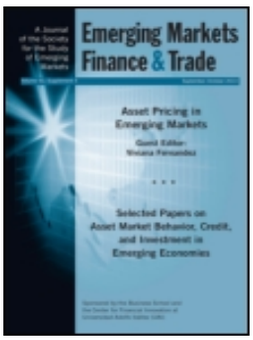

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Dollarization, Foreign Ownership, and Competition in the Banking Industry in Latin America

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ABSTRACT: We estimate the correlation of foreign bank penetration and dollarization with competition in the banking industry in sixteen Latin American countries during the period 1995–2008. We apply Boone's methodology to compute the intensity of competition. Our results suggest that in countries with an initial low level of competition, foreign ownership tends to foster rivalry among banks, whereas the opposite is true in countries with an initial high level of competition. The adoption of dollarization or a currency board, which reduces transaction costs and facilitates financial integration, has a positive correlation with competition. This is the case for Ecuador, El Salvador, and Argentina.

KEY WORDS: banking, competition, dollarization, foreign investment

Introduction

A competitive financial sector is a crucial component of market economies. Competition in the banking sector allows consumers to smooth consumption and firms to finance their investment at a lower cost. Most of the literature supports a positive relationship between competition and productivity. Under a more competitive environment, firms have higher incentives to innovate and to reduce managerial slack. In the banking industry, the effect of competition also works through a different channel since banks are a primary source of credit for firms. Prompt access to credit funds, as an indicator of competition, allows more firms to finance innovation in process and product.[1]

The purpose of this article is to reassess the relationships among foreign ownership, dollarization, and competition in the banking industry. Using the Fitch-IBCA Bankscope data set that provides bank-level annual financial information, we apply Boone's (2008) methodology to compute the intensity of competition in the banking sector over time in sixteen countries in Latin America. According to Boone's technique, the relationship between efficiency and profits should be stronger in more competitive markets. The data allow us to run a difference in difference model to test the relationship between foreign bank penetration and dollarization/currency board on the intensity of competition in the banking sector. During the '80s and '90, the Latin American banking industry has gone through important changes. First, consolidation has led to a more concentrated market. Second, several countries have experienced significant increases in foreign bank penetration, as reported by Yeyati and Micco (2007). Finally, some countries, Argentina (1991–2001), Ecuador (2000–), and El Salvador (2001–) have undertaken more radical reforms. They dollarized or introduced a currency board. We use these changes on foreign bank penetration and dollarization to study their correlation with competition.

We find that foreign banks tend to increase market rivalry in countries with an initial low level of competition, whereas the opposite occurs in countries with an initial high level of competition. In our sample, the adoption of dollarization or currency boards has a positive correlation with competition. This is the case for Ecuador, El Salvador, and Argentina.

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On the measurement of competition, we employ the methodology proposed by Boone (2008). In this method, the relationship between efficiency and profits/revenues should be stronger in more-competitive markets. The Boone indicator is based on the notion that more-efficient firms, with lower marginal costs, gain higher market shares or profits in relation to their less-efficient rivals. As competition becomes stronger, there is a reallocation of output from less-efficient to more-efficient firms.

Levine (1996) provides a conceptual framework to analyze the potential costs and benefits of foreign bank entry. On the benefits side, he emphasizes how foreign banks can play a useful role in promoting capital inflows and competition. Foreign entry may foster competition by different channels. Foreign banks may have superior access to funds from aboard. Holding a more diversified portfolio may also induce banks to a more aggressive lending policy. Finally, the entry of a foreign bank, without ties or relationship with local firms, may destabilize collusive agreements among domestic banks. Concerns about foreign banks are associated with the risk of capital outflows increasing the economic volatility. Other possible negative effects of the presence of foreign banks are the crowding out of domestic banks and the greater difficulty of the government to lead the economy.

Hawkins and Mihaljek (2001) claim that global market and technological developments, macroeconomic pressures, and banking crises in the 1990s in developing countries have forced the banking industry and regulators to change their old way of doing business. These trends have pushed authorities to deregulate the banking industry at the national level and to open up financial markets to foreign competition. As a result, borders between financial products, banks, and nonbank financial institutions and the geographical locations of financial institutions have started to break down. These changes have significantly increased competitive pressures on banks in the emerging economies and have led to deep changes in the structure of the banking industry. Claessens and Laeven (2004) use Panzar and Rosse's (1987) methodology and find that greater foreign bank presence and fewer activity restrictions in the banking industry render a more competitive banking system.

A high degree of financial integration, which comes not only through foreign banks but also through cross border lending, enhances competition among financial institutions, as well as among financial market infrastructures, and reduces the costs of financial intermediation. Using a theoretical model, Arellano and Heathcote (2010) claim that dollarization lowers transaction and information costs, encouraging trade and financial integration. Berg and Borensztein (2000) claim that one of the most profound effects of Panama's dollarization is the close integration of its banking system with that of the United States and indeed with the rest of the world, particularly since a major liberalization in 1969–70. Dollarization expands the array of financial options open to emerging market governments and firms and should therefore increase competition.

The data allow us to estimate the Boone indicator for each Latin American country. Also, using a difference in difference approach, we test how foreign bank penetration and dollarization affect competition in these economies. For the 1995–2008 period, our results show that since the dollarization in the year 2000, Ecuador has been the most competitive banking system in our sample while Costa Rica, Honduras, and Venezuela have been the least competitive. In countries with a higher initial level of competition, foreign penetration appears to have led to a less competitive industry. Following Yeyati and Micco (2007), we argue that foreign banks in more-competitive countries increase the degree of product differentiation to reap oligopolistic rents. In countries with a low initial competition level, foreign penetration improves competition. In less-competitive countries, foreign banks can steal rents from domestic banks just by being more aggressive in their prices.

We present evidence that dollarization and currency boards are positively correlated with competition in the banking industry. This is the case for countries that implemented such reforms: Argentina, Ecuador, and El Salvador. Our results are in line with the idea that dollarization reduces transaction costs, increases financial integration, and therefore offers firms and households more financial options.

The contribution of this article to the current literature is twofold. This article is the first that applies a new measure of competition, the Boone indicator, to a set of banks in sixteen Latin American countries, and it focuses on the evolution of this indicator within countries. Second, and more important, using the evolution of the Boone indicator, this article provides new insights about the

suggestive correlation between foreign penetration and dollarization on competition in the commercial banking system in Latin American countries.

Measuring Competition in the Banking Industry

The empirical literature in industrial organization supports the use of structural models to analyze competition in the banking sector. Panzar and Rosse (1987) develop a methodology that measures how changes in input prices are reflected in revenues earned by banks. The authors propose a parameter dubbed the "H-statistic," which is defined as the sum of the elasticities of the reduced-form revenue function with respect to factor prices. Under perfect competition, H = 1 since any increase or reduction in production costs will be passed through revenues on a one-to-one basis. In the case of monopoly, the H-statistic is negative since the monopolist reacts by reducing output and revenues following an increase in the input price. In the intermediate case of monopolistic or imperfect competition, the value of H will range between zero and one. Claessens and Laeven (2004) apply this methodology to a set of banks from fifty countries during the period 1994–2001. They find that most of the banking markets behave according to a monopolistic competition model, obtaining an H–statistic that varies between 0.6 and 0.8. Yeyati and Micco (2007) obtain H-statistic values ranging between 0.50 and 0.87 for a set of seven Latin American countries during the period 1993–2001.

Bresnahan (1982) introduces conjectural variation models to characterize the level of competition. In these models, the conjectural coefficient represents the reaction of rivals with respect to the level of output chosen by a particular firm. A negative parameter—where one firm expects the other to reduce its output—corresponds to a competitive scenario since it induces firms to place more output in the market. Inversely, a positive value signals a less competitive market since firms will be reluctant to increase their output due to their rival's reaction. Examples of conjectural variation methods applied to the banking industry are the works of Berg and Kim (1998) in Norway, Shaffer (1989) in the United States and Shaffer (1993) in Canada. In Latin America, Spiller and Favaro (1984) employ this methodology to estimate how competition in the Uruguayan market is affected by the entry of foreign firms.

Boone (2008) proposes a novel methodology that infers the degree of competition in a market from the relationship between profits and efficiency of firms.

Under a more competitive market, more-efficient firms get a higher market share and higher profits relative to less-efficient firms. This reallocation of output from less- to more-efficient firms, which enlarges the difference in market shares between them, is called the selection effect of competition as explained by Aghion and Schankerman (2004). Thus, in a scenario of strong competition, such as Bertrand with homogenous product, the efficient firm gets 100 percent of the market and obtains positive profits, whereas the inefficient firm gets 0 percent of the market share and profits. It is a result of the Bertrand model with homogenous product and asymmetric firms in term of cost. On the contrary, under a less competitive scenario such as Cournot, an inefficient firm gets a strictly positive market share and profits but at a lower magnitude than those of the efficient firm. It is a result of the Cournot model with homogenous product and asymmetric firms in term of cost. The Boone method is representative of the efficiency hypothesis, which predicts that a firm's profitability is driven by its own efficiency. In the banking industry, the Boone (2008) technique has been applied by Schaeck and Cihák (2010) for banks in Europe and the United States and by Van Leuvensteijn et al. (2007) for European countries. In Latin America, Oda and Silva (2010) and Paz (2009) have measured competition using this method for banking markets in Peru and Chile.

The Data

Our main source of data is the Fitch-IBCA Bankscope (BSC) data set. Our data set covers 516 commercial banks with lending activities and deposits in sixteen Latin American Countries: Argentina, Bolivia, Brazil, Chile, Colombia, Costa Rica, Dominican Republic, Ecuador, El Salvador, Guatemala, Honduras, Mexico, Paraguay, Peru, Uruguay, and Venezuela (see Table 1). The unbalanced panel data set has 4.526 bank-year observations of all commercial banks in the Fitch-IBCA Bankscope.

Table 1. Number of commercial banks by country and year

Country	1995	1996	1997	1998	1999	2000	2001	2002	2003	2004	2005	2006	2007	2008	Total
Argentina	34	37	35	46	47	47	47	42	41	36	37	45	44	48	586
Bolivia	8	9	9	9	9	9	9	9	9	9	9	8	9	9	124
Brazil	74	82	86	86	88	98	99	98	76	71	73	77	79	38	1,125
Chile	21	21	22	21	20	21	20	19	19	20	20	21	21	18	284
Colombia	10	10	11	12	13	13	13	14	15	15	15	14	14	13	182
Costa Rica	0	0	0	11	9	12	11	12	12	12	12	12	12	5	120
Dominican Republic	7	7	7	7	9	18	17	21	17	22	21	20	17	16	206
Ecuador	0	0	27	28	22	19	19	21	21	20	20	21	18	19	255
El Salvador	0	0	0	0	0	7	8	8	8	9	9	9	9	9	76
Guatemala	0	0	0	0	24	22	20	21	21	19	19	17	15	12	190
Honduras	6	8	6	10	11	14	15	16	16	15	16	16	17	12	178
Mexico	14	19	20	26	26	29	29	31	29	30	26	31	28	14	352
Paraguay	6	5	8	15	14	15	14	13	12	12	12	12	12	11	161
Peru	10	12	12	12	13	13	11	11	11	12	11	12	11	12	163
Uruguay	4	3	5	8	9	24	25	26	24	19	19	23	17	0	206
Venezuela	7	8	9	13	33	34	34	32	30	29	24	28	29	8	318
Total	201	221	257	304	347	395	391	394	361	350	343	366	352	244	4,526

Source. Fitch-IBCA Bankscope data set.

Table 2 provides a brief description of the main variables related to the commercial activities of banks such as revenue market share, foreign market share, and variable cost. Revenue is the sum of interest income, commission income, fee income, trading income, and other operating income. Variable cost includes overhead costs, interest expense, and commission expense. The average market share in our sample is 7 percent, ranging from 1.2 percent in Brazil to 15 percent in El Salvador. El Salvador, Paraguay, and Uruguay present the largest share of foreign penetration in the data (over 80 percent).

Considering this panel data set, in the next section we are going to describe the empirical model used to measure the competition in the banking sector through the Boone indicator.

The Empirical Model

Boone et al. (2005) estimate the following equation[2]:

$$Ln(\pi) = \alpha + \beta AVC_i + \in_i,$$

where profits (π) are revenues minus labor and intermediate costs (variable cost) and (AVC_i) represent variable costs share. The latter is a proxy for marginal cost in case of constant return to scale $(c_i q_i / p_i q_i = c_i / p_i = 1 - \textit{profit margin})$. Empirical studies focusing on developed and Latin American countries find that economies of scale in the banking sector are exhausted at a relatively small size and then remain constant; therefore, we can use average costs as a proxy for marginal cost[3]. More-efficient firms have higher profit margins and therefore lower values of c_i/p_i.

Table 2. Mean values of main variables by country

Country name	Average market share per bank (percent)	Average market share of foreign bank (percent)	Standard deviation market share of foreign bank (percent)	Total variable cost over total income (percent)	Wages over total income (percent)
Argentina	2.4	60.3	11.7	73.2	12.2
Bolivia	11.3	25.5	3.8	74.5	13.9
Brazil	1.2	36.3	5.2	84.1	13.0
Chile	4.9	48.4	9.0	65.7	11.7
Colombia	7.7	18.8	5.4	64.7	12.3
Costa Rica	10.7	37.8	5.0	76.3	16.5
Dominican Republic	6.8	17.2	4.0	75.9	16.2
Ecuador	5.1	2.9	1.8	70.2	12.6
El Salvador	14.7	81.0	20.7	75.0	14.8
Guatemala	6.3	12.6	7.4	83.6	
Honduras	7.9	33.3	6.6	66.4	15.1
Mexico	4.0	60.5	21.0	72.4	2.9
Paraguay	8.7	86.3	5.7	89.8	8.9
Peru	8.6	58.4	5.6	61.9	15.8
Uruguay	6.3	89.7	11.5	87.9	8.9
Venezuela	4.4	23.7	7.8	71.9	17.0
Average	6.9	43.3	8.3	74.6	12.8

Source: Fitch-IBCA Bankscope data set.

The parameter β is the elasticity of a firm's profits with respect to its cost level. A higher value of this profit elasticity, in absolute value, signals more intense competition. Van Leuvensteijn et al. (2007) work with revenue-based market shares instead of profits. Market share is always positive, whereas profit could be negative, creating a sample bias. In our econometric exercises, we use market share as our dependent variable.

We compute the Boone estimator (β), which is defined as the elasticity of bank market share with respect to the proxy of marginal cost (variable costs share).

Our empirical study is conducted in three steps. In the first step, using the pooled sample of banks, we compute one Boone estimator (β_j) per country. We use (β_j) to compare competition across countries. In the second step, we compute one Boone indicator per country-year (β_{jt}) to describe the evolution of competition within countries over time. Finally, in the third step, we estimate a Boone indicator that is a linear function of foreign bank penetration and dollarization. In all cases, we exploit the panel structure of the pooled data of commercial banks.

Our main specification allows for a time-varying Boone indicator at the country level. Changes over time are key to the purposes of our article. From a methodological perspective, the Boone indicator depends on industry-specific characteristics. Thus, by extension, the degree to which cross-national variations reveal differences in competition is not straightforward. As a result, a simple cross-national comparison is likely to lead to misleading conclusions unless we control for other country-specific characteristics. To avoid this omitted variable problem, we use a difference in difference approach. Within-country variation provides useful information about the evolution of competition and its determinants.

Degree of Competition Across Countries

To compare the degree of competition across countries, we compute the elasticity of a firm's revenues with respect to its cost level for each country over the entire period 1995–2008 (β_j)[4]

$$Ln(share_{ijt}) = \alpha + \beta_j d_j \ln(mc_{ijt}) + \eta_{jt} + \eta_i + \varepsilon_{\text{ijt},} \quad t = 1, .., \text{T } i = 1, \ldots, \text{N},$$

where *i, j*, and *t* refer to bank, country, and year, respectively; *dj* is a dummy that takes a value of one for country *j*; and η_{jt} and η_i are country-time and firm fixed-effects. $Share_{it}$ is the market share of the revenue; *mcit* is a proxy of efficiency. We use variable cost over total revenue of bank as a proxy for marginal costs. We include country-time dummies (η_{jt}) to control for country-specific shocks; ε_{ijt} is a random disturbance term.

Evolution of Competition Within Countries

To analyze how competition has changed over time in each country, we allow the elasticity to vary over time (1995–2008)(β_{jt})[5]:

$$Ln(share_{ijt}) = \alpha + \beta_{jt} d_j \ln(mc_{ijt}) + \eta_{jt} + \eta_i + \varepsilon_{\text{ijt},} \quad t = 1, .., \text{T } i = 1, \ldots, \text{N}$$

Foreign Penetration and Dollarization

Finally, to test the correlations of foreign penetration and dollarization on competition, we estimate the following equation:

$$Ln(share_{ijt}) = \alpha + (\beta_j + \gamma_j x_{jt}) \ln(mc_{ijt}) + \eta_{jt} + \eta_i + \varepsilon_{\text{ijt},} \quad t = 1, .., \text{T i} = 1, \ldots, \text{N},$$

where x_{jt} includes market share of foreign banks and a dummy that takes a value one if country *j* is dollarized in year *t*.

If γ_j is positive and statistically significant at the conventional level, foreign penetration and/or dollarization are negatively correlated with the intensity of competition.

In this estimation, it is important to consider that we do not have a causal effect of the impact of these policy changes on the effects of competition on outcomes. The results suggest correlations because this event could be correlated with another phenomena in the same period.

Endogeneity

Previous models have endogeneity problems between marginal costs and market share. Endogeneity could come from three sources: (1) The variable of interest, *ln(mcijt)*, is related to the dependent variable we use (market share [revenue]). Thus, β would be biased downward. (2) The presence of time-invariant unobserved heterogeneity across banks (fixed effects (*ηi*)): quality of management due to innate abilities and business experience, which is constant in time but different across banks. (3) Time-variant shocks (*ξ ijt*): investment in new technology, causing the bank to gain higher market shares and decrease the marginal costs, or investment in quality, which can increase market shares but at the same time increases marginal cost (Hay and Liu 1997).

The error term of the model, *εijt*, could be written as

$$\varepsilon_{ijt} = \eta_i + \xi_{ijt}$$

If we assume that $corr(\eta_i, mc_{ijt}) = 0$ and $corr(\xi_{ijt}, mc_{ijt}) = 0$, we can estimate the models using random effect and obtain a consistent estimator of the intensity of competition, but if $corr(\eta_i, mc_{ijt}) \neq 0$ and $corr(\xi_{ijt}, mc_{ijt}) = 0$, we need to account for bank fixed effects.[6] In the case that $corr(\eta_i, mc_{ijt}) \neq 0$ and $corr(\xi_{ijt}, mc_{ijt}) \neq 0$, the literature uses the generalized method of moments (GMM) or the Anderson-Hsiao approach (which is a special case of the GMM), using lagged values of the explanatory variables as instruments.[7]

In order to estimate the model using GMM, one necessary condition is to have an instrument that, once controlled for the other covariates, is uncorrelated with the error term of the structural equation; that is, the instrument must be exogenous. Nonetheless, instrument exogeneity is not sufficient to identify the causal effect. The instrument must also be relevant; that is, it must be correlated with the endogenous variable. Instruments that fulfill both conditions are called valid. In some cases, the instrument is only weakly correlated with the endogenous variable, raising the problem of weak instruments. Although we can identify the causal effects, in the presence of weak instruments, inference can be misleading. With weak instruments, the sampling distribution of GMM coefficients is in general nonnormal; the endogeneity bias increases; and standard errors are unreliable. We check for weak instruments comparing the Kleibergen-Paap F statistic with the 95 percent confidence interval, the rule of thumb value of ten, and the critical values computed by Stock and Yogo (2005).[8]

As a benchmark, we first estimate Equation 2 using fixed effects. Next we apply the Anderson-Hsiao approach to take into account fixed effects and the potential correlation between our proxy for marginal costs and the error term $corr(\xi_{ijt}, mc_{ijt}) \neq 0$.

If the error term is uncorrelated with the marginal costs proxy, both models are unbiased although the fixed effects model is more efficient. If the error term is correlated, the fixed effects model is biased, in which case we have to use the Anderson-Hsiao model. For the latter, we see whether the whole model is underidentified.[9] We find a serious problem in this estimation. However, there are some countries in which the instruments are relevant.[10] For robustness, we test whether the fixed effects coefficients are equal to the Anderson-Hsiao coefficients in these cases. We cannot reject the null hypothesis that both sets of coefficients are equal.

Finally, if individual specific effects are uncorrelated with the independent variables, random effect is more efficient than fixed effect. We use Durbin-Wu-Hausman (DWH) to test the null hypothesis *corr(ηi,mcit)* = 0. We reject the null hypothesis at standard levels of significance and therefore cannot use the random effect model.

Estimation Results

Table 3 presents Boone estimators by country using bank, fixed effects (column 1), random effect (column 2), and Anderson-Hsiao (column 3) models. In all empirical models, we include country-time dummies. Under the null hypothesis that both bank idiosyncratic effects $corr(\eta_i, mc_{ijt}) \neq 0$ and marginal costs $corr(\xi_{ijt}, mc_{ijt}) \neq 0$ are correlated with the error term, only Anderson-Hsiao coefficients are unbiased.

Under the null hypothesis that only the bank idiosyncratic effects are correlated with the error term $corr(\eta_i, mc_{ijt}) \neq 0$, the Anderson-Hsiao and fixed effect coefficients are unbiased, although the fixed effect model is more efficient.

Finally, under the null that neither idiosyncratic effects nor marginal costs are correlated with the error term, all three models are unbiased, although the random effect is the most efficient model.

For the Anderson-Hsiao model, we first compute the *T*-statistic for the AR(1) and AR(2) of error terms in first differences. By construction, AR(1) should be significant. Serial correlation in the first-difference errors at an order higher than one implies that the moment conditions used by Anderson-Hsiao are not valid; in our sample, the *T*-statistic for the AR(2) is −0.41 and is not significantly different from zero. Next we test whether the correlation between the difference and the second lag of marginal cost in level is different from zero and whether this correlation is weak.[11] The Kleibergen-Paap underidentification LM statistic (0.65) and Wald *F*-test (0.01) show that the model as a whole is under- and weakly identified; therefore, the instruments may be inadequate to identify the equation.

In column 4, we present the *F*-test for each first stage result in the Anderson-Hsiao model, and we conclude that the instruments are correlated with the potential endogenous regressors in some countries. The DWH test in the Anderson-Hsiao model is 19.9; we cannot reject the null that difference of marginal cost may be treated as exogenous, but these results have to be included with caution because of the problem with the instruments. For robustness and due to the problems of underidentification and weak instruments in the Anderson-Hsiao model as a whole, in each country in which the correlation is different from zero and the instruments are not weak (column 5), we applied the Wu-Hausman *F*-test. In these cases we cannot reject the null that the fixed effects coefficient is equal to the one estimated using the Anderson-Hsiao model. The previous statement is true whether we use the 95 percent confidence interval for the standard *F*-test, the rule of thumb of a value of ten for the *F*-test, or the stricter critical value computed by Stock and Yogo to test whether we are in the presence of weak instruments (16.38 in our case). For example, in the case of Argentina, the Kleibergen-Paap *F*-statistic (22.28) rejects the null that we are using a weak instrument, and the Wu-Hausman *F*-statistic cannot reject the null that the fixed effect coefficient is statistically equal to the Anderson-Hsiao coefficient.

Table 3 also presents the DWH test for the null hypothesis that fixed effect (FE) coefficients (column 1) and random effect coefficients (column 2) are equal. In our sample, the DWH is 25.8; therefore, we reject the null.

Summing up, previous results suggest that FE is the best econometric technique to compute the Boone estimator in our sample. Once we account for FE, we do not find evidence of correlation between our proxy of marginal costs (mc_{ijt}) and the error term (ξ_{ijt}).

This result is consistent with the empirical literature cited in the Empirical Model section, which find that economies of scale in the banking sector are exhausted at a relatively small size and then remain constant, so once we control for bank and country-time fixed effect we do not have a reverse causality between marginal costs and market share. This could be explained by the fact that the marginal costs in the Latin American countries are mainly determined by the changes of international cost of funds.

Table 3. Boone statistics by country, 1995–2008

	Bank fixed effect	Bank random effect	Anderson-Hsiao	Test of excluded instruments	Kleibergen-Paap	Wu-Hausman
	Model	Model	Model	*F*-statistic +	*F*-statistic ++	*F*-statistic +++
Argentina	–1.05***	–1.05***	–0.83***	21.98	22.28	0.78
x Marg.Cost (ln)	(0.00)	(0.00)	(0.00)			
Brazil	–0.37**	–0.33**	0.08	10.22	10.61	0.17
x Marg.Cost (ln)	(0.02)	(0.04)	(0.85)			
Bolivia	–1.31***	–1.31***	–1.23*	0.33	0.30	0.02**
x Marg.Cost (ln)	(0.00)	(0.00)	(0.09)			
Chile	–1.05***	–1.04***	–0.981***	18.51	17.85	0.66
x Marg.Cost (ln)	(0.00)	(0.00)	(0.00)			
Colombia	–1.21***	–1.24***	1060	4.00	3.65	0.14
x Marg.Cost (ln)	(0.00)	(0.00)	(0.36)			
Costa Rica	–0.04	–0.09	–5.08	1.04	0.92	0.03**
x Marg.Cost (ln)	(0.94)	(0.85)	(0.25)			
Dominican Republic	–1.65***	–1.70***	–1.54	3.57	3.34	0.67
x Marg.Cost (ln)	(0.00)	(0.00)	(0.14)			
Ecuador	–1.93***	–1.91***	–1.15	2.02	1.94	0.84
x Marg.Cost (ln)	(0.00)	(0.00)	(0.26)			
El Salvador	–0.99**	–1.10***	–1.31**	15.99	13.19	0.65
x Marg.Cost (ln)	(0.01)	(0.01)	(0.03)			
Guatemala	–1.01**	–1.00**	–2.06	0.19	0.18	0.74
x Marg.Cost (ln)	(0.03)	(0.03)	(0.76)			
Honduras	0.49**	0.18	1.07**	6.93	6.32	0.20
x Marg.Cost (ln)	(0.00)	(0.57)	(0.02)			
Mexico	–1.29***	–1.36***	–1.80***	12.53	12.30	0.03**
x Marg.Cost (ln)	(0.00)	(0.00)	(0.00)			
Paraguay	–0.67**	–0.68**	0.43	7.20	6.43	0.49
x Marg.Cost (ln)	(0.04)	(0.04)	(0.70)			
Peru	–1.23***	–1.30***	–0.91***	37.02	33.52	0.57
x Marg.Cost (ln)	(0.00)	(0.00)	(0.00)			
Uruguay	–0.94***	–0.86***	0.98*	57.45	53.48	0.08*
x Marg.Cost (ln)	(0.00)	(0.00)	(0.07)			
Venezuela	–0.19	–0.19	0.37	14.38	14.52	0.18

x Marg.Cost (ln)	(0.58)	(0.58)	(0.52)
Observations	4,526	4,526	3,251
Durbin–Wu–Hausman test		25.8**	19.9
T-statistic for AR(1) in first differences			–2.59**
T-statistic for AR(2) in first differences			–0.28
Kleibergen-Paap rk LM statistic *p*-value			0.652
Kleibergen-Paap rk Wald *F*-statistic			0.012

Notes: Bank cluster standard errors. *P*-value in parentheses. All regressions include country-year dummies. *Significant at 10 percent; **significant at 5 percent; ***significant at 1 percent.
+To measure the instrument relevance.
++For weak instruments.
+++For difference in coefficients between fixed effect and Anderson-Hsiao models.

Degree of Competition Across Countries

Column 1 in Table 4 shows that Ecuador presents the highest Boone estimator (in absolute value), although its level is not statistically different from those computed for Argentina, Bolivia, Chile, Colombia, Dominican Republic, El Salvador, Guatemala, Mexico, and Peru (see Table A.1 in appendix). Honduras presents the lowest competition index in the sample. Its competition coefficient is significantly different from all other countries except Costa Rica and Venezuela at the standard confidence level.[12]

These results are in line with Micco et al. (2004), who find that Venezuela has the third-highest interest spreads in the world (18.3 percent), which could be the result of a very low degree of competition in the commercial banking sector.[13]

Van Leuvensteijn et al. (2007) compute Boone estimators for eight developed countries (France, Germany, Italy, Japan, Netherlands, Spain, United Kingdom, and United States) during the period 1994–2004. They find that Japan has the lowest competitive banking sector with a Boone estimator equal to -0.72 and that the United States has the most competitive market with an estimator of -5.41. In our sample of sixteen Latin American countries, no country reaches the level of competition existing in the United States.

Development of Competition Over Time

As mentioned in the Empirical Model section, cross-national comparisons can be misleading due to accounting and regulatory differences across countries. We henceforth center our analysis on results that can be inferred from the dynamic dimension.

Table 4 shows the Boone indicator across countries over time using bank fixed effects. Results at the country level, based on Table 3, are still valid although there are differences within countries over time. For each country, we reject the null hypothesis that Boone estimators are equal for a given country over time (see *F*-test). We also reject that the difference between the highest and lowest values of the Boone indicator for each country is statistically different from zero (see Max. Diff.).[14]

In Ecuador, competition has increased since 1999 (see Table 4). These results, initially surprising, may be explained by Ecuador's dollarization in 2000. As Quispe-Angoli and Whistler (2006) point out, official dollarization lowers transaction and information costs, encouraging trade and financial integration. A less opaque market increases competition. Table 4 shows that Ecuador's Boone estimator increases in absolute terms after 2000, confirming this hypothesis. Also, during the first half of the '90s, due to the 1999 financial crisis and the postcrisis improvement in regulation, total credit shrank in Ecuador. If lending contraction is done in small and opaque clients, competition should also increase.

Boone estimators for Brazil, Bolivia, Colombia, and El Salvador have decreased since 1999 implying an increase in competition. The results show that the level of competition in Mexico increased abruptly in 1996, after the Tequila crisis when the country implemented a large-scale banking reform. Competition continued to increase until 2000 then remained flat until 2007 when it decreased.

El Salvador, which dollarized the economy in 2001, presents a relatively low Boone estimator in spite of its level of development. Just after the dollarization, the Boone estimator was not significantly different from zero although it started to decrease over time. The Boone estimator for El Salvador became statistically different from zero in 2006, five years after dollarization.

Argentina, Chile, and the Dominican Republic experienced a decrease in competition during the period 1995–2000. After 2000, the Boone estimator increased in each of these countries. Competition in Peru fell during 1995 and 2002 and then increased until 2006. Argentina and Chile had on average the same Boone indicator during the whole period, but since 2000, our estimates show a higher increase in the level of competition in Chile. Competition fell in Argentina after 2000. Argentina experienced an economic and financial crisis in 1999. During the same year, the real gross domestic

Table 4. Boone estimator over time with banks' EE

Year	Argentina	Brazil	Bolivia	Chile	Colombia	Costa Rica	Ecuador	Dominican Republic	El Salvador	Guatemala	Honduras	Mexico	Paraguay	Peru	Uruguay	Venezuela
1995	−2.02**	0.20	−0.27	−1.43***	−0.90			−2.36***			1.34**	3.49***	1.44***	−0.22	−0.98	1.25
	(0.03)	(0.64)	(0.89)	(0.00)	(0.31)			(0.01)			(0.05)	(0.01)	(0.47)	(0.65)	(0.44)	(0.32)
1996	−1.27***	0.25	1.33	−1.31***	−0.26			−4.70***			0.67	−1.31***	−0.10	−2.05***	−2.58***	−0.13
	(0.01)	(0.48)	(0.57)	(0.00)	(0.71)			(0.01)			(0.16)	(0.02)	(0.95)	(0.00)	(0.01)	(0.75)
1997	−0.70	−0.35	−3.44	−1.22***	−0.16		−4.31***	0.07			−0.07	−0.85**	−0.78***	−1.26**	−1.75***	−3.08**
	(0.14)	(0.33)	(0.36)	(0.00)	(0.79)		(0.00)	(0.96)			(0.73)	(0.05)	(0.29)	(0.04)	(0.01)	(0.05)
1998	−1.25***	−0.45	1.74	−0.54*	−0.07	−0.79	−0.78	−2.78*			0.18	−1.28**	−0.97**	0.04	−1.38***	−2.58**
	(0.00)	(0.18)	(0.20)	(0.07)	(0.89)	(0.15)	(0.30)	(0.07)			(0.74)	(0.03)	(0.04)	(0.93)	(0.00)	(0.03)
1999	−1.17***	0.11	3.17***	−0.67***	−1.03*	−1.98***	−0.10	−3.98***		−1.92	0.57	−2.05***	−2.43***	−1.25*	−0.74**	−0.46
	(0.00)	(0.57)	(0.00)	(0.00)	(0.08)	(0.00)	(0.91)	(0.01)		(0.03)	(0.23)	(0.00)	(0.00)	(0.07)	(0.04)	(0.34)
2000	−0.66***	0.13	0.29	−0.52	−0.99**	−0.60	−1.06***	−1.20	−0.03	−1.12*	0.54	−1.33***	−1.24**	−0.40**	−0.74	−0.07
	(0.01)	(0.74)	(0.61)	(0.16)	(0.02)	(0.29)	(0.01)	(0.14)	(0.94)	(0.09)	(0.19)	(0.00)	(0.05)	(0.04)	(0.15)	(0.87)
2001	−0.73***	0.06	−0.48*	−1.11***	−1.12**	0.91	−2.05***	−0.92**	−0.17	−2.01***	0.73*	−1.31***	−2.22	−0.44	0.83	−0.72*
	(0.00)	(0.86)	(0.08)	(0.00)	(0.04)	(0.62)	(0.00)	(0.02)	(0.40)	(0.00)	(0.08)	(0.00)	(0.27)	(0.48)	(0.21)	(0.10)
2002	−0.81***	0.42	−0.80**	−0.91***	−0.94***	0.55	−1.69***	−1.94***	−0.01	−1.19	0.66	−1.92***	−2.54	−0.01	−2.25	−0.01
	(0.00)	(0.11)	(0.03)	(0.00)	(0.01)	(0.47)	(0.00)	(0.00)	(0.99)	(0.13)	(0.15)	(0.00)	(0.63)	(0.92)	(0.17)	(0.98)
2003	−0.91***	−0.35	−1.59***	−1.27***	−1.41***	1.17**	−2.46***	−1.30***	−0.62	−0.99	0.62	−1.37***	−12.21	−0.93***	1.01	0.66
	(0.00)	(0.51)	(0.01)	(0.00)	(0.00)	(0.03)	(0.00)	(0.00)	(0.45)	(0.11)	(0.17)	(0.00)	(0.12)	(0.01)	(0.29)	(0.34)
2004	−1.06***	−0.78*	−0.86***	−1.26***	−1.18**	−0.29	−3.04***	−1.25**	−0.57	−0.51	0.40	−0.81*	−2.4	−1.24**	0.05	1.25
	(0.00)	(0.07)	(0.00)	(0.00)	(0.02)	(0.62)	(0.00)	(0.04)	(0.29)	(0.5)	(0.62)	(0.09)	(0.23)	(0.02)	(0.91)	(0.21)
2005	−1.25***	−1.55***	−0.09	−1.44***	−1.08**	−0.16	−2.04**	−1.70***	−0.96**	−0.68	−0.99*	−0.91	−2.36**	−0.94***	0.82	0.22
	(0.00)	(0.00)	(0.92)	(0.00)	(0.04)	(0.64)	(0.02)	(0.00)	(0.04)	(0.37)	(0.09)	(0.18)	(0.02)	(0.00)	(0.38)	(0.71)
2006	−1.20***	−0.82	−1.96	−1.01***	−2.33**	0.40	−2.51***	−1.99***	−1.34***	−0.55	−0.82	−1.23***	−0.33***	−1.42	−0.93	0.82
	(0.00)	(0.14)	(0.23)	(0.00)	(0.01)	(0.57)	(0.01)	(0.00)	(0.00)	(0.63)	(0.19)	(0.00)	(0.00)	(0.00)	(0.02)	(0.22)
2007	−1.09***	−1.96***	−1.84***	−1.02***	−1.97**	1.06	−2.63***	−3.21***	−1.27*	0.04	0.07	−0.62**	−4.72**	−0.55*	−1.01***	1.30*
	(0.00)	(0.00)	(0.01)	(0.00)	(0.05)	(0.24)	(0.01)	(0.00)	(0.06)	(0.98)	(0.91)	(0.05)	(0.03)	(0.03)	(0.09)	(0.09)
2008	−1.41***	−0.14	−1.47***	−2.02***	−2.74*	0.93***	−3.45***	−3.07*	−1.28*	−0.98	−0.03	−0.45	1.33	−0.01		1.35**
	(0.00)	(0.85)	(0.00)	(0.00)	(0.10)	(0.01)	(0.00)	(0.08)	(0.06)	(0.11)	(0.93)	(0.31)	(0.82)	(0.99)		(0.04)
Observations								4,526								
Max. Diff. (Prob>F) +	0.03**	0.00***	0.027**	0.01***	0.06*	0.00***	0.01**	0.00***	0.00***	0.07*	0.01***	0.00***	0.06*	0.00***	0.00***	0.01***
F-Test (Prob>F) ++	0.00***	0.00***	0.00***	0.00***	0.00***	0.00***	0.00***	0.00***	0.00***	0.41	0.01***	0.00***	0.00***	0.00***	0.00***	0.00***

Notes: Bank Cluster standard errors. *P*-value in parentheses. All coefficients are estimated together. *Significant at 10 percent; **significant at 5 percent; ***significant at 1 percent.
+Max. Diff. is the *p*-value of the null hypothesis that the difference between the largest and smallest Boone estimators for a given country are equal.
++*F*-Test is the *p*-value of the null hypothesis that all Boone estimators for a given country are equal.

product dropped by 4 percent. The crisis caused the government's fall, a default on the country's foreign debt, widespread unemployment, riots, the rise of alternative currencies, and the end of the currency board (fixed exchange rate to the U.S. dollar) by the end of 2001. The economic and financial crisis may have caused the decrease in competition in 2000, but it remained low relative to previous years because of the country's abandonment of the currency board, which reduced transaction costs and allowed firms to access international financial markets.

Costa Rica, Uruguay, and Venezuela show a significantly increasing trend since 1997, indicating a decline in competition in the banking sector until 2008. This trend has changed in Uruguay in the last few years. Paraguay shows an increase in competition during the whole period.

Effects of Foreign Penetration and Dollarization

To study the effect of foreign bank penetration and dollarization on competition, Table 5 and 6 show the estimation of Equation (3) using bank FE.

Table 5 suggests that foreign penetration has a positive correlation with the intensity of competition in Brazil and Venezuela during the period; in contrast, foreign penetration has a negative correlation with the degree of competition in Argentina, Costa Rica, and El Salvador.[15] These results are in line with Micco et al.'s (2004) interpretation. In countries with a higher initial level of competition, foreign banks can increase the degree of product differentiation and reduce competition, but in countries where the initial level of competition is lower (Brazil and Venezuela), foreign banks can substitute national products and therefore increase the level of competition.[16] Our results show that dollarization has a positive correlation with competition in Argentina and Ecuador. In the case of El Salvador, results suggest that competition is positively correlated with dollarization although the effect is only significant at the 15 percent level.

Table 5. Competition, foreign penetration, and dollarization at the country level

	Country Dummy		Foreign Mkt. Share		Dollarization	
	x Marg.Cost (ln)		x Marg.Cost (ln)		x Marg.Cost (ln)	
Argentina	–2.53***	(0.00)	2.73**	(0.03)	–0.52**	(0.05)
Brazil	1.34	(0.13)	–4.58*	(0.06)		
Bolivia	2.32	(0.62)	–14.00	(0.46)		
Chile	–1.46***	(0.00)	0.68	(0.20)		
Colombia	–2.35*	(0.07)	6.61	(0.19)		
Costa Rica	–3.31**	(0.04)	9.21***	(0.00)		
Dominican Republic	–2.68*	(0.08)	6.02	(0.50)		
Ecuador	–1.02	(0.27)	25.5	(0.13)	–2.02**	(0.02)
El Salvador	–1.22***	(0.00)	1.15**	(0.04)	–0.66 ϕ	(0.11)
Guatemala	–1.68*	(0.00)	4.95	(0.28)		
Honduras	0.38	(0.45)	0.33	(0.84)		
Mexico	–0.82	(0.18)	–0.62	(0.40)		
Paraguay	6.63	(0.11)	–8.56*	(0.08)		
Peru	–1.29*	(0.64)	0.10	(0.98)		
Uruguay	–1.40**	(0.28)	0.48	(0.73)		
Venezuela	–1.86***	(0.01)	–8.54***	(0.02)		
Observation:	4,526					

Notes: Model: Bank fixed effect with time dummies. Coefficients are estimated together. Bank cluster standard errors. *P*-value in parentheses. ϕ Significant at 15 percent; *significant at 10 percent; **significant at 5 percent; ***significant at 1 percent.

Table 6. Boone estimator, and foreign penetration and dollarization

	Bank fixed effect	Bank fixed effect
	Model	Model
Argentina	–1,751***	–1.47***
x Marg.Cost (ln)	(0.198)	(0.376)
Brazil	1,408	1.13*
x Marg.Cost (ln)	(0.945)	(0.667)
Bolivia	–1,195***	–1,271***
x Marg.Cost (ln)	(0.238)	(0.260)
Chile	–1,853***	–2,015***
x Marg.Cost (ln)	(0.207)	(0.451)
Colombia	–1,128***	–1,105***
x Marg.Cost (ln)	(0.194)	(0.184)
Costa Rica	1,653	1,489**
x Marg.Cost (ln)	(1,102)	(0.672)
Dominican.Republic	–1,572***	–1.35***
x Marg.Cost (ln)	(0.278)	(0.226)
Ecuador	–1,603***	0.141
x Marg.Cost (ln)	(0.225)	(0.393)
El Salvador	–1,635***	–1,565***
x Marg.Cost (ln)	(0.257)	(0.475)
Guatemala	–1,191**	–0.796*
x Marg.Cost (ln)	(0.574)	(0.419)
Honduras	2.01**	1,347**
x Marg.Cost (ln)	(0.909)	(0.534)
Mexico	–0.943*	–0.611
x Marg.Cost (ln)	(0.488)	–0.538
Paraguay	3,397	1,967
x Marg.Cost (ln)	(2,346)	(1,438)
Peru	–0.973**	–0.943*
x Marg.Cost (ln)	(0.427)	(0.492)
Uruguay	–2,202***	–0.475
x Marg.Cost (ln)	(0.497)	–0.638
Venezuela	0.957	1,124**
x Marg.Cost (ln)	(0.689)	(0.450)
For.Mkt.Share (1)	–4,769*	–3,03*
x Marg.Cost (ln)	(2.722)	(1.659)
For.Mkt.Share Medium Competition (2)	4,314	4,272**
x Marg.Cost (ln)	(2,785)	(1.822)
For.Mkt.Share Low Competition (3)	6,108**	4,617**
x Marg.Cost (ln)	(2.743)	(1.779)
Dollarization	–0.368***	–0.66**
x Marg.Cost (ln)	(0.130)	(0.322)
Observations	4,526	4,526
Weight	—	Revenue country Market Share
test (1) + (2) = 0 (Prob>F)	0.44	0.10
test (1) + (3) = 0 (Prob>F)	0.00	0.01

Notes. All regressions include country-year dummies. Bank cluster standard errors. Standard errors in parentheses. *Significant at 10 percent; **significant at 5 percent; ***significant at 1 percent.

Table 6 implies that dollarization has the same effect in all countries (Argentina, Ecuador, and El Salvador) but allows for different correlations of foreign bank penetration in countries with different levels of competition. Using results from Table 3, we create three groups of countries to determine whether the strategies used by the foreign banks depend on the level of competition in each country. The first group is composed of Bolivia, Colombia, the Dominican Republic, Ecuador, Mexico, and Peru (high competition); the second group includes Argentina, Chile, El Salvador, Guatemala, and Uruguay (medium competition); and the third includes Brazil, Costa Rica, Honduras, Paraguay, and Venezuela (low competition). Regressions include country-year dummies.

As predicted by previous results, column 1 shows that for countries in the base group (Low Competition Countries, the omitted one), foreign bank penetration has a positive correlation with competition (reduces Boone coefficient). For countries in the second group (medium level of competition), foreign bank penetration has a negative correlation with competition. For countries in the last group (high level of competition), foreign bank penetration is negatively correlated with the level of competition too, which is significant at the standard confidence level. The coefficient for Dollarization is negative and significant at the standard level. These results confirm the hypothesis that countries that dollarized or implemented the currency board could increase competition in their banking system. Dollarization lowers transaction and information costs, encouraging financial integration, and therefore expands the array of financial options open to emerging market governments and firms, ultimately increasing competition.

The main variables of interest are the interaction of marginal costs (*ln*) with foreign bank penetration and dollarization. These last two variables only vary across country and year and do not change across banks within a country in a given year. For robustness, column 2 repeats the previous regression weighting by market share at the country level. In this regression, each country-year has the same weight. All previous results remain.

Summing up, while foreign penetration has had a different correlation with the intensity of competition across countries with different initial levels of competition, dollarization has a significant positive correlation on competition.

Conclusion

We obtained suggestive correlations between competition and two institutional variables of the banking industry in Latin American countries: foreign penetration and dollarization. We use a new methodology to compute the level of competition: the Boone indicator.

Our main results are inferred from time variation within countries of the Boone indicator. In our sample, foreign bank penetration has a positive correlation with the intensity of competition across countries. In countries with an initial low level of competition, foreign ownership spurs rivalry among banks, whereas the opposite is true for countries with an initial high level of competition.

The presence of foreign banks has a positive correlation with the intensity of competition in Brazil and Venezuela, which initially had less-competitive banking industries. In contrast, in El Salvador, which has an intermediate level of competition, foreign penetration has a negative correlation with the degree of competition. Following Yeyati and Micco (2007), we argue that foreign banks in more-competitive countries increase the degree of product differentiation to reap oligopolistic rents. On the contrary, in less competitive countries, foreign banks can steal rents from domestic firms just by being more aggressive in their prices.

We present evidence that dollarization and currency boards are positively correlated with competition in the banking industry. This is the case in countries that implemented such reforms: Argentina, Ecuador, and El Salvador. These results are in line with the idea that dollarization reduces transaction costs, increases financial integration, and extends the financial options from which firms and households can choose.

Comparing the degree of competition across countries, our results suggest that Ecuador, after dollarization, has the most competitive commercial banking sector among the Latin American

countries, although the Boone estimator is not statistically different from other countries. Costa Rica, Honduras, and Venezuela present the lowest levels of competition in our sample.

Comparing our results with Van Leuvensteijn et al. (2007), we conclude that no country in Latin America reaches the level of competition existing in the United States. In addition, the least competitive country in Latin America has a much lower level of competition in comparison with Japan, which has one of the least competitive banking sectors in Leuvensteijn et al.'s (2007) sample.

However, as noted in the empirical literature, cross-national comparisons can be misleading due to accounting and regulatory differences, among others reasons, across countries.

Notes

1. See Boldrin and Levine (2008).
2. Instead of using the relation between profits of firm i and some reference j, they estimate log profits. This is equivalent to estimating the relative profits because using the log profits, the reference profit is absorbed into the constant term. They use that because in practice, it is problematic to specify this reference profit.
3. For developed countries, Berger and Hannan (1998), Peristiani (1997), Rhoades (1998), and Shaffer (1993) find that cost scale economies are exhausted at around $10 billion in assets. For Latin America, Micco et al. (2004) shows that there are substantial scale economies for small banks that have less than $150 million in assets. However, banks that have between $150 million and $8 billion in assets have similar overhead costs, indicating that economies of scale are not at work for these banks.
4. Except for Costa Rica (1998–2008), Ecuador (1997–2008), Guatemala (1999–2008) and El Salvador (2000–2008).
5. Except for Costa Rica (1998–2008), Ecuador (1997–2008), Guatemala (1999–2008) and El Salvador (2000–2008).
6. Yeyati and Micco (2007).
7. See Arellano and Bond (1991); Schaeck and Cihák (2010); Van Leuvensteijn et al. (2007).
8. Staiger and Stock (1997) suggest an *F*-statistic less than ten is problematic and a value of five or less is a sign of extreme finite-sample bias.
9. Kleinberg-Paap LM Statistic.
10. Test of exclude instruments to measure the instrument relevance.
11. We compare test of exclude instruments Angrist-Pischke F statistic to measure the instrument relevance with Stock and Yogo test.
12. The differences are significant at the 10 percent level.
13. Interest rate spreads are measured as net interest income divided by the average of loans and deposits, 1995–2002.
14. The significance refers to the 90 percent level of confidence.
15. In all the other countries, the effect is positive but not significant, except in the case of Paraguay and Bolivia.
16. Kraft (2004) finds a positive effect on competition of foreign bank penetration in Croatia.

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Appendix

Boone Estimates per Country: Difference Test

	Bolivia	Brazil	Chile	Colombia	Costa Rica	Dominican Republic	Ecuador	El Salvador	Guatemala	Honduras	Mexico	Paraguay	Peru	Uruguay	Venezuela
Argentina	0.04***	0.00***	0.96	0.69	0.06*	0.08*	0.11	0.90	0.93	0.00***	0.13	0.25	0.61	0.67	0.01***
Bolivia		0.00***	0.07*	0.79	0.02**	0.34	0.26	0.45	0.51	0.00***	0.89	0.06*	0.82	0.15	0.00***
Brazil			0.00***	0.05*	0.55	0.00***	0.01**	0.14	0.18	0.00***	0.00***	0.40	0.02**	0.05**	0.63
Chile				0.71	0.06*	0.09*	0.11	0.89	0.92	0.00***	0.18	0.26	0.63	0.66	0.02**
Colombia					0.07*	0.39	0.28	0.71	0.74	0.00***	0.84	0.29	0.96	0.56	0.05**
Costa Rica						0.01***	0.01***	0.14	0.16	0.33	0.02**	0.31	0.06*	0.12	0.82
Dominican Republic							0.66	0.21	0.26	0.00***	0.32	0.04**	0.39	0.09*	0.00***
Ecuador								0.17	0.19	0.00***	0.25	0.05**	0.28	0.10*	0.01***
El Salvador									0.99	0.00***	0.49	0.52	0.66	0.90	0.12
Guatemala										0.00***	0.55	0.54	0.70	0.89	0.15
Honduras											0.00***	0.00***	0.00***	0.00***	0.08*
Mexico												0.08*	0.88	0.21	0.00***
Paraguay													0.24	0.50	0.31
Peru														0.49	0.03**
Uruguay															0.07*

Notes. *T*-test for the null that both countries have the same Boone Estimator. Coefficient from column 1 in Table 3. (Prob > t): *Significant at 10 percent; **significant at 5 percent; ***significant at 1 percent.