Legal reform, contract enforcement and firm size in Mexico*

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Abstract

The variation in legal system quality across states in Mexico is used to examine the relationship between judicial quality and firm size over the course of the 2000s, when systemic changes were taking place. Using economic census microdata and survey-based measures of legal institutions, a robust effect of judicial quality is observed on the firm size distribution and efficiency, instrumenting for underlying historical determinants of institutions. Indicative evidence is found that the effect is strongest in more capital-intensive industries. Market size and distance-to-market are also found to matter for firm size outcomes, consistent with the new trade literature.

Keywords: legal institutions, judicial efficiency, firm size, market potential, distance.

JEL Classification Numbers: F12, K4, L11, 012.

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1 Introduction

Contract enforcement is essential to the efficient functioning of decentralised markets, and legal systems provide necessary institutions to support such enforcement. The quality of the judiciary is based not only on *de jure* laws and regulations, but also on their *de facto* implementation, which can often differ considerably from statutes in countries that are still developing and have relatively weak state capacity. In theory, without a high-quality judiciary, transaction costs may be prohibitive, deterring market transactions and firm entry, inhibiting competition and trade. The literature on growth and development has long argued for a fundamental role of "deep" institutions such as the judiciary in determining long-run economic outcomes (Acemoglu *et al.*, 2005), yet identification of the mechanisms at play is often difficult (see OECD, 2012; Palumbo *et al.*, forthcoming).

Increasing firm scale or size is the main channel through which the most efficient firms can expand their production, by taking on capital and labor as they grow. This up-scaling may be motivated by competition with less efficient firms, who give up market share, particularly when they exit the market. Such dynamics are thought to be a main driver of aggregate productivity growth in open economies (Melitz, 2003; Melitz and Ottaviano, 2008), though there is also evidence of substantial within-firm productivity gains induced by domestic and foreign competition pressures (Harrison *et al.*, 2011; Ben Yahmed and Dougherty, 2013).

This paper examines the link between legal systems and firm size in a large developing economy – Mexico – where legal system quality and enforcement not only varies across states, but is also in the process of being substantially reformed. The paper proceeds by following closely the key reference in the field, Laeven and Woodruff (2007), which also focused on Mexico. While the earlier paper focused exclusively on the year 1998, we look at an elapsed and more recent time period – 2003 to 2009 – when major institutional changes were taking place, using insights from the international trade literature to interpret the mechanisms at work. Our study relates to Kumar *et al.* (2002), who carried out an analysis of the effect of court efficiency on average firm size across 15 European jurisdictions, as well as studies of Italy and Spain, where firms in provinces with more efficient courts are found to be larger (Fabbri, 2010; Bürker and Minerva, 2012; Giacomelli and Menon, 2012; García-Posada and Mora-Sanguinetti, 2013).

Taking advantage of the state and time variation in the new Mexico dataset, we find that firms in states with higher judicial quality tend to be substantially larger than those in remaining states, as did Laeven and Woodruff (2007). However, this result is not only robust to alternative measures of firm size and instrumentation for the potential endogeneity of legal system quality, but also to the use of panel data estimators that exploit the evolution of judicial quality over time. Additionally, we find that firms in more capital intensive industries are more likely to benefit from higher quality judicial systems, consistent with insights from the incomplete contracting literature, suggesting that hold-up problems may be limiting the scaling up of firms.

The paper proceeds by considering the theoretical linkage between firm size and judicial quality in the next section. In the third section, we discuss the data used in the analysis. In the fourth section, the estimation strategy, empirical results and their implications are discussed. The fifth section concludes.

2 Firm size and legal systems

The industrial organization and new institutional economics literatures both give support for the idea that average firm size should be positively related to legal system characteristics (see Kumar *et al.*, 2002). We focus on a particular model that cuts across these literatures, based on Laeven and Woodruff's (2007) adaptation of Lucas's (1978) model of firm size, which views the legal system as reducing the investment risk faced by entrepreneurs who invest an increasing share of their wealth in an enterprise. The model predicts that improvements in the legal system will cause an increase in the demand for capital and labor from all entrepreneurs. This in turn puts upward pressure on wage and rental rates, inducing entrepreneurs with low ability to leave self-employment for wage work in incorporated firms. As a result, average firm size increases.

A related adaptation of the Lucas (1978) model by Quintin (2008) sees the contractual framework as imperfect, with a variable degree of enforcement across jurisdictions, and proxies the quality of the legal system as an exogenous probability that agents will default. They calibrate their model to the firm size distribution in the United States, Mexico and Argentina, and show that differences in enforcement in the model explain a sizable part of the observed differences in economy-wide firm size distributions.

Delving more into the mechanisms that determine the link between the legal system and firm size, the role of capital intensity appears to be critical. The Grossman-Hart-Moore property rights theory of the firm emphasizes the importance of hold-up costs in contracting, which can make capital-intensive and input-dependent investments especially risky. However, the evidence for this idea in the context of legal systems is mixed.[1] Laeven and Woodruff (2007) find no significant effect of increasing capital intensity or decreasing vertical integration, but do find a role for risk diversification through incorporation in affecting the incentives of entrepreneurs, increasing average firm size. Kumar *et al.* (2002) even find some evidence in the opposite direction: that firms in more capital intensive industries are less affected by judicial quality, which they speculate to be attributable to physical capital needing less protection than do intangible assets. In contrast, our study re-examines these questions, finding new evidence supporting the idea that hold-up costs may be more substantial in capital intensive manufacturing industries, making the quality of the legal system even more important for these sectors to support larger firm sizes.

Limited guidance is available on the shape of the firm size distribution. The model of Guner *et al.* (2008) suggests that the firm size distribution should be more dispersed when there are fewer restrictions on capital use. However, Bürker and Minerva (2012) find that in Italian provinces with shorter civil trials – one narrow measure of judicial quality – the firm size distribution is more compact. We, however, find evidence that the shape of the distribution tends to be more dispersed when firms are located in Mexican states with better judicial quality, using our fairly broad measure. Next we turn to the data at hand.

[1] Evidence from Nunn (2007) and Ma *et al.* (2010) also supports the idea that hold-up costs from a weak judiciary may distort the comparative advantage of some industries through an influence on their input structure, though outcomes are measured in terms of exporting rather than increasing firm size. Such findings may suggest that improving legal institutions allows firms to better specialize by reducing transactions costs (see Williamson, 2005).

3 Data

The key data used in this study are economic census microdata, for measuring firm size and characteristics, and survey-based data that measure judicial quality for contract enforcement, along with state-level demographic, distance and gravity-type data that are included as control variables.

3.1 Bin-level economic census data

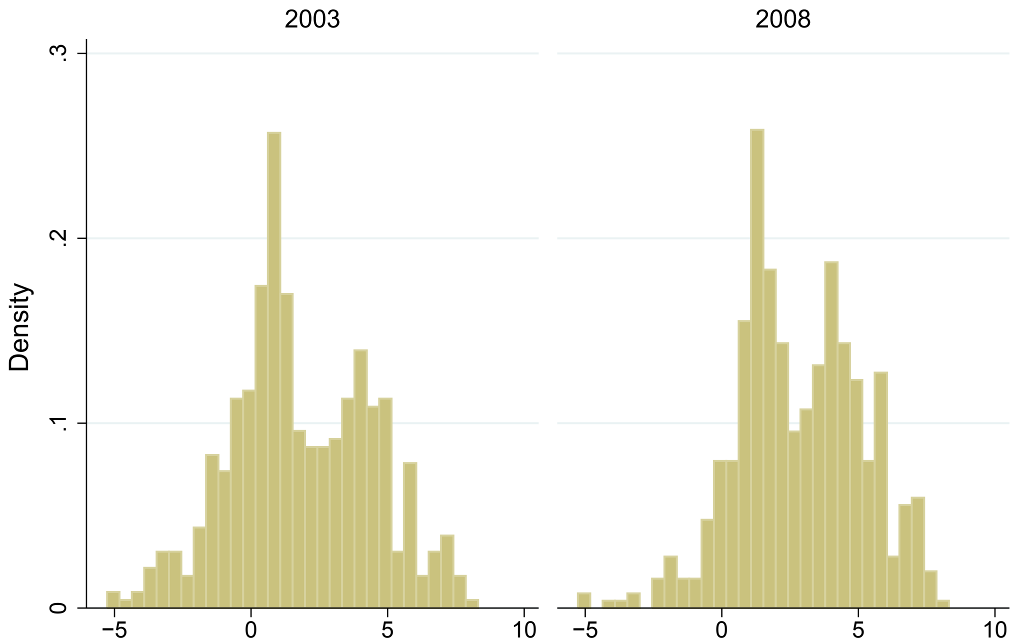
Disaggregated data from the Mexican economic census were sourced from the *Instituto Nacional de Estadística y Geografía* (INEGI) in Aguacalientes, for the census years 2003 and 2008. The economic census enumerates all fixed establishments in Mexico every five years, and we sourced the information it collects from firms on output, employment, fixed assets, and intermediate inputs. While we gained access to the unit-level microdata, due to the complexity of confidentiality procedures, we chose to use the data at the firm-size bin level, within each industry, state and year cell. This data is collected and enumerated at the establishment level, and multi-plant firm data is only available at the level of establishments. However, since 97.5% of firms in Mexico are single-establishment, the data correspond closely to purely firm-level data.[2]

This data is available for all 31 Mexican states and the *Distrito Federal* (Mexico City) from the level of "sub-sector", or three-digit industry. Within manufacturing, where we focus most of our analysis, there are up to 21 such industries in each state. These industries are then stratified by firm size, in the following size "bins": 0 to 2; 3 to 5; 6 to 10; 11 to 15; 16 to 20; 21 to 30; 31 to 50; 51 to 100; 101 to 250; 251 to 500; 501 to 1000; and over 1000. The bin-level data allow for computation of a weighted-average firm size at the industry level by state.

An employee-weighted firm size $EWFS$, following the approach of Kumar *et al.* (2002)

[2] We were able to obtain an extract of the 2008 data that removes the multi-plant firms from the sample, and robustness checks using this data confirm our key findings.

Figure 1: Distribution of log employee-weighted firm size



Source: Calculations from INEGI economic census data.

and Laeven and Woodruff (2007), is specified as follows, for state s, industry i and time t:

$$EWFS_{s,i,t} = \sum_{b=1}^{n} \left(\frac{N^{emp}_{b,s,i,t}}{N^{emp}_{s,i,t}} \right) \left(\frac{N^{emp}_{b,s,i,t}}{N^{firms}_{b,s,i,t}} \right) \tag{1}$$

where b is a firm-size bin, and $N^{emp}_{b,s,i,t}$ captures the employment in a single bin, for all bins with more than three firms,[3] and up to a dozen bins. $N^{emp}_{s,i,t}$ captures the employment in a state-industry-year cell, and $N^{firms}_{b,s,i,t}$ captures the total number of firms in a bin. The formula weights average firm size by the share of employment in each firm size bin. We use the natural log of EWFS, and the distribution of this variable in manufacturing is shown for 2003 and 2008 in Figure 1. The distribution is normal according to standard tests, although qualitatively there appears to be some indication that it could be bi-modal.[4]

[3] When three or fewer units are present in a firm bin (true for 15% of establishments), the firm count data are suppressed, and we exclude these bins from our firm size analysis. These mostly relate to industries where there are few firms in any size class, and thus are unlikely to affect the analysis. However, in some cases, the largest size class has a smaller number of firms, and in these cases we lose some representativeness of this size class, which is partially compensated for by the weighting of firm size.

[4] As a result, when we use log($EWFS$) as a dependent variable below, we also carry out quantile regressions, though there does not appear to be any significant difference across quartiles in the main coefficients of interest.

Equation (1) gives greater weight than a simple average to those bins that contain larger firms. Alternative measures, including a simple average of firm size across bins and the average size of firms in the bin with the median worker, are also computed.

3.2 Measures of judicial quality

The Intituto Technológico Autónomo de México (ITAM) and the law firm Gaxiola, Moraila and Associates (GMA) have cooperated since 2001 with Moody's Investors Service to measure the efficiency of state institutions involved in the administration of justice every 2-3 years, for 2001, 2003, 2006, 2008 and 2011. These studies focus on the adequacy of legal administration, as it relates to the enforceability of commercial contracts and mortgages disputed in state court in Mexico. The sources of information for the measures comprise expert opinion surveys completed by litigation attorneys in each of the federal entities, information supplied by each state tribunal, data collected by the researchers and on-site visits by ITAM in each state.

The four key factors considered in the measures are:

1. *Institutional quality – 50%*: factors within and outside the control of the judicial branch in a state that affect its ability to carry out its functions. These include the perceived quality of tribunals' judges and magistrates, their expertise in commercial cases, the criteria required for the promotion of judges and the nomination of magistrates and the impartiality of persons in both of these positions.

2. *Duration of cases – 40%*: the average time and backlog involved in processing a typical case related to contract enforcement.

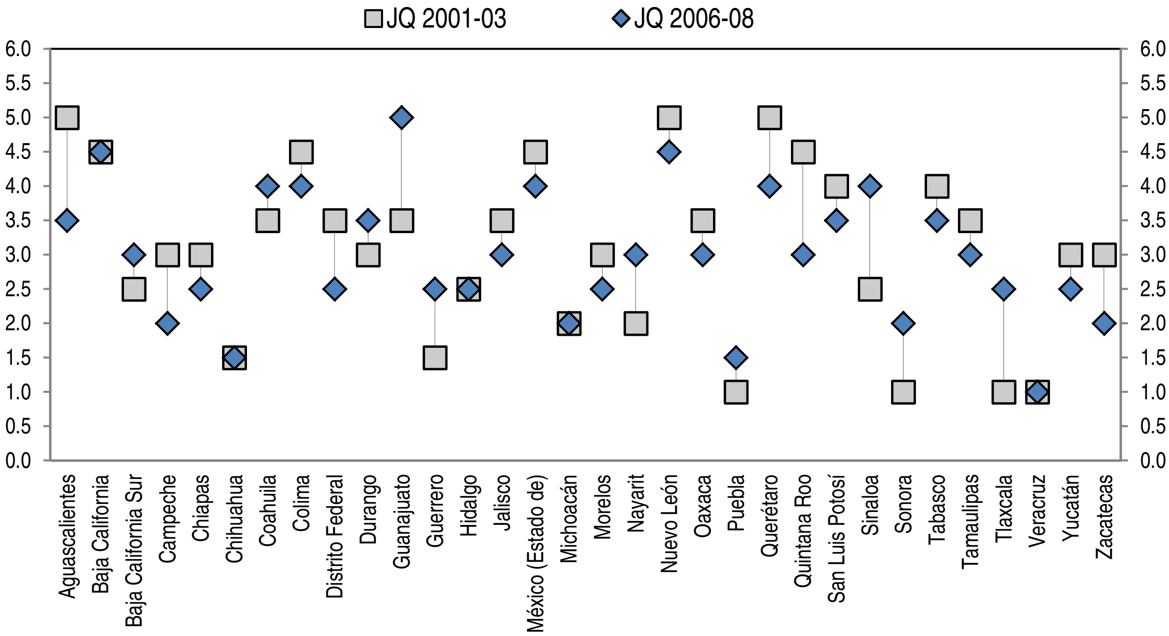
3. *Quantity and efficiency in the use of resources – 10%*: human and physical resources devoted to the judicial branch, including an assessment of their quality by legal practitioners.

4. *Enforcement of resolutions – adjustment to the criteria*: an evaluation of the support provided by the executive branch in obtaining final enforcement of verdicts.

These factors are combined into a five-level score by Moody's (2011), from which we code judicial quality from worst to best as follows:

- *Lowest quality (EC5)*: scored 1
- *Below-average quality (EC4)*: scored 2
- *Average quality (EC3)*: scored 3

Figure 2: Evolution of judicial quality measure by state
Rated on a scale of lowest (1) to highest (5) quality



Source: Moody's based on ITAM and Gaxiola, Moraila and Associates survey.

- *Above-average quality (EC2)*: scored 4
- *Highest quality (EC1)*: scored 5

While the scoring system remains consistent over time, Moody's recommends some caution in interpreting the changes in EC over time, due to possible sampling variation and methodological changes in some of the underlying indicators across vintages. As a result, we give initial emphasis to the pooled results, and include state and time dummies in the panel-based estimates.

The scores that we use are based on an inversion of the digit associated with the Moody's EC rating (see Table A.1), and we disregard the "+" that some states receive for being at the top end of a given rating. Using our scoring of the Moody's measure of contract enforcement, a higher score for judicial quality JQ represents "better" state-level institutions.

Comparison across states of the scores for judicial quality in each of the state-level entities, in Figure 2, shows considerable heterogeneity for the period immediately preceding 2003 and 2008. The measures for 2006 and 2008 are averaged by state and shown with diamonds, and the measures for 2001 and 2003 are averaged and shown with squares.

There is evidence of a deterioration in the state-level scores for two-thirds of the states over this period, consistent with anecdotal and qualitative evidence. The reasons for this worrying devolution of judicial quality varied from state-to-state and period-to-period, and lacked a common driver. The overall decline in the state-level indicators appears to be related to the slow pace of legal reform in many states, that created an important degree of variation across states and time, which we take advantage of in our empirical estimates.

During the mid-2000s, a number of states started reforming their civil and criminal judicial systems, in advance of a 2008 constitutional amendment creating a federal framework for reforming the criminal justice system. However, there was less attention to civil justice procedures and contract enforcement issues. This seems to have been partly due to a lack of sufficient will to undertake reforms in better-performing states, which may have triggered neglect: there is some evidence that on average, states with better judiciaries were the least likely to undertake the reforms (see OECD, 2013). A comprehensive account of the criminal justice reforms by Ingram (2013) summarized the pace as: "highly uneven, unstable, and can alter quickly over the span of a few months, depending on budgetary conditions, federal incentives, legislative calendars, local political conditions, media coverage, societal support, and other factors." Moreover, the period covered in this study coincided with a rise in drug trafficking-related violence in many of the northern states. As a result, we include the crime (murder) rate as an additional control variable.

A map of the most recent vintage of the judicial quality scores for 2011, shown in Figure 3, reveals few obvious patterns in the spatial distribution of judicial quality, although there is some indication that states closer to the border with the United States may have tended to score more highly.

In addition to legal system quality, we use a standard measure of financial market depth – the ratio of private credit to GDP, from the Banco de México. This measure uses a straightforward measure of financial market development, and gives a useful additional robustness check to our estimates of judicial quality, to ensure that they are not instead partly capturing the development of financial market institutions.

Figure 3: Map of judicial quality measure by state in 2011
Rated on a scale of lowest (1) to highest (5) quality

Source: Moody's based on ITAM and Gaxiola, Moraila and Associates surveys.

3.3 Instruments

Judicial quality cannot be considered to be exogenous to economic outcomes such as investment and firm size, so we employ an instrumentation strategy taking inspiration from Acemoglu *et al.* (2005), and also used by Laeven and Woodruff (2007) in their earlier study. The key instruments are indigenous state-level population in 1900, and the number of crops with large economies of scale in 1939. The justification for their use is based, first, on the use of the *encomienda* system imported from Europe, that treated indigenous labor as a resource to be used by the ruling elite. Hence, the presence of a larger share of indigenous people could be expected to be associated with a worse institutional environment. The second instrument is based on the presence of substantial production of crops that had sufficiently large economies of scale that they led to substantial distortions in the distribution of land and income. Thus, where there was more cultivation of sugar, coffee, rice and cotton as revealed in the 1940 census, we expect that political institutions and thus legal system should be worse. The correlation between both instruments is low (-0.11), and together they explain an appreciable share of the state-level variation in judicial quality.

3.4 Geographic controls

Mexico's firm size distribution has been found to be distorted, and skewed towards small firms, especially when compared with the United States (see Hsieh and Klenow, 2012). However, this is in part due to its considerably smaller market size, as measured by either GDP or population. Theoretical work in the trade literature has demonstrated that in a monopolistically competitive model with firm heterogeneity, average firm size is larger and dispersion is higher in larger markets (Melitz and Ottaviano, 2008). Thus, we control for market size using the log of state population, or alternatively the log of total state GDP, though the later is more likely to be endogenous.

The firm geography literature also makes predictions about export market success and consequently firm size (see Redding and Venables, 2004), and we thus use several different variables to proxy distance to market and foreign market potential. These controls include the following: (*i*) distance to the nearest major point of entry into the United States, from Rios and Romo (2008); (*ii*) the average distance to the closest of one of the 10 largest cities in Mexico, weighted by the inverse of the distance, from the same source; and (*iii*) foreign market potential, as estimated by Escobar Gamboa (2010) using the (Head and Mayer, 2004) method. We also use GDP per capita and murders per capita to proxy the level of development and the crime rate, sourced from the INEGI and OECD Regional databases, respectively. Next we turn to the estimations.

4 Estimation strategy and results

The empirical analysis starts with a basic estimation of the firm size equation and its distribution, then a series of additional variables are introduced to examine the robustness of the relationship, and alternative measures are explored, as well as interactions. While the basic analysis closely follows the approach of Laeven and Woodruff (2007), additional control variables are used, and the panel nature of the present dataset is fully exploited. Finally, a production function is estimated that looks at efficiency implications.

Summary statistics for the variables used in the main analysis are shown in Table 1. They show that across industries, the average employment-weighted firm size is 119, while the simple average is 56. Yet the typical size of a firm (in the median size bin) is only 20. The average industry has slightly over 100 firms in total.

Table 1: Summary statistics for pooled 2003 and 2008 data

VARIABLE	Obs.	Mean	Std. Dev.	Min.	Max.
Industry/state-level /a					
Employment-weighted firm size	1,062	119.0	328.1	0.005	4,076
Log employment-weighted firm size	1,062	2.4	2.49	-5.27	8.31
Simple average firm size	1,062	56.3	115.4	1.1	1,952
Typical firm size (median bin)	1,062	19.6	72.5	1.1	1,952
Number of employees	1,062	8,172	13,833	14.0	161,347
Number of firms (establishments)	1,062	108.6	1,517	3.0	19,451
Log capital intensity (K/L) /b	538	1.72	1.65	-4.81	5.39
Log vertical integration (GO/VA) /b	550	1.03	.369	.156	3.79
Log value added /b	550	12.9	2.31	4.76	18.0
Log gross output /b	555	13.9	2.36	5.77	18.9
State-level					
Judicial quality (JQ)	64	3.1	1.11	1	5
Log JQ score	64	1.0	0.44	0	1.61
Market size (log population)	64	14.8	0.76	13.1	16.5
Foreign market potential	32	.097	.246	.014	1.35
Log international distance to market	32	6.2	1.90	0	7.74
Log domestic distance to market	32	7.1	1.4	4.1	9.6
Log private credit as a share of GDP	64	-1.9	0.9	-4.6	0.10
Log real GDP in millions of pesos	64	19.0	0.84	17.5	21.1
Log GDP per capita	64	8.8	0.48	8.1	10.8
Log murder rate	64	2.1	0.42	0.86	2.90
Log indigenous share in 1900	32	0.1	0.19	0	0.69
Number of large-scale crops in 1939	32	1.7	1.2	0	4

/a For industry and state pairs where firm size data is available. **/b** 2008 only.
/c The correlation of employment-weighted firm size with the unweighted measure is 0.944, while with the median measure it is 0.862.

4.1 Basic estimates

The default estimation equation, using measured firm size for each state s, industry i, and time t cell, is as follows:

$$firm\_size_{s,i,t} = \alpha_{i,t} + \beta B_{s,t} + \gamma \Gamma_{s,i,t} + \epsilon_{s,i,t} \tag{2}$$

where $\alpha_{i,t}$ is an industry-year fixed effect, $B_{s,t}$ is a vector of state-level variables, $\Gamma_{s,i,t}$ is a vector of variables that vary by state and industry, when applicable; $\epsilon_{s,i,t}$ is the error term. Since the data vary over time, not only a pooled estimate of the equation is carried out, but also panel estimates.

In the first set of regressions, shown in Table 2 (columns 1 to 3), we set $firm\_size$ as $\log(EWFS)$, and regress it on judicial quality and market size, using the pooled 2003 and 2008 data, with industry-year fixed effects. Estimates of the equation are shown using both ordinary least squares (OLS) and two-stage least squares instrumental variables (IV) methods. All estimates show a significant positive coefficient on judicial quality (JQ), supporting our hypothesis that states with better legal institutions should have larger firms on average. Market size using state population is also found to be significant, with systematically high levels of significance. All reported standard errors are clustered by state and are thus robust to intra-state correlations.

The OLS estimates of JQ appear to be substantially biased downwards, as the IV estimates show much higher coefficients on JQ, which is well-identified using the standard Hansen overidentification test once both instruments are included. F-tests on the first stage are also far above the "weak instrument" threshold of 10 where the performance of 2SLS estimators tends to be less robust (see Stock *et al.*, 2002). The first stage equation shows the expected signs on both of the instruments when they are included together.

The second set of regressions in Table 2 (columns 4 and 5) include our preferred gravity variable, the log of the distance to the nearest point of entry to the United States, which we call distance to international markets, since the US border is the departure point for most of Mexico's exports. This variable is also highly significant (and negative), along with state-level market size (positive).

Our preferred specification is the final IV estimate (column 5), which shows a coefficient of 2.613 on judicial quality, and represents an average increase of 17% in terms of weighted average firm size for each one-step increase in judicial quality, or an increase by just over two-thirds of the weighted average firm size of 119 if the legal system in a low-performing state were to improve from the worst to best practice in judicial quality. An illustration of the impact of such a dramatic reform on the predicted firm size distribution is shown

Table 2: Baseline estimates of the effect of judicial quality

	Dependent variable: weighted average firm size				
	(1)	(2)	(3)	(4)	(5)
VARIABLES	OLS-1	IV-1a	IV-1b	OLS-2	IV-2
Judicial quality (JQ)	0.662**	3.411*	3.232***	0.705***	2.613***
	(0.319)	(1.705)	(0.563)	(0.247)	(0.478)
Market size	1.321***	1.484***	1.473***	1.274***	1.381***
	(0.126)	(0.173)	(0.109)	(0.148)	(0.107)
Distance to int'l markets				-0.210***	-0.231***
				(0.074)	(0.045)
Constant	-17.144***	-22.337***	-21.998***	-15.159***	-18.556***
	(1.832)	(3.817)	(1.672)	(2.312)	(1.619)
Observations	1,062	1,062	1,062	1,062	1,062
R-squared	0.368	0.374	0.412	0.393	0.420
Instrumented	No	Yes	Yes	No	Yes
IV First Stage: JQ					
Indigenous		-0.555	-0.603*		-0.732**
		(0.332)	(0.297)		(0.319)
Crops			-0.131**		-0.137**
			(0.055)		(0.056)
Market size		-0.068	-0.008		0.003
		(0.103)	(0.080)		(0.079)
Distance to int'l markets					0.040
					(0.045)
Constant		2.097	1.442		1.064
		(1.510)	(1.195)		(1.239)
Observations (States)		32	32		32
Hansen overidentification test (p-value)			0.692		0.192
F-statistic (first-stage 2SLS)		60.8	112.8		132.3
Partial R-squared		0.056	0.182		0.207
Industry-year dummies	Yes	Yes	Yes	Yes	Yes

Robust standard errors in parentheses, clustered by state. *** p<0.01, ** p<0.05, * p<0.1

in Figure 4, plotted on a log scale. This plot shows that the shape of the distribution also shifts with the increase in judicial quality, becoming more dispersed, as hypothesized.[5]

4.2 Robustness tests

Several sets of robustness checks are carried out: (*i*) to account for the possibility of additional geographic and gravity-type effects; (*ii*) with alternative measures of firm size

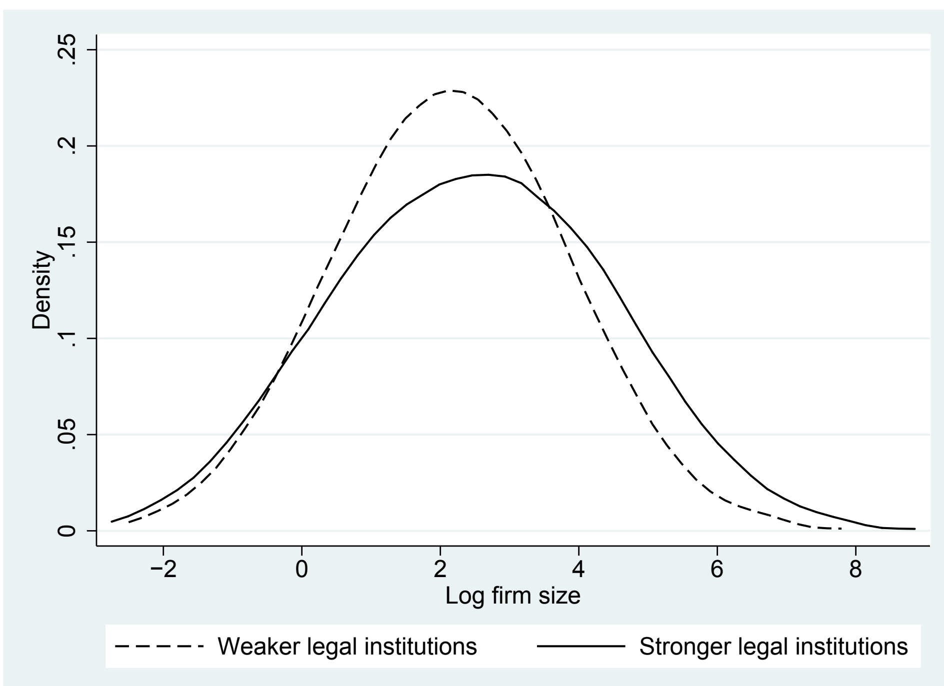
[5] A quantile regression was also run using the preferred equation specification, and no significant difference was found in the estimated JQ coefficient for firms at the first and third quartiles, as compared with the median firm. However, the estimated coefficients were slightly larger for firms in the lower half of the distribution, which is consistent the the idea that they face greater effective barriers to up-scaling.

and institutions; (*iii*) using various panel-type specifications with respect to time.

The first set of these regressions are shown in Table 3, Panel A, compared with the preferred IV specification above. As with the previous results, and all of those in this table, F-tests on the first stage results are also well above the critical threshold of 10. Using an additional distance variable, for distance to domestic markets (column 2), does reduce the size of the estimated (IV) coefficient. This is likely a result of multicolinearity, as these two variables are highly correlated (see Rios and Romo, 2008), and partly for this reason, foreign market potential is often preferred (Benassy-Quere *et al.*, 2005). Thus, we also estimate the equation using Escobar Gamboa's (2010) estimate of foreign market potential (FMP) in 2002 for each Mexican state (column 3), and while FMP is not significant, judicial quality remains strongly significant. When total GDP (in the year of observation) is used in place of total population (column 4), the effect of JQ remains significant, although the size of the coefficient is diminished.

Alternative estimates of the quality of related institutions shows that the effects we

Figure 4: Predicted employee-weighted firm size distribution
Density function conditional on presence in best or worst-practice state



Source: Calculations using the "preferred" equation estimate in Table 2, column 5.

Table 3: Robustness checks of preferred specification

Panel A

VARIABLES	(1) Preferred specif- ication	(2) with domestic distance	(3) with market potential	(4) using GDP for market size	(5) with private credit
Judicial quality	2.613***	1.172**	3.126***	1.256**	2.354***
	(0.478)	(0.566)	(0.562)	(0.486)	(0.552)
Market size	1.381***	0.767***	1.463***		1.305***
	(0.107)	(0.229)	(0.103)		(0.132)
Distance to int'l markets	-0.231***	-0.193***		-0.174***	-0.225***
	(0.045)	(0.032)		(0.059)	(0.043)
Distance to domestic markets		-0.441***			
		(0.140)			
Foreign market potential			0.369		
			(0.284)		
GDP size				1.190***	
				(0.151)	
Private credit					0.165
					(0.182)
Constant	-18.556***	-5.103	-21.778***	-19.682***	-16.824***
	(1.619)	(4.697)	(1.553)	(3.049)	(2.389)
Observations	1,062	1,062	1,062	1,062	1,062
R-squared	0.420	0.438	0.419	0.411	0.426
Instrumented	Yes	Yes	Yes	Yes	Yes
Industry-year dummies	Yes	Yes	Yes	Yes	Yes

Panel B

VARIABLES	(6) with crime rate	(7) with GDP per capita	(8) Average firm size	(9) Typical firm size	(10) All sectors
Judicial quality	2.486***	2.272***	1.819***	1.097***	1.374***
	(0.448)	(0.568)	(0.349)	(0.222)	(0.342)
Market size	1.313***	1.402***	0.915***	0.554***	0.707***
	(0.114)	(0.123)	(0.073)	(0.052)	(0.096)
Distance to int'l markets	-0.243***	-0.212***	-0.193***	-0.136***	-0.040**
	(0.050)	(0.050)	(0.029)	(0.020)	(0.017)
Murders per capita	-0.524**				
	(0.218)				
GDP per capita		0.347			
		(0.278)			
Constant	-16.240***	-21.677***	-10.672***	-5.830***	-10.232***
	(1.929)	(3.100)	(1.143)	(0.790)	(1.635)
Observations	1,062	1,062	1,062	1,062	3,566
R-squared	0.423	0.428	0.536	0.528	0.494
Instrumented	Yes	Yes	Yes	Yes	Yes
Industry-year dummies	Yes	Yes	Yes	Yes	Yes

Robust standard errors in parentheses, clustered by state. *** p<0.01, ** p<0.05, * p<0.1

are observing are unlikely to be spurious. Private credit as a share of GDP directly measures financial market depth, and can also be thought of as complementary to the quality of contract enforcement (see Laeven and Woodruff, 2007). When we add this variable (column 5), it is not significant and the effect of JQ on weighted firm size remains significant.

Adding the crime rate, Table 3, Panel B (column 6) – measured as the annual number of murders per capita – shows that it indeed is associated with significantly reduced average firm size, as expected, yet it does not substantially diminish the impact of JQ on firm size. Controlling for the overall level of development using GDP per capita (column 7) reduces slightly the size of the estimated coefficient on JQ, but it remains large and significant, though the coefficient on GDP per capita is not significant.

Estimates using alternative measures of firm size are shown in columns 8 and 9. The estimates using a simple average firm size (column 8) and typical firm size (column 9) still support significantly positive effects of judicial quality, and the R-squared for these regressions is even higher than for our preferred employment-weighted firm size.

In all of the estimates so far, only manufacturing industries have been included. This is intended to ensure that the firms included were roughly comparable, since they are primarily goods-producing sectors, where economies of scale are clearer. If we include all services, the sample size is about four times larger; however, these firms are not as comparable given the extreme heterogeneity of their activities with respect to "optimal" firm size (since this includes many sole proprietorship-dominated service sectors). Nevertheless, if we carry out an estimate using all industries (Column 10), the coefficient on JQ remains significant and positive. However, distance to international markets has a smaller coefficient, presumably indicating the fact some of these sectors do not engage extensively in international trade.

4.3 Panel estimates

So far the estimates have been based on pooled estimates over 2003 and 2008; while this is an improvement over the earlier results by Laeven and Woodruff (2007) who examined only a single year (1998), in order to more fully take advantage of the time dimension of the judicial quality data, panel data models are estimated using those control variables that vary over time. Our preferred results make use of a technique by Baum *et al.* (2007)

Table 4: Robustness checks taking into account changes over time

Panel A

VARIABLES	(1) Random effects (RE)	(2) RE with crime rate	(3) RE without instruments	(4) Fixed effects (FE)	(5) FE with crime rate
Judicial quality	2.720***	2.587***	0.636***	0.667**	0.677**
	(0.382)	(0.371)	(0.231)	(0.283)	(0.286)
Market size	1.403***	1.332***	1.291***	15.447***	15.455***
	(0.103)	(0.103)	(0.159)	(1.329)	(1.329)
Distance to int'l markets	-0.237***	-0.249***	-0.217***		
	(0.040)	(0.040)	(0.075)		
Murders per capita		-0.542***			-0.082
		(0.172)			(0.360)
Constant	-20.108***	-17.722***	-17.599***		
	(1.709)	(1.798)	(2.662)		
Observations	1,062	1,062	1,062	998	998
R-squared	0.317	0.331	0.390	0.216	0.216
Number of groups	563	563	563	499	499
Instrumented	Yes	Yes	No	No	No
Industry-year dummies	Yes	Yes	Yes	No /b	No /b

Panel B

VARIABLES	(6) Using first differences	(7) using GDP size	(8) with private credit	(9) with crime rate	(10) with GDP per capita
Change in judicial quality	0.340*	0.350*	0.376*	0.349*	0.338*
	(0.187)	(0.182)	(0.192)	(0.197)	(0.182)
Change in market size	2.966		3.227*	2.977	3.716*
	(1.793)		(1.737)	(1.813)	(1.942)
Change in GDP size		2.441			
		(1.484)			
Change in private credit			-0.177		
			(0.215)		
Change in crime rate				-0.078	
				(0.561)	
Change in GDP per capita					1.830
					(1.550)
Constant	-1.204	-1.504*	-1.097	-1.194	-1.493*
	(0.735)	(0.826)	(0.726)	(0.727)	(0.811)
Observations	499	499	499	499	499
R-squared	0.110	0.111	0.111	0.110	0.113
Instrumented	No	No	No	No	No
Industry dummies	Yes	Yes	Yes	Yes	Yes

/a Robust standard errors in parentheses, clustered by state. *** p<0.01, ** p<0.05, * p<0.1
/b Fixed effects estimators include both industry and year dummies

that allows the use of instrumental variables with a panel estimator. Since the instruments are time invariant, we can only employ this technique with random effects and in levels. The random effects are defined by industry and state groupings. The results, shown in Table 4, Panel A, show that using either random effects (columns 1 to 3) or fixed effects (columns 4 and 5), the strong positive relationship remains between judicial quality and firm size. As with the pooled estimates, the results are robust to the inclusion of the crime rate (columns 2 and 5).

The size of the coefficients with both the IV-2SLS and OLS estimators is comparable to those from the pooled results shown in Table 2, though the size of the effect is lower when using fixed effects, or when the relationship is estimated in first differences (Panel B).[6] Various specifications were tried with the first difference specification (columns 6 to 10), and the estimated effect of judicial quality on firm size growth is similar and significantly positive in each case, suggesting that there may be dynamic effects. While the coefficient on the change in market size is significant in two of the five differenced specifications, the other control variables are not significant.

4.4 Interactions

Next we explore the mechanisms that may explain the effects that we have been observing, using the ratio of fixed assets to employment to measure capital intensity, and the ratio of gross output to value added to measure the degree of intermediate input use, or decreasing vertical integration. Extending the estimation equation to include interactions terms with judicial quality gives us:

$$firm\_size_{s,i} = \alpha_i + \beta B_s + \gamma \Gamma_{s,i} + \xi X_{s,i} \Omega_s + \epsilon_{s,i} \tag{3}$$

where the variables are as in equation (3), except an extra term $X_{s,i}\Omega_s$ is included, where $X_{s,i}$ is the log of the ratio of fixed assets to employment for each state s and industry i, or alternatively, the ratio of gross output to value added, and Ω_s is judicial quality in state

[6] While the estimates use employment-weighted firm size as the dependent variable, these results are robust to the use of average firm size and typical firm size as the dependent variable.

s, that may also be included in the vector B_s. We estimate this equation using only 2008 data, due to the availability of capital intensity.

Table 5: Interactions with capital intensity and vertical integration
2008 data only

VARIABLES	(1) Preferred 2008 eqn.	(2) Without IV	(3) Capital intensity	(4) only X term	(5) Vertical integration	(6) only X term
Judicial quality	2.972***	0.951***	0.282		0.853**	
	(0.739)	(0.287)	(0.290)		(0.380)	
JQ X Capital intensity			0.285***	0.145**		
			(0.049)	(0.056)		
JQ X Vertical integration					0.085	-0.012
					(0.271)	(0.277)
Market size	1.216***	1.118***	1.039***	0.983***	1.089***	0.955***
	(0.123)	(0.146)	(0.136)	(0.033)	(0.146)	(0.018)
Distance to int'l markets	-0.279***	-0.249***	-0.253***	-0.164***	-0.276***	-0.016
	(0.058)	(0.084)	(0.078)	(0.020)	(0.074)	(0.026)
Constant	-13.732***	-10.219***	-14.155***	-12.573***	-13.935***	-12.297***
	(2.015)	(2.591)	(2.392)	(1.038)	(2.567)	(1.026)
Observations	555	555	538	538	550	550
R-squared	0.414	0.390	0.426	0.517	0.394	0.504
Instrumented	Yes	No	No	No	No	No
State dummies	No	No	No	Yes	No	Yes
Industry dummies	Yes	Yes	Yes	Yes	Yes	Yes

Robust standard errors in parentheses, clustered by state. *** p<0.01, ** p<0.05, * p<0.1

The estimates including capital intensity imply that the capital intensity of a firms' activities may be a dominant mechanism determining the effect of judicial quality on firm size. For comparison, in Table 5, the preferred IV and OLS equation is estimated using only 2008 data (columns 1 and 2). The OLS estimate would appear to be a lower-bound estimate, due to OLS bias. In the estimate when JQ and its interaction with capital intensity are both included (column 3), only the interaction term is significant. When the interaction term is estimated by itself (with state dummies), it remains significant, though the coefficient is smaller. Effects of decreasing vertical integration (or heightened input use) in affecting firm size through judicial quality are not observed in our estimates (columns 5 and 6) when the direct effect of JQ is included, nor when the interaction is estimated by itself (with state dummies). These findings contrast with those of Laeven and Woodruff (2007), who find little role for capital intensity, and find more evidence for vertical integration.

An important robustness check examines whether there may be selection biases through the channel of multi-plant and foreign-invested firms, who could potentially select the states with better-quality judicial systems as the home of their investments. Moreover, foreign companies may not necessarily be subject to the weaknesses of state courts, since some of them rely on arbitration to handle certain types of contractual disputes. At the same time, we also limit the sample to single-plant firms, since this better accounts for the possibility that a firm with plants in multiple states may arbitrage across different state courts. The results of these estimates (Table A.2, columns 1 to 6) re-confirm the basic findings we obtained using the full sample, with a similar coefficient size when using instruments. In addition, the effect of capital intensity is statistically significant but smaller compared to the main ones, while the results with vertical integration are unsurprisingly even more ambiguous, without the multi-plant firms.[7]

4.5 Efficiency effects

At the outset, we argued that removing barriers to firm growth was a "good" in that it facilitates the expansion of more efficient firms, promoting entrepreneurship and productivity. While this link is well-known, we have adequate factor input data in 2008 to perform a straightforward test, using production function estimates at the industry level, following the approach of Laeven and Woodruff (2007). Thus, a translog gross output (Y) production function of the following form is estimated on manufacturing firms only:

$$\begin{aligned}\log(Y) \quad = \quad & (\alpha + \gamma\Omega) + \beta_K \log(K) + \beta_L \log(L) + \beta_M \log(M) + \\ & \beta_{KL} \log(K)\log(L) + \beta_{KM} \log(K)\log(M) + \beta_{LM} \log(L)\log(M) + \\ & \beta_{KK} \log(K)^2 + \beta_{LL} \log(L)^2 + \beta_{MM} \log(M)^2 + \epsilon\end{aligned} \tag{4}$$

where the β coefficients are the elasticities on the production factors: capital K, measured as fixed assets;[8] labor L, measured as employment; and intermediate materials M,

[7] Estimates of the key equations presented in the paper were made using this more limited dataset, and with it, all coefficients on JQ still remained significant, and of comparable magnitudes. Nevertheless, we choose to present the results with the broader dataset since they cover both years.

[8] Gross fixed capital formation is used to proxy fixed assets.

approximated by the difference between gross output and value added. The constant α is the average level of TFP and γ is the effect of state-level judicial quality Ω on TFP, which we instrument with indigenous population in 1900 and large scale crops in 1939, as above. The error term ϵ captures the residual TFP. We similarly estimate a translog value added production function, replacing Y with value added, with intermediate materials M in the equation set to unity, so that only capital and labor inputs remain.

Table 6: Production function estimates with judicial quality
2008 data only

VARIABLES	(1) Value added production function	(2) Gross output production function
Judicial quality	0.727***	0.226**
	(0.200)	(0.107)
$\log(K)$	0.062	0.218***
	(0.078)	(0.083)
$\log(L)$	1.383***	0.415***
	(0.118)	(0.094)
$\log(K)\times\log(L)$	-0.019	0.004
	(0.020)	(0.013)
$\log(M)$		0.523***
		(0.113)
$\log(K)\times\log(M)$		-0.045***
		(0.014)
$\log(L)\times\log(M)$		-0.027
		(0.020)
$\log(K)^2$	0.030***	0.021***
	(0.007)	(0.006)
$\log(L)^2$	-0.039**	0.006
	(0.017)	(0.013)
$\log(M)^2$		0.035***
		(0.010)
Constant	1.633***	0.941**
	(0.421)	(0.385)
Observations	594	605
R-squared	0.896	0.975

/a Instrumental variable estimates.
/b Robust standard errors shown in parentheses, clustered by state.
*** p<0.01, ** p<0.05, * p<0.1

The results, shown in Table 6, reveal a strongly positive effect on TFP of differences in state-level judicial quality. The estimated size of the effects of judicial quality in value added (column 1) and gross output (column 2) specifications – of about 2% of GDP per step of the judicial quality score – are roughly equivalent since the ratio between the coefficients is similar to the average ratio (in the dataset) of gross output to value added of three (3.12).

The gross output specification is usually preferred in industry-level data, and it seems to be more robust here (column 2), both in terms of its R-squared as well as its significant capital coefficients, that are of a similar magnitude to what is commonly found in emerging country microdata.

5 Conclusion

This paper has found a robust relationship between the increasing quality of the legal system and higher average firm size in Mexico, strongly supporting versions of the Lucas (1978) model of firm size that incorporate contractual uncertainty in investment decisions. These effects are estimated using bin-level census data and state-level measures of judicial quality observed over a period of five years using panel estimates, bolstering previous results that relied on cross-sections only. The findings are strengthened with the inclusion of geographic controls, historical instruments and the use of alternative measures of firm size and judicial quality. Moreover, evidence is found that firms in capital-intensive industries are affected the most by lower judicial quality. This is consistent with hold-up problems in contract enforcement limiting investment in states and regions that lack an adequate quality legal system.

The size of the estimated effects on firm size and efficiency are substantial. Moving from worst to best-practice judicial quality is estimated to be able to boost average firm size by two-thirds, widen the dispersion of its distribution, and boost the weakest states' GDP by as much as 8% through higher TFP, and possibly more if there are dynamic effects.

In future work, we would also like to be able to decompose judicial quality further, and identify in more detail the effects of specific reform channels. This will require greater data availability about judicial processes and the implementation of recent systemic reforms.

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A Appendix

Table A.1: Indicators of judicial quality for contract enforcement

	Ejecutabilidad Contractual ("EC")				
FEDERAL ENTITY	2001	2003	2006	2008	2011
Aguascalientes	EC1	EC1	EC2+	EC3	EC2
Baja California	EC2+	EC1	EC2+	EC1	EC1
Baja California Sur	EC3	EC4+	EC4	EC2+	EC3
Campeche	EC4+	EC2	EC3	EC5	EC2
Chiapas	EC3	EC3	EC3+	EC4+	EC4
Chihuahua	EC5	EC4+	EC4	EC5	EC4
Coahuila	EC2	EC3+	EC2+	EC2	EC2
Colima	EC2+	EC1	EC2+	EC2	EC2
Distrito Federal	EC3+	EC2+	EC3+	EC4+	EC2
Durango	EC4+	EC2+	EC2+	EC3+	EC2
Guanajuato	EC3+	EC2	EC1	EC1	EC2
Guerrero	EC5	EC4+	EC3	EC4+	EC3
Hidalgo	EC4+	EC3	EC4	EC3+	EC3
Jalisco	EC2	EC3	EC3	EC3+	EC3
México (Estado de)	EC2	EC1	EC1	EC3+	EC2
Michoacán	EC5	EC3+	EC4+	EC4+	EC3
Morelos	EC4+	EC2	EC4	EC3+	EC3
Nayarit	EC4+	EC4	EC2	EC4	EC1
Nuevo León	EC1	EC1	EC1	EC2+	EC3
Oaxaca	EC3	EC2	EC3	EC3+	EC4
Puebla	EC5	EC5	EC4	EC5	EC3
Querétaro	EC1	EC1	EC1	EC3	EC4
Quintana Roo	EC1	EC2	EC4+	EC2	EC5
San Luis Potosí	EC2	EC2	EC3	EC2+	EC3
Sinaloa	EC3	EC4+	EC3+	EC1	EC2
Sonora	EC5	EC5	EC4+	EC4+	EC3
Tabasco	EC2	EC2	EC3+	EC2	EC3
Tamaulipas	EC2+	EC3	EC2+	EC4	EC2
Tlaxcala	EC5	EC5	EC4	EC3	EC5
Veracruz	EC5	EC5	EC5	EC5	EC4
Yucatán	EC4+	EC2	EC3	EC4	EC5
Zacatecas	EC4	EC2+	EC4	EC4	EC5

Source: Moody's (2011) and its surveys with ITAM and GMA.

Table A.2: Interactions with capital intensity and vertical integration
Single-plant domestic firms – 2008 data only

VARIABLES	(1) Preferred 2008 eqn.	(2) Without IV	(3) Capital intensity	(4) only X term	(5) Vertical integration	(6) only X term
Judicial quality	2.567***	0.837***	-0.290		0.457	
	(0.535)	(0.268)	(0.524)		(0.365)	
EC X Capital intensity			0.219**	0.026		
			(0.085)	(0.083)		
EC X Vertical integration					0.362	0.096
					(0.252)	(0.239)
Market size	1.014***	0.920***	0.881***	0.871***	0.900***	0.873***
	(0.100)	(0.126)	(0.121)	(0.039)	(0.125)	(0.027)
Distance to int'l markets	-0.050	-0.034	-0.025	0.216***	-0.039	0.215***
	(0.058)	(0.085)	(0.080)	(0.015)	(0.084)	(0.013)
Constant	-14.687***	-11.616***	-11.019***	-11.283***	-11.312***	-11.278***
	(1.668)	(2.055)	(1.991)	(0.327)	(2.031)	(0.334)
Observations	541	541	541	541	541	541
R-squared	0.345	0.309	0.326	0.446	0.314	0.446
Instrumented	Yes	No	No	No	No	No
State dummies	No	No	No	Yes	No	No
Industry dummies	Yes	Yes	Yes	Yes	Yes	Yes

Robust standard errors in parentheses, clustered by state. *** p<0.01, ** p<0.05, * p<0.1