-1} + \eta_i + \phi_t + v_{it} \qquad \mid \alpha_1 \mid < 1 \\ i &= 1, 2, ...88 \text{ and } t = 2...8 \end{aligned} \tag{2}$$

where in our application y_{it} will be a measure of financial development and x_{it} will be openness. The model allows for unobserved heterogeneity through the individual effect η_i capturing the combined effect of time-invariant omitted variables. ϕ_t is a common time effect, while v_{it} is the disturbance term. We assume that x_{it} is potentially correlated with the individual effect η_i and may be correlated with v_{it}, but is uncorrelated with future shocks $v_{it+1,}\, v_{it+2,}$ ... Under these assumptions, x_{it-1} is predetermined with respect to v_{it} and the errors can be assumed to satisfy sequential moment conditions of the form

$$E(v_{it} \mid y_i^{t-1},\ x_i^{t-1},\ \eta_i, \phi_t) = 0 \tag{3}$$

where $y_i^{t-1} = (y_{i1}, y_{i2}...., y_{i,t-1})'$ and $x_i^{t-1} = (x_{i1}, x_{i2}...., x_{i,t-1})'$.

When these moment conditions are satisfied, the transient errors v_{it} are conditionally serially uncorrelated, for any $j > 0$

$$E(v_{it}v_{it-j} \mid\ y_i^{t-1},\ x_i^{t-1},\ \eta_i, \phi_t) = 0 \tag{4}$$

and this implies (by the law of iterated expectations) that

$$E(v_{it}v_{it-j}) = 0 \tag{5}$$

[15] We draw on some aspects of Arellano's presentation, including his notation. Another useful introduction can be found in Bond (2002), and a briefer overview in Durlauf et al. (2005).

Under these assumptions, the model can be estimated by first-differencing equation (2) to eliminate the individual effects, and then using lagged levels of y_{it} and x_{it} dated $t-2$ as instruments. But a more efficient GMM estimator can be obtained by using more of the available moment conditions, as proposed by Arellano and Bond (1991). They suggest using all available lagged levels of x_{it} and y_{it} dated $t-2$ and earlier, so that the relevant moment conditions take the form

$$\begin{aligned} E(y_{it-s}\Delta v_{it}) &= 0 \qquad\qquad s \geq 2; t = 3,..,8 \\ E(x_{it-s}\Delta v_{it}) &= 0 \qquad\qquad s \geq 2; t = 3,..,8 \end{aligned} \tag{6}$$

(There are also six moment conditions associated with the period-specific constants, which we omit for simplicity.)

We call this estimator DIF-GMM. Given the strict assumptions needed for identification, it is important that specification tests are applied to the estimated models. First of all, we use the standard Arellano and Bond (1991) tests for serial correlation in the first-differenced errors. We expect to find first-order serial correlation in the differenced errors, but second-order serial correlation would imply that (4), and therefore some of the moment conditions, are invalid. We can also use a Sargan-type test to assess the model specification and overidentifying restrictions (also known in the GMM context as Hansen's J test).

A number of limitations of DIF-GMM should be noted. A well-known property of two-step GMM estimators is that the standard errors may be severely biased downwards in small samples. To address this problem, we adopt the Windmeijer (2005) finite sample correction to the standard errors. All reported standard errors and test statistics are heteroskedasticity-robust.

Importantly, when using DIF-GMM, we also experiment with restricted instrument sets. This can help to avoid the overfitting biases that are sometimes associated with using all the available (linear) moment conditions. Throughout the paper, our 'reduced' instrument sets use no lags dated further back than $t-4$. Using fewer moment conditions can also help to improve the power of Sargan-type tests; see for example Bowsher (2002).

A perhaps more fundamental weakness of DIF-GMM is that lagged levels of the series may be weak instruments for first differences, especially when the series are highly persistent, or the variance of the individual effects (η_i) is high relative to the variance of the transient shocks (v_{it}). In this case, identification of the model can sometimes be improved by making additional assumptions on the initial conditions of the process. Under assumptions developed in Arellano and Bover (1995) and Blundell and Bond (1998), the "system GMM" estimator can be used to alleviate the weak instruments problem.

This estimator (which we call SYS-GMM) adds a system of equations in levels to that in first differences. To achieve identification, the lagged first-differences of the series (y_{it}, x_{it}) dated $t-1$ are used as instruments in the untransformed (levels) equations. The additional moment conditions are

$$\begin{aligned} E[\Delta y_{it-1}(\eta_i + v_{it})] &= 0 \qquad t = 3,..,8 & (7) \\ E[\Delta x_{it-1}(\eta_i + v_{it})] &= 0 \qquad t = 3,..,8 & (8) \end{aligned}$$

Note that, given these moment conditions, differences lagged two periods or more are then redundant as instruments for the levels equations, because the corresponding moment conditions are linear combinations of those already in use.

The simulation results in Blundell and Bond (1998) suggest that the combined or system GMM estimator is more robust than difference GMM to weak instrument biases, and this method has become increasingly popular in the cross-country empirical literature. Note that the validity of the additional moment conditions used by this estimator (or a subset of them) can be tested using an incremental Sargan statistic. Also, in implementing SYS-GMM, we will again restrict the instrument set, to avoid overfitting biases. As before, in the transformed equations, we use only lagged levels at dates $t-2$, $t-3$, and $t-4$ as instruments.

Our later empirical work will suggest that the AR(1) specification is invalid for this particular application. We therefore move to an AR(2) model, with two lags of the dependent variable, and two lags of openness. This model is given by:

$$\begin{aligned} y_{it} &= \alpha_1 y_{it-1} + \alpha_2 y_{it-2} + \beta_1 x_{it-1} + \beta_2 x_{it-2} + \eta_i + \phi_t + v_{it} \\ i &= 1, 2, ...88 \text{ and } t = 3...8 & (9) \end{aligned}$$

and so the first-differenced equation is:

$$\begin{aligned} \Delta y_{it} &= \alpha_1 \Delta y_{it-1} + \alpha_2 \Delta y_{it-2} + \beta_1 \Delta x_{it-1} + \beta_2 \Delta x_{it-2} + \phi_t - \phi_{t-1} + \Delta v_{it} \\ i &= 1, 2, ...88 \text{ and } t = 4...8 & (10) \end{aligned}$$

Now the relevant moment conditions for DIF-GMM are:

$$\begin{aligned} E(y_{it-s}\Delta v_{it}) &= 0 \qquad s \geq 2; t = 4,..,8 & (11) \\ E(x_{it-s}\Delta v_{it}) &= 0 \qquad s \geq 2; t = 4,..,8 \end{aligned}$$

Note that, perhaps surprisingly, we can continue to use moment conditions based on the lagged levels y_{it-2} and x_{it-2} for the equations in first differences, even when y_{it-2} and x_{it-2} are included in the untransformed model. Arellano (2003, section 6.7) is an example of this approach. To see how this works, it can alternatively be interpreted as exploiting the exogeneity of Δy_{it-2} and Δx_{it-2} in the first-differenced equations, together with the use of lagged levels dated $t-3$ and earlier:

$$\begin{aligned}
E(\Delta y_{it-2}\Delta v_{it}) &= 0 \qquad\qquad t = 4,..,8 \\
E(\Delta x_{it-2}\Delta v_{it}) &= 0 \qquad\qquad t = 4,..,8 \\
E(y_{it-s}\Delta v_{it}) &= 0 \qquad\qquad s \geq 3; t = 4,..,8 \\
E(x_{it-s}\Delta v_{it}) &= 0 \qquad\qquad s \geq 3; t = 4,..,8
\end{aligned}$$

But clearly these moment conditions are linear combinations of the set (11). This alternative, equivalent way of writing down the moment conditions helps to clarify that, in the AR(2) model, there is a sense in which identification relies on lagged levels dated $t-3$ and earlier. (This can be seen more explicitly by considering how a 2SLS approach would be implemented.) Since these longer lags may be weak instruments, this again suggests the potential usefulness of the system GMM estimator in this context.

In the AR(2) model, one hypothesis of economic interest is the joint null $\beta_1 = \beta_2 = 0$, which can be interpreted as a panel data test for Granger-causality. Although a Wald-type test of this restriction could be implemented, we use an alternative approach. This is to estimate both the unrestricted and the restricted models using the same moment conditions, and compare their (two-step) Sargan statistics using an incremental Sargan test of the form:

$$D_{RU} = n(J(\tilde{\gamma}) - J(\hat{\gamma}))$$

where $J(\tilde{\gamma})$ is the minimized GMM criterion for the restricted model, $J(\hat{\gamma})$ for the unrestricted model, and n is the number of observations. Under the null, D_{RU} is asymptotically distributed as χ^2_r where r is the number of restrictions. The intuition for the test is that, if the parameter restrictions are valid, the moment conditions should remain valid even in the restricted model. For more details, see Bond et al. (2001a) and Bond and Windmeijer (2005).

We now turn to some additional issues of interpretation raised by the use of an AR(2) model. First of all, we may be interested in the stability of the estimated model. For stability we require the roots of the relevant lag polynomial, namely the bracketed term on the left-hand side of:

$$\left(1 - \alpha_1 L - \alpha_2 L^2\right) y_{it} = \beta_1 x_{it-1} + \beta_2 x_{it-2} + \eta_i + \phi_t + v_{it}$$

to lie outside the unit circle. This can easily be checked by either dynamic simulation or direct calculation; see for example Hamilton (1994, p. 17-18) on the latter. The majority of our estimated models are stable, the main exceptions arising when pooled OLS is used rather than our preferred fixed-effects methods.

If the model is stable, we can calculate a point estimate for the long-run effect of openness on financial development:

$$\beta_{LR} = \frac{\beta_1 + \beta_2}{(1 - \alpha_1 - \alpha_2)}$$

We can estimate an approximate standard error for this long-run effect using the delta method. However, we should note that the long-run effect, as a nonlinear function of the model parameters, may be imprecisely estimated. In particular, if the sum $\alpha_1+\alpha_2$ is close to one and imprecisely estimated, the data can appear consistent with very high values for the long-run effect, because the ratio quickly blows up as $\alpha_1+\alpha_2$ nears unity. Since the confidence interval generated by the delta method is, perhaps mistakenly, assumed to be symmetric around the point estimate, it may well overlap zero too often.

For this reason, we complement the delta method with a test of the restriction $\beta_1 + \beta_2 = 0$. If this restriction is rejected, it suggests that there is evidence for a long-run effect of openness on financial development. If the restriction holds and the parameters are non-zero ($\beta_2 = -\beta_1 \neq 0$) then financial development depends on the change in openness, and not on its level. This would be consistent with a story in which increases in openness, through restructuring and investment, lead to a short-term boom in lending that does not persist into the long run. Again, we will test this restriction on the coefficients using a criterion-based approach.

In some of our estimated models, the lag polynomial associated with the dependent variable has one root close to unity, indicating a high degree of persistence. This makes it especially relevant to ask how the estimation methods will perform when the data-generating process is characterized by a high degree of persistence. The GMM estimators will remain consistent, because the relevant asymptotics here are for large N, fixed T. A qualification is that in the case of an exact unit root, the moment conditions used in DIF-GMM are no longer sufficient to identify the parameters. Identification may require the use of mean stationarity assumptions, as in the system GMM estimator; see Arellano (2003, p. 116) for the AR(1) case.

Finally, we can test for unobserved heterogeneity using a procedure originally suggested by Holtz-Eakin (1988). In the absence of individual effects, the following additional moment conditions become valid, corresponding to the use of lagged levels as instruments in the levels equations:

$$\begin{aligned} E[y_{it-1}(y_{it} - \alpha_1 y_{it-1} - \alpha_2 y_{it-2} - \beta_1 x_{it-1} - \beta_2 x_{it-2} - \phi_t\ )] &= 0 \\ E[x_{it-1}(y_{it} - \alpha_1 y_{it-1} - \alpha_2 y_{it-2} - \beta_1 x_{it-1} - \beta_2 x_{it-2} - \phi_t\ )] &= 0 \\ t &= 3, .., 8 \end{aligned} \tag{12}$$

The validity of these additional moment conditions, which can be tested using an incremental Sargan test relative to difference or system GMM, then becomes a simple test for the presence of unobserved heterogeneity (where the null is no heterogeneity). One motivation for using this test is that, if individual effects are not present, pooled OLS will be a consistent estimator, although not fully efficient given the likely presence of heteroskedasticity.

6 Panel data results

This section presents the results of our panel data analysis. The main findings can be summarized as follows. In the short run (here, 5-10 years) increases in openness are followed by increases in financial depth. For the whole sample, and a sample of lower-income countries, this result is robust across different estimation methods, and to variation in the choice of moment conditions. Not surprisingly, the long-run effect is estimated less precisely, but in most cases we can reject the restriction that $\beta_1 + \beta_2 = 0$. This suggests that the effect of openness on financial development persists into the long run, although the extent of support for this hypothesis varies with the precise choice of moment conditions.

We will consider OLS estimates, Within Groups (WG) estimates, and two versions of difference GMM. All models include a full set of time dummies. The first version of DIF-GMM uses all available linear moment conditions, while the second does not use any lags dated further back than $t - 4$. We will also consider three versions of system GMM, the first using both sets of moment conditions (7) and (8). The second two versions use separately either (7), which we call SYS-GMM-1, or (8), which we call SYS-GMM-2. This approach helps to avoid overfitting and also reflects the possibility that the system GMM assumptions may be incorrect. For example, if the countries with relatively small individual effects (and hence tending to be less developed financially) are also the countries in which openness has increased the fastest, the moment conditions in (8) would be invalid.

First of all, we consider an AR(1) model of the form often studied in the cross-country empirical literature. These results are shown in Table 6.[16] The table provides several reasons to believe this model is badly mis-specified. One source for concern is the evidence for second-order serial correlation. In the case of system GMM, in the final column, the Sargan statistic also suggests that either the model specification or the moment conditions are invalid.[17]

[16] All our panel data estimates are obtained using the `xtabond2` package for `Stata`, written by David Roodman. We were able to obtain almost identical results and test statistics using the `DPD for PcGive` software described in Doornik et al. (2002).

[17] More subtly, it is worth noting that the autoregressive parameter estimated by DIF-GMM or SYS-

With these problems in mind, we do not give further consideration to AR(1) models. We turn instead to AR(2) models of the form:

$$\begin{aligned} y_{it} &= \alpha_1 y_{it-1} + \alpha_2 y_{it-2} + \beta_1 x_{it-1} + \beta_2 x_{it-2} + \eta_i + \phi_t + v_{it} \\ i &= 1, 2, ...88 \text{ and } t = 3...8 \end{aligned}$$

We will be especially interested in whether the level of financial development depends primarily on the *change* in openness, or the level. To examine this, we will test the restriction that $\beta_1 + \beta_2 = 0$, in which case $\beta_1 x_{it-1} + \beta_2 x_{it-2} = \beta_1 \Delta x_{it-1}$. We will see that the degree of support for this interpretation is limited. It is always strongest when pooled OLS is used, typically less strong under system GMM, and weakest under difference GMM. In general, when we allow for fixed effects, we find evidence that the effects of openness persist into the long run.

Table 7 presents results for the whole sample of 88 countries.[18] Recall that, whenever we apply the GMM estimators, we report tests for serial correlation (first- and second-order), the Sargan statistic, a test for Granger causality ($\beta_1 = \beta_2 = 0$), a test for a long-run effect ($\beta_1 + \beta_2 = 0$) and the implied point estimate and approximate standard error of the long-run effect. We report the tests of restrictions on parameters, and long-run effects, for OLS and Within Groups also. In the OLS and WG cases, the tests of restrictions are based on conventional Wald tests.

Looking at Table 7, the first point to note is that, across all the estimation methods, we find strong evidence that increases in openness are associated with increases in financial depth in the short-run. The coefficient on the first lag of openness is positively signed and significantly different from zero across all seven models reported in Table 7. The effect of the second lag of openness varies more across the models. In the system GMM estimates, especially, we find evidence that the coefficient on the second lag of openness is negatively signed, which points to the importance of specifying the dynamics carefully. The Granger-causality test rejects the null of non-causality at the 5% level for all seven models.

The long-run effect is less clear. In the final two rows of the Table, we report the point estimate and approximate standard error associated with the long-run effect. In the case of the WG estimates, and the two varieties of DIF-GMM estimates, we find a

GMM does not lie in the interval defined by the OLS and WG estimates. Since the OLS and WG estimates of the autoregressive parameter should be biased in opposite directions, a consistent estimate of this parameter might be expected to lie between these two extremes (see Bond et al. 2001b for more discussion). Again, this tends to call into question the validity of the moment conditions, and may also hint at weak instrument biases.

[18]Note that, following our earlier discussion, it may again be tempting to compare the first autoregressive parameter in the GMM estimates with that obtained under WG and OLS; but this would not have the same justification in the AR(2) model as in the AR(1) model.

stable long-run effect, and one that is significantly different from zero at the 5% level in columns 2 and 4. When we turn to the different versions of SYS-GMM (columns 5-7) the long-run effect is less precisely estimated.[19] Note, however, that whenever we allow for fixed effects, the restriction that $\beta_1 + \beta_2 = 0$ is rejected at the 20% level, and typically at the 5% level. This suggests that the effect of greater openness persists into the long run.

Our simple heterogeneity tests reject the null of no individual effects in column 7, and are close to doing so in the other two cases. This suggests that, as expected, unobserved heterogeneity is present, and the pooled OLS estimates are likely to be inconsistent. An alternative way to address this problem is to look for samples that are more likely to be homogenous. For this reason, we follow our earlier cross-section analysis, and split the sample into a higher-income group (Table 8) and a lower-income group (Table 9).

We consider the higher-income group first, in Table 8. In this group of countries, the evidence for a significant short-run effect is much weaker than before. We typically cannot reject the null hypothesis that the two lags of openness are jointly insignificant, or the alternative null that $\beta_1 + \beta_2 = 0$. Consistent with this pattern, the long-run effects are imprecisely estimated for all estimation methods other than Within Groups.

At first glance, these findings might seem to go against our earlier OLS and GMM cross-section results, in which openness and financial development were positively associated in the higher-income sample. One reason for weaker results in the higher-income panel may be that its cross-section dimension ($N = 35$) is unusually small for an application of GMM. It is interesting to note that our test for individual effects does not reject the null of a common intercept at conventional levels. If taken at face value, the lack of variation in the individual effects in the higher-income group implies that panel data methods are not required. Simple cross-section or pooled OLS estimates may provide a reasonable estimate of the long-run effect, and have the advantage that they retain the information in the "between" variation.

The results for the lower-income group are shown in Table 9. In this sample, we find much stronger evidence that increased openness leads to greater financial depth in the short run. This effect is significant at the 10% level in all seven models. Again we find evidence that the coefficient on the second lag of openness is negatively signed, and can see that it is hard to estimate the long-run effect precisely in a sample of this size, with the exception of the SYS-GMM-2 estimates in column 7. But the restriction that $\beta_1 + \beta_2 = 0$ is usually either rejected, or is close to being rejected.

The diagnostic tests support the identifying assumptions, and overall we conclude

[19] This arises partly because of the significantly negative coefficient on the second lag of openness (corresponding to β_2), which makes a zero long-run effect harder to reject, and partly because $\alpha_1 + \alpha_2 \approx 1$ in these models, which tends to blow up the ratio of coefficients that forms the long-run effect.

that there is strong evidence that changes in openness are followed by changes in financial depth, which persist at least over the medium term. In the whole sample, or the subset of poorer countries, there is fairly strong evidence that the effect persists into the long run. In the group of higher-income countries, the estimates are less precise, which may reflect the small cross-section dimension of the relevant panel. An alternative explanation is that the stronger cross-section results for the higher-income sample are an artifact of individual effects, but we find relatively little statistical evidence for these individual effects in the higher-income panel.

7 Conclusions

The determinants of financial development have become a focus of recent research. In this paper, we have examined whether finance is influenced by external trade, perhaps for the political economy reasons identified by Rajan and Zingales (2003), or the risk diversification considerations emphasized by Svaleryd and Vlachos (2002). The first sections of the paper emphasize cross-section results, in which the Frankel-Romer measure of "natural openness" is used to isolate exogenous variation in openness. Whether using OLS or instrumental variable procedures, we find strong evidence that trade promotes bank-based financial development in higher-income countries, but not in the lower-income group.

These findings may be contaminated by omitted variables that are correlated with openness. The main contribution of the paper is to use panel data to examine whether increases in openness are followed by increases in financial development, and whether this effect persists into the long run. We find strong support for this hypothesis in the sample as a whole, and in the lower-income group. The results are robust to several different estimation methods, and the short-run effects in particular are not sensitive to the precise choice of moment conditions. For the higher-income group, the panel data results are weaker than in the cross-section.

Our results suggest that the long-run effects of trade go beyond those envisaged in traditional models. But as with much of the empirical literature on financial development, it is not clear whether the effects work primarily through the demand for external finance, or through improving the supply side, or some interaction of the two. For example, it may be that greater openness is associated with changes in sectoral structure that increase the demand for external finance. Alternatively, it may be that increased exposure to foreign competition influences the supply-side more directly, as envisaged in Rajan and Zingales (2003). Discriminating between the various explanations is a difficult task, but may be an interesting area for further research, given the findings we present here.

References

[1] Acemoglu, D. and Zilibotti, F. (1999). “Information accumulation in development,” *Journal of Economic Growth*, 4(1), 5-38.

[2] Aizenman, J. (2004). “Financial opening and development: evidence and policy controversies,” *American Economic Review*, 94(2), 65-70.

[3] Aizenman, J. and Noy, I. (2003). “Endogenous financial openness: efficiency and political economy considerations,” NBER working paper no. 10144.

[4] Alessandria, G. and J. Qian (2005), “Endogenous financial intermediation and real effects of capital account liberalization,” *Journal of International Economics*, forthcoming.

[5] Arellano, M. and S. R. Bond (1991), “Some tests of specification for panel data: Monte carlo evidence and an application to employment equations”, *Review of Economic Studies*, 58: 277-297.

[6] Arellano, M. (2003). *Panel data econometrics*. Oxford University Press, Oxford.

[7] Beck, T. (2002), “Financial development and international trade. Is there a link?”, *Journal of International Economics*, 57: 107-131

[8] Beck, T. (2003), “Financial dependence and international trade”, *Review of International Economics*, 11: 296-316.

[9] Beck, T., A. Demirguc-Kunt, and R. Levine (2000). “A new database on the structure and development of the financial sector,” *World Bank Economic Review*, 14(3), 597-605.

[10] Beck, T., Demirguc-Kunt, A. and Levine, R. (2003). “Law, endowments, and finance,” *Journal of Financial Economics*, 70, 137-181.

[11] Berkowitz, D., K. Pistor and J.-F. Richard (2003), "Economic development, legality, and the transplant effect", *European Economic Review*, 47(1), 165-95.

[12] Blundell, R. and S. Bond (1998), ”Initial conditions and moment restrictions in dynamic panel data models”, *Journal of Econometrics*, 87: 115-43.

[13] Bond, S. R. (2002), ”Dynamic panel data models: a guide to micro data methods and practice”, *Portuguese Economic Journal*, 1: 141-162.

[14] Bond, S. R., C. G. Bowsher, and F. Windmeijer (2001a), ”Criterion-based inference for GMM in autoregressive panel data models”, *Economics Letters*, 73: 379-88.

[15] Bond, S. R., A. Hoeffler and J. R. W. Temple (2001b), "GMM estimation of empirical growth models", CEPR discussion paper no. 3048.

[16] Bond, S. R. and Windmeijer, F. (2005). “Reliable inference for GMM estimators? Finite sample properties of alternative test procedures in linear panel data models,” *Econometric Reviews*, 24(1), 1-37.

[17] Bowsher, C. G. (2002). “On testing overidentifying restrictions in dynamic panel data models”, *Economics Letters* 77, 211-220.

[18] Boyd, J. H., R. Levine and Bruce D. Smith (2001), “The impact of inflation on financial sector performance”, *Journal of Monetary Economics*, 47, 221-48.

[19] Chinn, M. D. and Ito, H. (2005). “What matters for financial development? Capital controls, institutions, and interactions,” NBER working paper no. 11370.

[20] Davidson, R. and J. G. MacKinnon (1993), *Estimation and inference in econometrics.* Oxford University Press.

[21] Dell’Ariccia, G. and Marquez, R. (2005). “Lending booms and lending standards,” CEPR discussion paper no. 5095.

[22] Demirguc-Kunt, A. and R. Levine (1999). “Bank-based and market-based financial systems: Cross-country comparisons”, World Bank Policy Research Working Paper No. 2143.

[23] Do, Q.-T. and Levchenko, A. A. (2004). “Trade and financial development,” World Bank working paper series 3347, April.

[24] Dollar, D. and Kraay, A. (2004). “Trade, growth, and poverty,” *Economic Journal*, 114, F22-F49.

[25] Doornik, J., Arellano, M. and Bond, S. (2002). “Panel data estimation using DPD for OX,” Manuscript, Nuffield College, Oxford.

[26] Durlauf, S. N., P. A. Johnson and J. R. W. Temple (2005) “Growth econometrics”, in P. Aghion and S. N. Durlauf (eds) *Handbook of Economic Growth*, North-Holland, forthcoming.

[27] Frankel, J. A. and D. Romer (1999), “Does trade cause growth?”, *American Economic Review*, 89: 379-98.

[28] Giavazzi, F. and Tabellini, G. (2004), "Economic and political liberalizations," CEPR discussion paper no. 4579.

[29] Hamilton, J. D. (1994). *Time series analysis*. Princeton University Press, Princeton, New Jersey.

[30] Henry, P. B. (2003). "Capital-account liberalization, the cost of capital, and economic growth," *American Economic Review*, 93(2), 91-96.

[31] Heston, A., Summers, R. and Aten, B. (2002). *Penn World Table Version 6.1*, Center for International Comparisons at the University of Pennsylvania (CICUP), October.

[32] Holtz-Eakin, D. (1988). "Testing for individual effects in autoregressive models," *Journal of Econometrics*, 39(3), 297-307.

[33] Huybens, E. and B.D. Smith (1999), "Inflation, financial markets and long-run real activity", *Journal of Monetary Economics*, 43: 283-315.

[34] Isham, J., Woolcock, M., Pritchett, L. and Busby, G. (2002). "The varieties of rentier experience: how natural resource export structures affect the political economy of economic growth," Manuscript, March.

[35] Ju, J. and Wei, S.-J. (2005). "Endowment versus finance: a wooden barrel theory of international trade," CEPR discussion paper no. 5109.

[36] King, R. G. and R. Levine (1993). "Finance and growth: Schumpeter might be right", *Quarterly Journal of Economics*, 108: 717-737.

[37] Kletzer, K. and Bardhan, P. K. (1987). "Credit markets and patterns of international trade," *Journal of Development Economics*, October.

[38] La Porta, R., Lopez-de-Silanes, F., Shleifer, A., and Vishny, R.W. (1998), "Law and finance", *Journal of Political Economy*, 106: 1113-1155.

[39] Law, S. H. and Demetriades, P. (2005). "Openness, institutions and financial development," University of Leicester, working paper no. 05/08.

[40] Levine, R. (1997). "Financial development and economic growth: views and agenda". *Journal of Economic Literature*, 35: 688-726.

[41] Levine, R. (2005). "Finance and growth: theory and evidence," in P. Aghion and S. N. Durlauf (eds) *Handbook of Economic Growth*, North-Holland, forthcoming.

[42] Levine, R., N. Loayza and T. Beck (2000). "Financial intermediation and growth: causality and causes," *Journal of Monetary Economics*, 46(1): 31-77.

[43] Levine, R. and Renelt, D. (1992). "A sensitivity analysis of cross-country growth regressions," *American Economic Review*, 82, 942-963.

[44] Levine, R. and S. Zervos (1998), "Stock markets, banks and economic growth", *American Economic Review*, 88: 537-58.

[45] Li, K., R. Morck, F. Yang and B. Yeung (2004). "Firm-specific variation and openness in emerging markets," *Review of Economics and Statistics*, 86(3), 658-669.

[46] Mayer, C. and O. Sussman (2001), "The assessment: finance, law and growth", *Oxford Review of Economic Policy*, 17(4), 457-66.

[47] McKinnon, R. I. (1973). *Money and capital in economc development*. Washington D.C.: Brookings Institution.

[48] Pagano, M. and Volpin, P. (2001a). "The political economy of finance," *Oxford Review of Economic Policy*, 17(4), 502-519.

[49] Pagano, M. and Volpin, P. (2001b). "The political economy of corporate governance," CEPR discussion paper no. 2682.

[50] Rajan, R. G. and L. Zingales (1998). "Financial dependence and growth," *American Economic Review*, 88: 559-86.

[51] Rajan, R. G. and Zingales, L. (2003). "The great reversals: the politics of financial development in the twentieth century," *Journal of Financial Economies*, 69, 5-50.

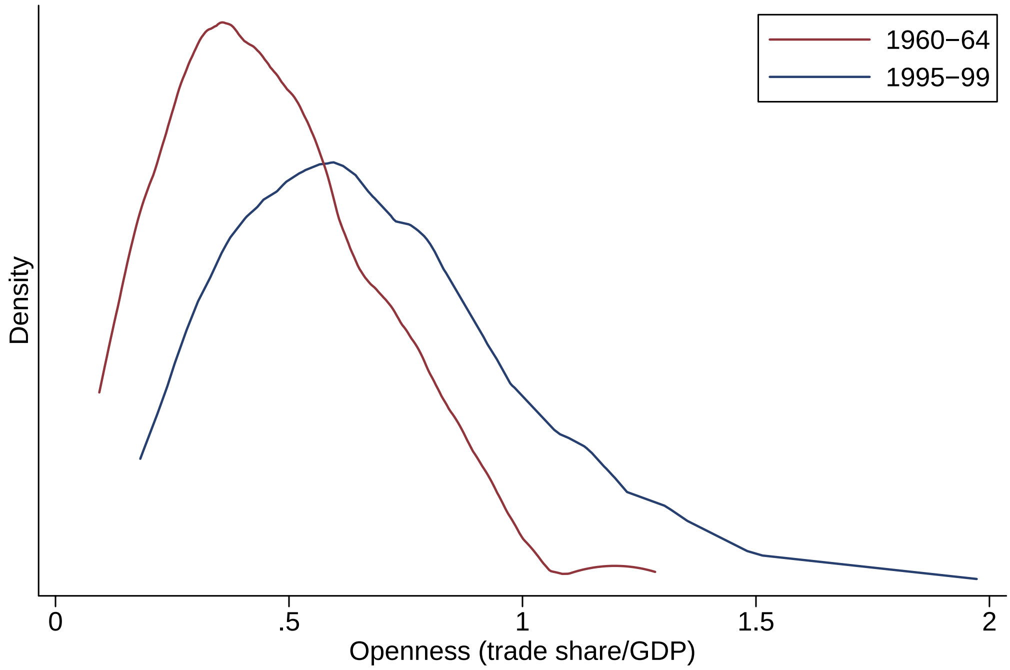
[52] Stulz, R. M. and Williamson, R. (2003). "Culture, openness, and finance," *Journal of Financial Economics*, 70, 313-349.

[53] Svaleryd, H. and Vlachos, J. (2002), "Markets for risk and openness to trade: how are they related?" *Journal of International Economics*, 57: 369-395.

[54] Windmeijer, F. (2005). "A finite sample correction for the variance of linear efficient two-step GMM estimators," *Journal of Econometrics*, 126(1), 25-51.

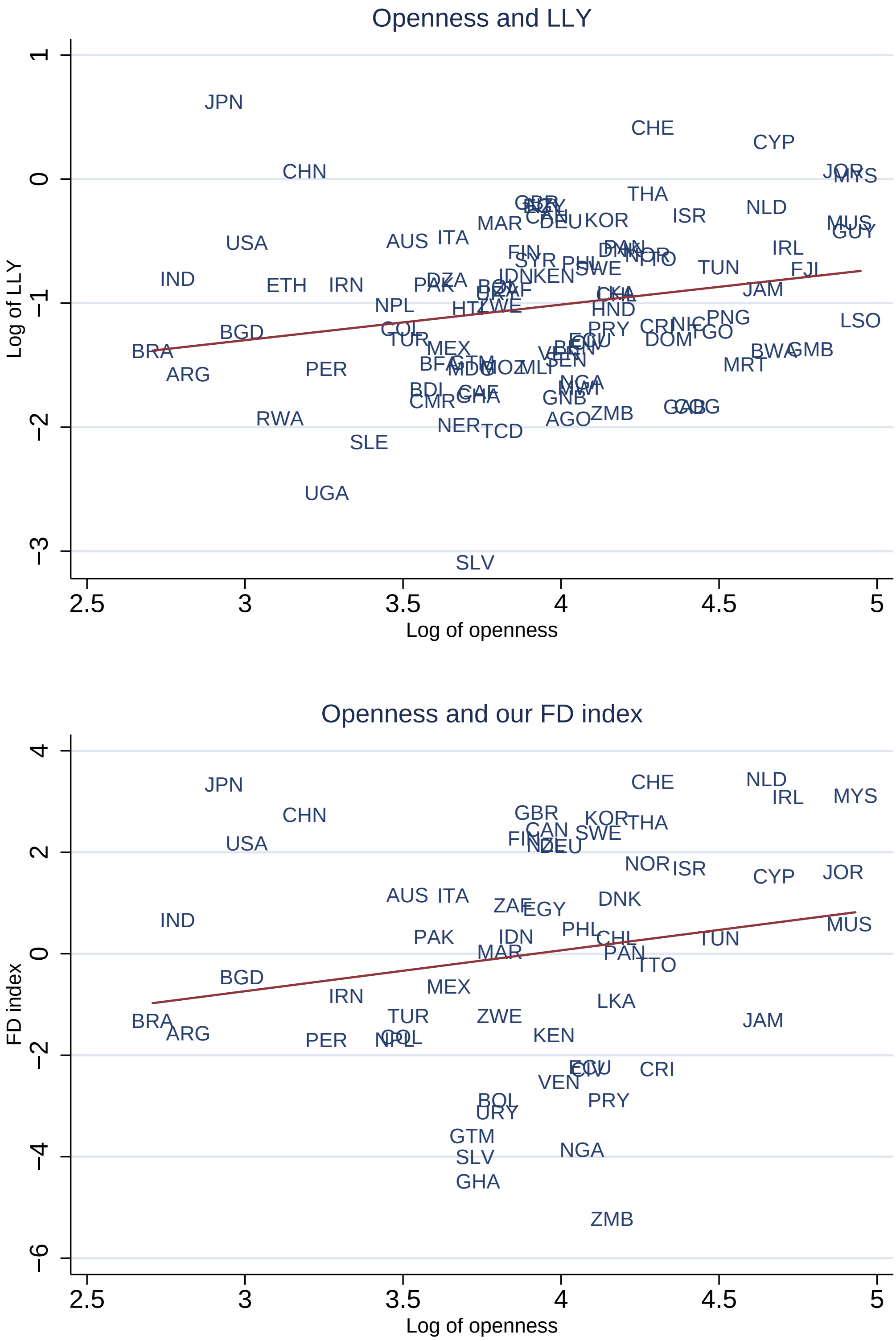
[55] Wynne, J. (2005). "Wealth as a determinant of comparative advantage," *American Economic Review*, 95(1), 226-254.

Figure 1: The increase in openness



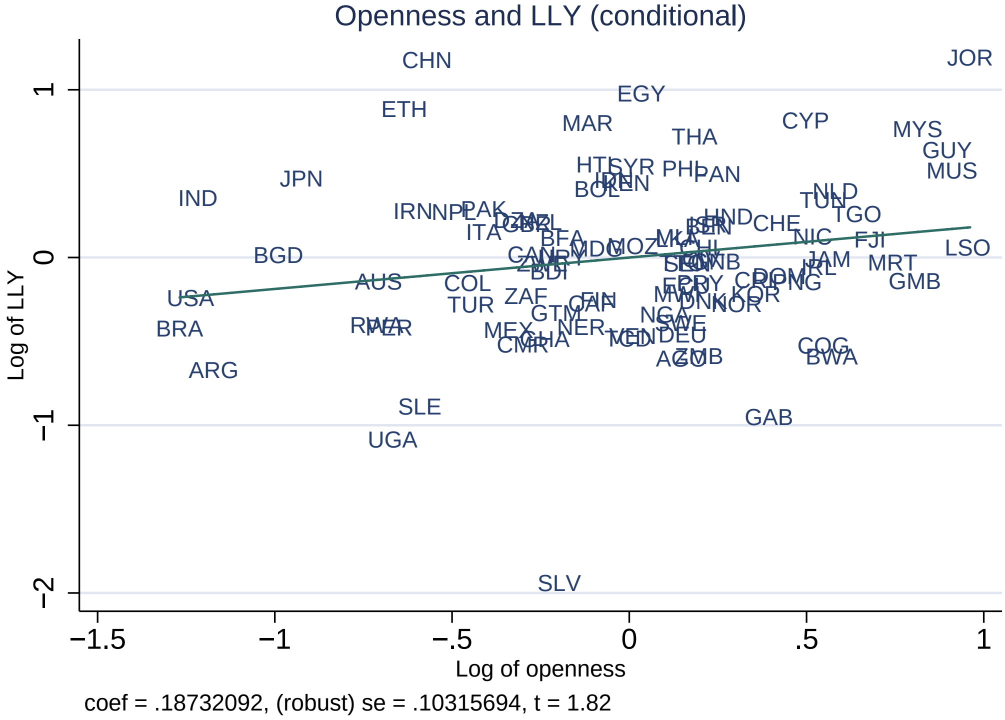
This figure shows kernel density plots of the distributions of average openness in 1960-64 (left) and 1995-99 (right). The figure shows the trend towards increased openness over this period.

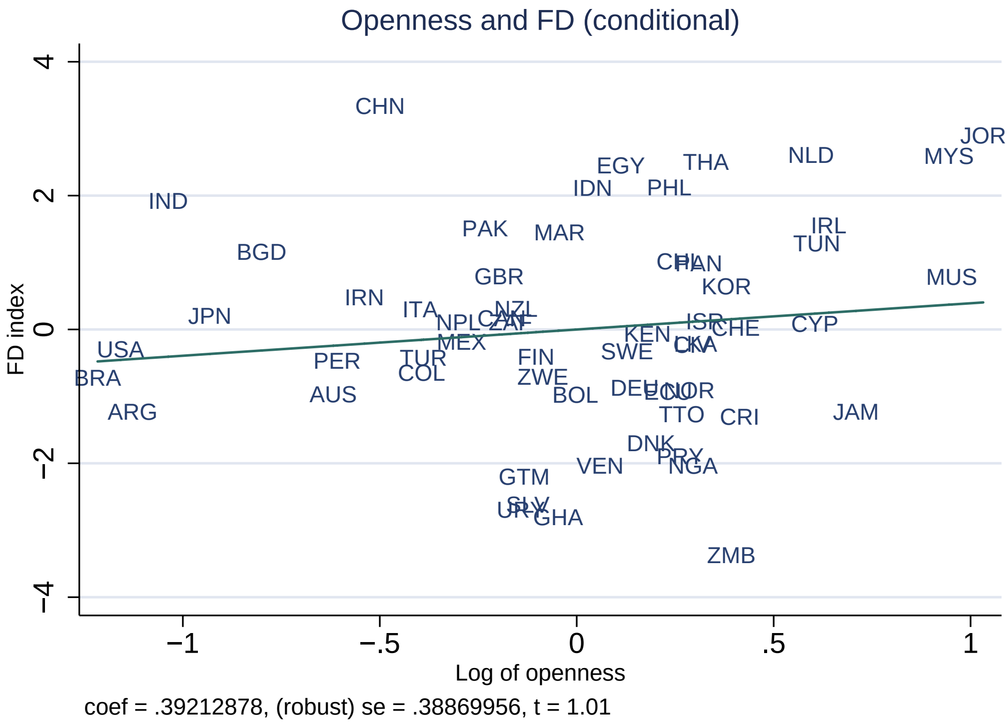
Figure 2: Trade and finance



These figures show scatter plots of the logarithm of LLY, and the aggregate index FD, against the logarithm of openness.

Figure 3: Partial scatter plots





These figures show partial scatter plots of the logarithm of LLY, and the aggregate index FD, against the logarithm of openness. These are the partial associations between trade and finance, conditional on the logarithm of GDP and legal origin dummies.

Table 1. The Variables

Variable	Description	Sources
OPENC	The sum of exports and imports over GDP (at current prices)	Penn World Table 6.1
OPENK	The sum of exports and imports over GDP (at international prices)	Penn World Table 6.1
RGDPCH	Average real GDP per capita over 1988-90 (cross-section study only)	Penn World Table 6.1
CTRADE	Natural propensity to trade, as derived by Frankel and Romer. Based on aggregated fitted values of bilateral trade equation.	Frankel and Romer (1999)
LLY	Ratio of liquid liabilities of financial system (currency plus demand and interest-bearing liabilities of banks and nonbanks) to GDP	FSD
PRIVO	Ratio of credit issued to private sector by banks and other financial intermediaries, to GDP	FSD
BTOT	Ratio of commercial bank assets to sum of commercial and central bank assets	FSD
OVC	Ratio of overhead costs to total assets of the banks	FSD
NIM	Bank interest income minus interest expenses over total assets	FSD
MCAP	Ratio of the value of listed shares to GDP	FSD
TVT	Ratio of the value of shares traded on domestic exchanges to GDP	FSD
TOR	Ratio of the value of shares traded to market capitalization	FSD
FD	First principal component of LLY, PRIVO, BTOT, NIM, MCAP, TVT, TOR (all in logs)	See text
FDSIZE	First principal component of LLY and MCAP (all in logs)	See text
FDEFF	First principal component of OVC, NIM, TVT and TOR (all in logs)	See text
FDBANK	First principal component of LLY, PRIVO, BTOT, OVC and NIM (all in logs)	See text
FDSTOCK	First principal component of MCAP, TVT and TOR (all in logs)	See text
FDEPTH	First principal component of LLY, PRIVO and BTOT. This index is used in the panel data analysis.	See text
LEGOR_UK	British legal origin	GDN
LEGOR_FR	French legal origin	GDN
LEGOR_GE	German legal origin	GDN
LEGOR_SC	Scandinavian legal origin	GDN

Key to Table 1: FSD – Financial Structure Database introduced by Beck, Demirguc-Kunt and Levine (2000). GDN – Global Development Network Database, 2002

Table 2. The indices of financial development

Measure	Proportion	LLY	PRIVO	BTOT	OVC	NIM	MCAP	TVT	TOR
FD	0.63	0.38	0.40	0.30	-0.32	-0.36	0.36	0.38	0.32
FDSIZE	0.80	0.71					0.71		
FDEFF	0.68				-0.48	-0.51		0.52	0.49
FDBANK	0.71	0.48	0.48	0.38	-0.43	-0.46			
FDSTOCK	0.86						0.55	0.62	0.56
FDEPTH	0.74	0.60	0.63	0.49					

Notes: This table shows how our various indicators of financial development (FD, FDSIZE, FDEFF, FDBANK, FDSTOCK, FDEPTH) are constructed, from the raw data on different measures of financial development. We construct indicators from the raw data using the first principal component of a number of variables, namely the linear combination of the variables that has the highest sample variance, subject to the constraint that the sum-of-squares of the coefficients equals unity. The table shows the weights that each index places on each of the (standardized) variables, and the proportion of the variance in the original data that is explained by the first principal component.

The raw measures used are the natural logarithms of LLY = the ratio of liquid liabilities of the financial system (currency plus demand and interest-bearing liabilities of banks and nonbanks) to GDP; PRIVO = the ratio of credit issued to the private sector by banks and other financial intermediaries to GDP; BTOT= the ratio of commercial bank assets to the sum of commercial bank and central bank assets; OVC = the ratio of overhead costs to total assets of the banks; NIM = the bank interest income minus interest expenses over total assets; MCAP = the ratio of the value of shares listed on domestic exchanges (market capitalization) to GDP; TVT = the ratio of the value of shares traded on domestic exchanges to GDP; TOR = the ratio of the value of shares traded on domestic exchanges to total market capitalization.

Table 3. Descriptive Statistics: 1990-2001

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A. Summary Statistics for Openness and financial development measures

Variable	Observation	Mean	Std. Dev.	Min	Max
OPENC	101	59.73	30.71	14.99	140.90
LLY	93	0.44	0.31	0.05	1.87
PRIVO	99	0.42	0.38	0.02	1.65
BTOT	98	0.77	0.21	0.17	1.00
OVC	92	0.05	0.02	0.01	0.11
NIM	91	0.06	0.03	0.01	0.17
MCAP	67	0.36	0.37	0.01	1.76
TVT	68	0.21	0.29	0.00	1.34
TOR	67	0.42	0.42	0.01	1.92
FD	59	0.00	2.25	-5.22	3.45
FDSIZE	61	0.00	1.26	-3.06	2.58
FDEFF	66	0.00	1.64	-3.44	3.16
FDBANK	82	0.00	1.88	-4.80	3.51
FDSTOCK	67	0.00	1.60	-3.99	2.41

B. Correlations between openness and the existing financial development measures

	OPENC	LLY	PRIVO	BTOT	OVC	NIM	MCAP	TVT	TOR
OPENC	1.00								
LLY	0.22	1.00							
PRIVO	0.18	0.86	1.00						
BTOT	0.15	0.50	0.64	1.00					
OVC	-0.21	-0.56	-0.54	-0.46	1.00				
NIM	-0.15	-0.57	-0.59	-0.51	0.86	1.00			
MCAP	0.24	0.58	0.71	0.41	-0.35	-0.40	1.00		
TVT	0.04	0.54	0.77	0.44	-0.32	-0.37	0.79	1.00	
TOR	-0.23	0.32	0.47	0.38	-0.29	-0.25	0.23	0.60	1.00

C. Correlations among openness and the new financial development measures

	OPENC	FD	FDSIZE	FDEFF	FDBANK	FDSTOCK
OPENC	1.00					
FD	0.27	1.00				
FDSIZE	0.30	0.92	1.00			
FDEFF	0.15	0.95	0.80	1.00		
FDBANK	0.27	0.94	0.84	0.87	1.00	
FDSTOCK	0.03	0.86	0.82	0.86	0.63	1.00

Table 4. External trade and financial development (whole sample), 1990-2001

	FD	FDSIZE	FDEFF	FDBANK	FDSTOCK
OPENC (OLS)	0.64	0.244	0.516	0.56	-0.497
	(0.08)*	-0.25	(0.04)**	(0.07)*	(0.06)*
OPENC (GMM)	0.138	0.139	0.238	0.532	-0.883
	(0.78)	(0.48)	(0.49)	(0.15)	(0.01)***
(a) R-squared	0.59	0.60	0.57	0.52	0.52
(b) First-stage F[1]	69.07	75.75	68.55	68.60	87.56
(c) Pagan-Hall[2]	0.32	0.13	0.14	0.27	0.10
(d) Wu-Hausman[3]	0.16	0.49	0.29	0.92	0.06
(e) C statistic[4]	0.17	0.47	0.28	0.91	0.08
Observations	57	56	62	79	64

Notes: This Table shows the point estimates and p-values for openness in OLS and GMM estimates, using five alternative measures of financial development as the dependent variable. The coefficients and heteroskedasticity-robust p values correspond to the natural logarithm of openness. Table 1 describes all variables in detail. Other explanatory variables included in each of the regressions are the natural logarithm of initial real GDP, and one or more dummy variables for legal origin. In the GMM estimates, the instrument is the Frankel-Romer measure of the propensity to trade (CTRADE). All regressions exclude outliers, as identified by a preliminary median regression (see text). Most of the diagnostics are based on 2SLS rather than GMM.

significant at 10%; ** significant at 5%; *** significant at 1%

[1] This tests the significance of the excluded instrument (CTRADE) in the first-stage regression of 2SLS.

[2] This tests the homoskedasticity of the system of equations when 2SLS is used. Based on Pagan and Hall (1983).

[3] This is a Wu-Hausman test that the difference between the OLS and 2SLS coefficients is not systematic.

[4] This is a GMM-based test of the null hypothesis that OPENC is orthogonal to the disturbances.

Table 5. External trade and financial development (subsamples), 1990-2001

A. Higher-income group

	FD	FDSIZE	FDEFF	FDBANK	FDSTOCK
OPENC (OLS)	0.88	0.472	0.442	1.563	-0.371
	(0.00)***	(0.01)***	(0.08)*	(0.00)***	(0.09)*
OPENC (GMM)	0.831	0.514	0.264	1.254	-0.431
	(0.00)***	(0.00)***	(0.27)	(0.00)***	(0.03)**
(a) R-squared	0.77	0.72	0.55	0.84	0.58
(b) First stage F	46.99	44.95	46.46	30.98	55.16
(c) Pagan-Hall	0.47	0.89	0.07	0.09	0.48
(d) Wu-Hausman	0.78	0.78	0.33	0.05	0.63
(e) C statistic	0.74	0.73	0.27	0.05	0.59
Observations	27	28	34	26	34

B. Lower-income group

	FD	FDSIZE	FDEFF	FDBANK	FDSTOCK
OPENC (OLS)	-1.023	-0.435	-1.565	-0.221	-0.753
	(0.13)	(0.23)	(0.01)***	(0.63)	(0.19)
OPENC (GMM)	-4.027	-1.357	-3.994	-0.723	-1.301
	(0.05)**	(0.02)**	(0.01)***	(0.36)	(0.25)
(a) R-squared	0.26	0.20	0.18	0.21	0.22
(b) First stage F	6.80	22.04	10.26	25.66	19.34
(c) Pagan-Hall	0.88	0.29	0.92	0.80	0.15
(d) Wu-Hausman	0.00	0.01	0.01	0.41	0.48
(e) C statistic	0.06	0.07	0.06	0.39	0.46
Observations	28	28	29	51	29

Notes: The upper panel is based on high-income and upper-middle countries while lower panel is for lower-middle and low-income countries, as classified in the GDN Growth Database. For other notes, please see Table 4.

Table 6. External trade and financial development (whole sample)

Dependent variable: FDEPTH	OLS levels	Within groups	DIF-GMM	DIF-GMM	SYS-GMM
Instrument set	None	None	Full	Reduced	Reduced
Observations	520	520	432	432	520
Lag 1 FDEPTH	1.032	0.775	0.689	0.453	1.100
	(61.620)***	(19.24) ***	(9.23) ***	(2.15) **	(24.71) ***
Lag 1 OPENC	0.071	0.621	0.973	1.114	0.421
	(1.14)	(3.52) ***	(3.80) ***	(1.92) *	(1.77) *
Serial correlation (m1) p-value			0.00	0.45	0.00
Serial correlation (m2) p-value			0.02	0.12	0.01
Sargan p-value			0.44	0.38	0.03
Granger causality p-value	0.25	0.00	0.00	0.08	0.08
LR effect point estimate	-2.229	2.765	3.126	2.038	-4.206
(Standard error)	(2.51)	(0.85) ***	(1.26) ***	(1.30)	(3.46)

Notes: 88 countries, 1960-1999. Year dummies are included in all models (coefficients not reported). Figures in parentheses below point estimates are t-ratios. * significant at 10%; ** significant at 5%; *** significant at 1%

The GMM results reported here are two-step estimates with heteroskedasticity-consistent standard errors and test statistics; the standard errors are based on the finite sample adjustment of Windmeijer (2005). m1 and m2 are tests for first-order and second-order serial correlation. First-order serial correlation is expected due to first-differencing, but identification of the models relies on the absence of second-order serial correlation.

The Sargan test is used to assess the overidentifying restrictions and is asymptotically distributed as χ^2. The test uses the minimised value of the corresponding two-step GMM estimator. The difference Sargan test is used to test the additional moment conditions used by the system GMM (SYS-GMM) estimator. The Granger causality test examines the null hypothesis that financial development is not Granger-caused by openness; the test statistic is criterion based, using restricted and unrestricted models (see text).

The LR effect is the point estimate of the long-run effect of openness on financial development. Its standard error is approximated using the delta method.

Table 7. External trade and financial development (whole sample)

Dependent variable: FDEPTH	OLS levels	WG	DIF-GMM	DIF-GMM	SYS-GMM	SYS-GMM-1	SYS-GMM-2
Instrument set	None	None	Full	Reduced	Reduced	Reduced	Reduced
Observations	432	432	346	344	432	432	432
Lag 1 FDEPTH	1.251	0.824	0.653	0.388	1.250	1.264	1.147
	(17.10) ***	(14.14) ***	(3.01) ***	(1.43)	(14.99) ***	(14.93) ***	(6.52) ***
Lag 2 FDEPTH	-0.250	-0.253	-0.205	-0.155	-0.294	-0.344	-0.258
	(-3.11) ***	(-4.01) ***	(-2.47) **	(-1.76) *	(-3.61) ***	(-4.20) ***	(-2.48) **
Lag 1 OPENC	0.731	0.803	1.240	1.235	0.968	1.174	0.914
	(3.10) ***	(3.65) ***	(2.48) **	(2.78) ***	(3.72) ***	(2.87) ***	(2.69) ***
Lag 2 OPENC	-0.652	0.027	0.037	0.222	-0.518	-0.547	-0.434
	(-2.96)***	(0.11)	(0.18)	(0.84)	(-2.05) **	(-2.55) **	(-1.94) *
Serial correlation (m1) p-value			0.06	0.55	0.00	0.00	0.00
Serial correlation (m2) p-value			0.86	0.95	0.36	0.53	0.36
Sargan p-value			0.50	0.29	0.20	0.26	0.11
Diff-Sargan p-value					0.21	0.31	0.05
Heterogeneity test p-value					0.16	0.23	0.05
Granger causality p-value	0.01	0.00	0.01	0.02	0.01	0.00	0.02
Test of $\beta_1+\beta_2=0$ p-value	0.25	0.00	0.00	0.01	0.16	0.04	0.14
LR effect point estimate	-47.369	1.935	2.314	1.900	10.328	7.856	4.330
(Standard error)	(581.61)	(0.59) ***	(1.49)	(0.86) **	(14.50)	(7.84)	(3.25)

Notes: 88 countries, 1960-1999. Year dummies are included in all models (coefficients not reported). Figures in parentheses below point estimates are t-ratios. * significant at 10%; ** significant at 5%; *** significant at 1%. The GMM results reported here are two-step estimates with heteroskedasticity-consistent standard errors and test statistics; the standard errors are based on the finite sample adjustment of Windmeijer (2005). m1 and m2 are tests for first-order and second-order serial correlation. First-order serial correlation is expected due to first-differencing, but identification of the models relies on the absence of second-order serial correlation. The Sargan test is used to assess the overidentifying restrictions and is asymptotically distributed as χ^2. The test uses the minimised value of the corresponding two-step GMM estimator. The difference Sargan test is used to test the additional moment conditions used by the system GMM estimators in which SYS GMM uses the standard moment conditions, while SYS GMM (Modified 1) only uses the lagged first-differences of FDEPTH dated t-1 as instruments in levels and SYS GMM (Modified 2) only uses lagged first-differences of OPENC dated t-1 as instruments in levels. The heterogeneity test is used to test the null that there are no individual effects (see text). The Granger causality test examines the null hypothesis that financial development is not Granger-caused by openness; the test statistic is criterion based, using restricted and unrestricted models (see text). The LR effect is the point estimate of the long-run effect of openness on financial development. Its standard error is approximated using the delta method.

Table 8. External trade and financial development (higher-income group)

Dependent variable: FDEPTH	OLS levels	WG	DIF-GMM	DIF-GMM	SYS-GMM	SYS-GMM-1	SYS-GMM-2
Instrument set	None	None	Full	Reduced	Reduced	Reduced	Reduced
Observations	185	185	150	150	185	185	185
Lag 1 FDEPTH	1.166	0.753	0.368	0.123	1.173	1.218	0.957
	(10.01) ***	(8.65) ***	(1.98)	(0.64)	(9.66) ***	(7.01) ***	(2.75) ***
Lag 2 FDEPTH	-0.183	-0.298	-0.273	-0.214	-0.149	-0.393	-0.154
	(-1.38)	(-3.25) ***	(-2.12) **	(-1.91) *	(-1.19)	(-3.02)***	(-1.10)
Lag 1 OPENC	0.686	1.237	0.842	0.734	0.814	0.734	0.586
	(1.58)	(3.10) ***	(1.11)	(1.31)	(1.47)	(1.60)	(0.75)
Lag 2 OPENC	-0.610	-0.013	-0.135	0.063	-0.374	-0.238	-0.573
	(-1.50)	(-0.03)	(-0.42)	(0.29)	(-1.04)	(-0.44)	(-1.16)
Serial correlation (m1) p-value			0.22	0.87	0.02	0.01	0.07
Serial correlation (m2) p-value			0.33	0.47	0.41	0.97	0.68
Sargan p-value			0.92	0.71	0.79	0.53	0.48
Diff-Sargan p-value					0.68	0.17	0.12
Heterogeneity test P-value					0.78	0.28	0.41
Granger causality p-value	0.29	0.00	0.54	0.66	0.56	0.55	1.00
Test of $\beta_1+\beta_2=0$ p-value	0.47	0.00	0.21	0.08	0.39	0.27	1.00
LR effect point estimate	4.574	2.245	0.780	0.731	-18.448	8.582	0.066
(Standard error)	(13.18)	(0.747) ***	(0.93)	(0.58)	(81.76)	(6.72)	(4.56)

Notes: 35 countries, 1960-1999. For other notes please see Table 7.

Table 9. External trade and financial development (lower-income group)

Dependent variable: FDEPTH	OLS levels	WG	DIF-GMM	DIF-GMM	SYS-GMM	SYS-GMM-1	SYS-GMM-2
Instrument set	None	None	Full	Reduced	Reduced	Reduced	Reduced
Observations	247	247	194	194	247	247	247
Lag 1 FDEPTH	1.279	0.849	1.054	0.930	1.202	1.250	0.961
	(15.46) ***	(10.16) ***	(4.40) ***	(2.33) **	(13.43) ***	(9.70) ***	(5.51) ***
Lag 2 FDEPTH	-0.356	-0.252	-0.306	-0.309	-0.366	-0.380	-0.258
	(-4.32) ***	(-2.77) ***	(-3.23) ***	(-2.61) ***	(-5.12) ***	(-4.64) ***	(-2.45) **
Lag 1 OPENC	0.756	0.640	2.258	2.281	0.933	1.262	1.307
	(2.99) ***	(2.52) ***	(2.65) ***	(1.72) *	(2.76) ***	(1.87) *	(3.39) ***
Lag 2 OPENC	-0.666	-0.207	-0.201	-0.204	-0.665	-0.703	-0.453
	(-2.73) ***	(-0.67)	(-0.55)	(-0.45)	(-2.99) ***	(-2.36) **	(-1.42)
Serial correlation (m1) p-value			0.02	0.15	0.00	0.00	0.01
Serial correlation (m2) p-value			0.20	0.34	0.30	0.31	0.35
Sargan p-value			0.74	0.40	0.53	0.47	0.42
Diff-Sargan p-value					0.65	0.57	0.45
Heterogeneity test P-value					0.51	0.32	0.13
Granger causality p-value	0.01	0.04	0.00	0.06	0.06	0.05	0.02
Test of $\beta_1+\beta_2=0$ p-value	0.40	0.19	0.01	0.01	0.31	0.04	0.17
LR effect point estimate	1.158	1.075	8.140	5.484	1.640	4.332	2.877
(Standard error)	(1.13)	(0.80)	(7.37)	(7.26)	(1.63)	(4.74)	(0.94)***

Notes: 53 countries, 1960-1999. For other notes please see Table 7.