Working Paper No. 2010/52

Foreign Banks and Credit Volatility

The Case of Latin American Countries

Meriem Haouat,[1] Diego Nicolas Moccero,[2] and Ramiro Sosa Navarro[3]

May 2010

Abstract

Foreign bank presence has substantially increased in Latin America during the second half of the 1990s, which has prompted an intense debate on its banking and macroeconomic consequences. In this paper, we apply ARCH techniques to jointly estimate the impact of foreign bank presence on the level and volatility of real credit in a panel of eight Latin American countries, using quarterly data over the period 1995:1-2001:4. Results show that, together with financial development, foreign bank presence has contributed to reduce real credit volatility, improving the buffer shock function of the banking sector. This finding is consistent with the fact that foreign banks are typically well diversified institutions holding higher quality assets and having access to a broad set of liquidity sources.

Keywords: foreign banks; credit volatility; Latin America; panel data; ARCH techniques

JEL classification: C33, E51, G21

[1]ISERAM, Institut Supérieur Européen de Gestion (ISEG), 28, rue des Francs-Bourgeois 75003 Paris, e-mail: meriem.haouat@iseg.fr; [2]OECD Economics Department, 2, rue André-Pascal, 75775 Paris CEDEX 16, e-mail: diego.moccero@oecd.org; and [3]EPEE, University of Evry Val d'Essonne, Boulevard François Mitterrand, e-mail: Ramiro.sosanavarro@univ-evry.fr

This study has been prepared within the UNU-WIDER project on the Southern Engines of Global Growth, directed by Amelia Santos-Paulino and Guanghua Wan.

UNU-WIDER gratefully acknowledges the financial contributions to the research programme by the governments of Denmark (Royal Ministry of Foreign Affairs), Finland (Ministry for Foreign Affairs), Sweden (Swedish International Development Cooperation Agency—Sida) and the United Kingdom (Department for International Development).

ISSN 1798-7237 ISBN 978-92-9230-289-4

Tables and figures appear at the end of the paper.

UNU World Institute for Development Economics Research (UNU-WIDER)
Katajanokanlaituri 6 B, 00160 Helsinki, Finland

Typescript prepared by Janis Vehmaan-Kreula at UNU-WIDER

1 Introduction

Foreign bank entry into developing countries has exploded since the 1990s, a trend that had been dominated in the previous decade by cross-border lending activities. International banking institutions have expanded their presence in several emerging market economies by establishing branches and subsidiaries, favoured by the liberalization of external sectors and the embracement of market-friendly policy reforms, including deregulation and privatization of the banking sector. Two regions that have been very active in attracting foreign direct investment into the banking industry have been Eastern Europe and Central Asia, and Latin America and the Caribbean. In the Middle East and North Africa, foreign bank presence has increased at a generally slower pace, while it remained stagnant or even declined in South and East Asia and the Pacific regions (Figure 1).

This unprecedented internationalization of the banking sector has prompted a debate on the potential consequences for the recipient countries. Multinational banks are likely to introduce better practices, and better risk evaluation and information technologies, helping boost the efficiency and diversification of banking services in host countries (Levine 1997; Goldberg 2007). They are also thought to have the potential to ameliorate emerging countries' banking regulation and supervision (Goldberg 2007). Moreover, they may lead to greater competition in markets that had been captured by domestic institutions, which should reduce banking costs and increase the overall efficiency of the system (Berger and Hannan 1989; Hannan 1991; Claessens et al. 2001; Corvoisier and Gropp 2002; Evanoff and Ors 2002; Demirgüç-Kunt et al. 2003). However, foreign banks may limit lending to small and medium-size enterprises (Berger et al. 2001), by attracting mainly the wealthier customers and leaving the riskiest borrowers to local banks, then weakening domestic banks (Claessens and Jansen 2000; Barajas et al. 2000; Detragiache et al. 2006).

Against this background, the present paper uses aggregate data to investigate the impact of foreign bank presence on credit volatility in a panel of eight Latin American countries during the period 1995-2001. The literature on the internationalization of the banking sector and the implications for banking and macroeconomic stability is inconclusive in this respect. There are those who argue that foreign banks are more likely to be *fickle lenders*, since they have better access to alternative business opportunities than domestic banks (Galindo et al. 2005). Moreover, they could potentially import shocks from their home countries, then contributing to destabilize domestic banking systems (Goldberg 2002; Martinez Peria et al. 2005). However, foreign banks may increase the buffer shock function of the banking sector in case of negative shocks, because they are typically well-diversified institutions with access to a broader set of liquidity sources (Detragiache and Gupta 2006; De Haas and Van Lelyveld 2006). Multinational banks may also allow for a faster recapitalization of local banks after a crisis, as was the case following the Mexican, Brazilian and Argentinean banking crisis of the 1990s (Peek and Rosengren 2000a). Finally, due to superior risk evaluation systems, better screening devices and home country supervision, foreign banks are likely to hold better quality assets (Crystal et al. 2001), and have the potential to avoid capital flight in case of domestic shocks (De Haas and Van Lelyveld 2004; Peek and Rosengreen 2000a).

Most of the applied papers on the link between foreign banks and stability of different macroeconomic aggregates are grounded on econometric models that analyse the first conditional moment, i.e. the mean, of the dependent variable (Micco and Panizza 2006; Galindo et al. 2005; Dages et al. 2000). The only study that has tried to account for the second conditional moment of the data, i.e. volatility, has done so using two-step methods, which are known to be inefficient (Morgan and Strahan 2003). We deviate from the previous empirical literature in that we apply ARCH techniques to model jointly the first and second conditional moment of real domestic credit. In order to shed light on the issue of credit volatility, the ARCH equation is extended to include the degree of development of the banking sector and the internationalization of the banking system among a broader set of regressors. To the knowledge of the authors, this is the first time that such tools are used to analyse the impact of foreign bank presence on macroeconomic volatility.

The main findings of the paper regarding credit volatility are as follows: (i) perhaps not surprisingly, banking crisis steeped-up real credit volatility, with public and foreign banks having no discernible effect (positive or negative) during these stressful periods; (ii) deeper banking systems result in lower credit volatility, a finding that is coherent with Denizer et al. (2002) for other macroeconomic variables, such as consumption, investment and real GDP; (iii) stabilizing effects predominate in a such a way that foreign bank presence reduced credit volatility in our panel of eight Latin American countries over the period 1995-2001.

The reminder of the paper is organized as follows. The next section provides an overview of theoretical considerations underlining the link between foreign direct investment in the banking sector and macroeconomic volatility. Section III details the data construction process, and Section IV presents the ARCH econometric model used in the estimations. Panel unit root tests are exposed in Section V, while Section VI describes the estimation results. Section VII concludes.

2 Foreign banks and volatility: a review of the literature

The theoretical literature examining the link between foreign direct investment in the banking sector and macroeconomic stability is rather limited. To the knowledge of the authors, only two studies deal with such an issue: Morgan et al. (2004), who design a banking model with capital-constrained intermediaries, and Galindo et al. (2005), who use a portfolio model. In both cases, foreign bank presence has contradictory effects on volatility, depending on the type of shock hitting the economy.

Morgan et al. (2004) (MSR) extend to two countries the banking model originally developed by Holmström and Tirole (1997) (HT) for a closed economy. The HT approach emphasizes that the banking system may become a separate channel of shocks transmission, apart from macroeconomic spillovers, such as changes in the monetary stance or exchange rate policy. In the model, firms have to choose between two sources of financing: bank or investors capital, which are not perfectly substitutable. Bank capital is the most expensive, because banks provide not only loans but also valuable monitoring services. Banks monitor investment projects by firms, financed by other firms in the economy, the banks and investors. For this reason, bank capital is called ‘informed capital’, while investor capital (not subject to monitoring) is called ‘uninformed capital’. Capital-constrained banking intermediaries are cardinal in this set up, since firms depend on their collateral (or capital) value to first raise bank (‘informed’) capital, in order to be able to have access to the much cheaper investor (‘uninformed’) capital. Indeed, the banking system may become a main source of instability in the economy, since any shock on banks will have immediate real effects on economic activity.

MSR extend the HT model to include another country, and let bank capital to be freely distributed between countries, while the amount of firm capital in each individual country is fixed. This set up is appropriate to study how bank capital shocks (financial shocks) and firm capital shocks (real or collateral shocks) affect the distribution of bank capital between countries, under an internationally integrated banking system. Uninformed investors in both countries have access to a worldwide securities market, with a quasi-unlimited supply of investment opportunities (the securities market rate of return is exogenous, equal in both countries and independent of country-specific shocks).

To isolate the effect of banking integration on economic volatility, MSR compare the impact of bank and firm capital shocks over credit and investment, both under an integrated and a national banking system. Indeed, suppose there is a negative bank capital shock in one of the countries. The impact on the amount of uninformed and informed capital invested in the affected country is smaller when the banking system is internationally integrated. Bank (informed) capital declines less in an integrated system because after a negative bank capital shock, its rate of return increases, attracting bank capital from the unaffected country and buffering the negative initial effect. The smaller reduction in informed capital, induces in turn a smaller contraction in investor (uninformed) capital. Given that banks monitor firms, bank capital constitutes a signal that firms exploit in order to attract investor capital. Consequently, in the case of a negative bank-capital shock, a multinational banking system helps promoting the stability of total credit and investment in the economy.

In contrast, the negative impact of a firm collateral shock is amplified under a multinational system. As before, two mechanisms are at play. First, the lower value of firm collateral decreases the bank capital rate of return after a negative shock in an integrated banking system. Therefore, banks will prefer lending their mobile capital in the unaffected country, where the bank capital rate of return is higher, and firms are backed by better collateral. As a consequence, bank capital is reduced in comparison with a national system, because in this case bank capital is immobile. Negative firm-capital shocks end in turn in larger declines of investor capital in an integrated banking system. As before, the supply of uninformed capital depends on the firms’ ability to leverage bank capital, which is directly linked to firm collateral. Therefore, bank credit and investment suffer more in a multinational banking system when negative firm-collateral shocks hit the economy.

Galindo et al. (2005) (GMP) extend Pyle’s (1971) portfolio model to many countries, to examine the behaviour of well diversified banks across nations in case of shocks to the host country. In their theoretical model, banks in each country have deposits and assets. They show that credit from well diversified foreign banks will be more stable when liquidity shocks (i.e., shocks to funding costs) hit the economy. In fact, multinational banks have access to a global pool of liquidity, so they may be less sensitive to a rise in deposit interest rates than domestic banks. In contrast, foreign banks may react more aggressively in the case of opportunity shocks (i.e., shocks to expected returns), worsening the impact of globalization on banking stability in the host country. A worldwide diversified bank is able to rapidly withdraw investments from a host country when there is a decline in expected returns, reassigning the capital to that part of the world with better economic prospects.

All in all, from the theoretical literature it seems clear that the final effect of foreign financial institutions on macroeconomic volatility depends on the type of shocks hitting the economy. In the MSR model, banking integration reduces credit and investment volatility under bank-capital shocks, but it increases it under firm-collateral shocks. In the GMP model, foreign banks may bring rewards in terms of greater stability when funding costs are affected in the host country, but may cause instability in the face of host opportunity shocks. Of course, the overall impact of banking integration on volatility is an empirical question, not free from implementation difficulties. As such, it is very hard to identify and isolate the types of shocks discussed above. These caveats, coupled with data availability problems have lead researchers to focus attention on the statistical significance of aggregate measures of foreign bank presence. If banking integration is not significant, this means that the stabilizing and destabilizing effects compensate each other, while if it is negatively signed and statistically significant at conventional statistical levels, stabilizing effects predominate and foreign banks improve the buffer function of the financial system. Indeed, we will use ARCH techniques to model together the determinants of the level and volatility of credit, and testing for the significance of foreign bank presence as a regressor in the variance equation.

3 The data

Our sample of Latin American countries includes Argentina, Bolivia, Brazil, Chile, Colombia, Costa Rica, Mexico and Peru, selection based on data availability. Aggregate information is quarterly and spans the period 1995:1-2001:4, for which a balanced panel is available.[1] As a consequence, we have 28 quarters and 8 countries resulting in 224 observations. Banking information was kindly provided by the Inter-American Development Bank (IADB), while macroeconomic data are available from the IMF's International Financial Statistics and national sources (i.e., central banks and national institutes of statistics).

Banking data were built using balance sheets of local financial institutions that report to the appropriate regulatory agency, and consist of a measure of foreign and public bank presence, and concentration of the banking system. Each bank is classified according to capital ownership in public and private, domestic, regional or foreign. Foreign Banks are those with more than 50 per cent of the capital owned by a G10 country. Some of the countries in the sample also host banks from other Latin American countries (regional banks). As in Galindo et al. (2005), we treat them as domestic banks, because the authors find that they behave like domestic than global well-diversified banks. We consider state-owned banks to be those with most of the capital in hands of the government. Using this classification, the measure of foreign- and public-banks presence is defined as their respective credit shares in the whole system. The measure of credit considered is direct credit by banks to private and public, non-financial

1 While for most of the countries banking data were available on a monthly basis, for Chile and Mexico data were only available at a quarterly frequency. Excluding these countries from the sample would have entailed a major loss, since these countries have been very active in attracting foreign banks.

institutions. Finally, the degree of concentration in the system is captured by the share of credits granted by the three largest banks over the total stock of credit.

As mentioned in the introduction, foreign banks have been very active in Latin America over the estimation sample, allowing assessing their impact on credit behaviour. The data show that foreign direct investment (FDI) into the Latin American banking sector was mainly encouraged by the process of deregulation and privatization of this industry that took place during the 1990s. As a result, the share of foreign banks on total credit more than doubled in Argentina, Brazil, Chile, Mexico and Peru between 1995 and 2001 (Figure 2).[2] This is against a slight increase in banking concentration over the region, with the exception of Brazil and Chile, where concentration has been reduced (Figure 3).

Macroeconomic data deemed relevant for the analysis consist of seasonally-adjusted real credit to the private sector by banking institutions (the dependent variable), seasonally-adjusted real domestic GDP, seasonally-adjusted U.S. real GDP, the Federal Funds Rate, the domestic fiscal balance, the spread between lending and borrowing rates, the degree of financial development, the bilateral real exchange rate with the U.S., and measures of banking and currency crisis.

Quarterly real credit is computed as the nominal credit stock average over the three months, deflated by the CPI. The Federal Funds Rate is expressed as per cent per annum, while the fiscal balance is computed as the four-quarter rolling-sum of the headline central government balance, over nominal GDP. The spread uses interest rates for local currency operations (in per cent per annum), and is computed as the difference between the lending and deposit rates, as a ratio to the deposit rate. Financial development in the economy is expressed as the percentage of the stock of total (public plus private) credit over GDP. The bilateral real exchange rate with the U.S. is computed using market nominal exchange rates (in national currency per U.S. dollar) and seasonally-adjusted consumer price indices. The banking crisis variable is a dummy that takes the value of one for each quarter of the year in which there was a banking crisis. Systemic banking crisis for 93 countries over the period 1970-2000 are dated in Caprio et al. (2005), while for 2001 we used information provided by Carstens et al. (2004). Finally, a proxy for a currency crisis was constructed based on Frankel and Rose (1996). Indeed, our currency crisis index takes a value of one if depreciation in a given quarter is higher than 10 per cent, which is in turn at least five percentage points higher than the depreciation of the previous quarter (Table 1).

4 The econometric model

Our interest lies in identifying the impact of foreign bank presence on credit volatility, controlling for additional factors affecting the mean and conditional variance of credit. The ARCH family models are particularly suitable for this purpose, since they allow estimating jointly the determinants of both the first and the second conditional moment of the data. Surprisingly, none of the studies on the subject has made use of this methodology.

---

2 A similar pattern emerges when assets, instead of credit, are used to measure foreign participation.

The econometric model to be estimated consists of the following equations:

$$y_{it} = \beta_0 + \phi y_{it-1} + X_{it}^{'}\beta_1 + \mu_i + u_{it}, \;\; i = 1,...,N, \;\; t = 1,...,T \tag{1}$$

$$u_{it} = \sigma_{it}\eta_{it} \tag{2}$$

$$\sigma_{it}^2 = \exp\left(\lambda_{0i} + z_{it}^{'}\lambda_1\right) + \alpha u_{it-1}^2 \tag{3}$$

$$\eta_{it} \sim N(0,1) \tag{4}$$

In the mean equation (1), y_{it} is the dependent variable, μ_i is an individual, country-specific fixed effect, X_{it} is a vector of explanatory variables, β_0 denotes the constant term, β_1 is a set of coefficients, u_{it} is a disturbance term, and N and T are the number of cross-sectional units and time periods in the panel, respectively. Equation (2) states the country-specific shock to follow an ARCH process. In equation (3), the conditional variance, λ_{0i} is an individual, country-specific fixed effect, while Z_{it} is a vector of explanatory variables.[3] Together, equations (1)-(4) constitute a panel ARMAX-ARCH model, which is estimated by quasi-maximum-likelihood techniques.[4]

The vector of macroeconomic explanatory variables assumed to affect credit dynamics in equation (1) includes real domestic GDP, the U.S. GDP, the Federal Funds Rate, the domestic fiscal balance, the spread between lending and borrowing rates, the bilateral real exchange rate with the U.S., and a measure of currency crisis. The lagged dependent variable is introduced in the regression to control for persistence in the level of private credit.

We expect banking lending to be procyclical, both with respect to local and international economic activity.[5] A higher foreign GDP captures a more benign international context, then leading to higher credit in domestic markets. We also expect increases in the international cost of money, measured by the Federal Funds Rate, to lead to a less buoyant credit activity in domestic economies. In the same vein should play increases in the real exchange rate, since they can be considered as another component of international lending costs (Dornbusch 1983). When the real exchange rate depreciates, the repayment of foreign loans becomes more expensive, making them less attractive. This may be a potential important channel of credit contraction in domestic markets since banks in the region have been extensively issuing foreign debt as a way of funding domestic lending operations.

With respect to the fiscal balance, the banking sector has proved to be a large provider of governmental funding in the countries under consideration. It would then be expected that better public account positions would crowd-in private credit. A higher interest rate

---

3 As usual, country specific fixed effects in the mean and conditional variance equation are intended to capture differences in institutions, regulations, culture and other economic factors not accounted for the explanatory variables.

4 Note that we estimate an ARCH(1) instead of a higher order or GARCH(*p*) model in order to keep the model as parsimonious as possible. Of course, residual diagnostic tests will indicate whether or not this model specification is appropriate.

5 See Horvath et al. (2002) and Goldberg (2007) for a revision of the theoretical literature on procyclical credit behaviour.

spread should negatively affect credit, by making fewer projects economically viable. A priori, we might expect currency crisis episodes in a particular country to be accompanied by a decline in banking lending, since they are associated with general losses of confidence in the system and a retrenching in deposits.

The banking variables assumed to enter the mean equation are the degree of financial development, the presence of public and foreign banks in the system, and the level of concentration. The inclusion of financial development in the mean equation is intended to capture the extent of financial market imperfections (information asymmetries, monitoring capacity, etc.). Aghion et al. (1999) develop a macroeconomic model based on micro foundations which combines financial market imperfections and unequal access to investment opportunities. They show that economies with less developed financial systems will tend to be more volatile (justifying the inclusion of financial development in the variance equation – see below), and will experience slower growth. Regarding foreign bank presence, empirical evidence is inconclusive up to the impact of foreign banks on credit dynamism. While Crystal et al. (2002) have found a positive effect (with foreign banks exhibiting a more robust loan growth than their national counterparts in a sample of Latin American countries over the second half of the 1990s), Detragiache et al. (2006) show that in poorer countries a stronger foreign bank presence is robustly associated with a slower credit growth to the private sector.[6] Finally, a more concentrated banking system is expected to reduce credit dynamism, because dominant players have much at stake in the event of negative shocks, then reducing risk taking behaviour and credit growth (Morgan and Strahan 2003).

Both theory and evidence highlight the importance of including interaction terms in equation (1) to account for the potential asymmetric behaviour of public and foreign banks in the event of crisis and external and internal shocks. For example, Micco and Panizza (2006) find that state-owned banks may play a useful credit-smoothing role because their lending is less responsive to macroeconomic shocks than lending by private banks. Regarding foreign banks, the empirical literature gives conflicting results as to whether or not they exacerbate credit expansions and contractions. Some empirical studies analysing cross border lending from international banks to developing countries show a positive relationship between the host country business cycles and international lending behaviour. Indeed, multinational banks may reduce lending to host countries which face economic difficulties, to reallocate their capital over markets with better economic prospects (Dahl and Shrieves 1999; Buch 2000; Jenneau and Micu 2002; Morgan and Strahan 2003). However, Micco and Panizza (2006) find that foreign banks have not contributed to exacerbating lending cyclicality in a sample of developing and developed countries over the period 1995-2002. Studies on foreign banks behaviour during financial crisis in host countries point that, internationally-diversified and more capitalized foreign banks did not reduce credit supply in periods of financial distress, contributing to greater credit stability (Dages et al. 2000; Peek and Rosengren 2000a;

6 This could be the case because of a 'cherry picking' behaviour.

Crystal et al. 2002; and Goldberg 2002).[7] Foreign banks might view crises as an opportunity to expand, then increasing their local market share.[8]

An internationally integrated banking system may also shape the way external interest rates and GDP shocks affect lending behaviour. For example, in the event of a Federal Funds Rate increase, these institutions may exacerbate the 'flight to quality' effect. Regarding foreign GDP, the literature gives conflicting results on the impact on credit by foreign banks and cross border lending. On the one hand, authors like Moshiriam (2001), Martinez Peria et al. (2005) and De Haas and Van Lelyveld (2006) find a negative push relationship between the home country economic cycle and cross-border and foreign-bank lending in host countries. Parent banks can expand their activities in the host country when facing worsening economic conditions in their own market, as a way to seek for external lending opportunities and boost profits. On the other hand, Dahl and Shrieves (1999) and Peek and Rosengren (1997, 2000b) find evidence of a positive push relationship between the home country economic cycle and cross-border and foreign-bank lending in host countries. The authors argue that this may be the case because economic turmoil in the home country can lead to deterioration in the parent's bank financial conditions, making them reduce foreign subsidiaries' activities. Empirical evidence for the region seems to go in this direction, since Goldberg (2002) observes that U.S. bank claims to Latin American countries increase as the U.S. economy grows faster.

4.1 The potential determinants of credit volatility

Regarding the variables in the conditional variance equation (3), we include those that have been assumed by the literature to have a burden on credit volatility. These are banking and currency crisis, the presence of foreign and state-owned banks, financial development and the degree of concentration in the banking system. While we expect crisis to be positively related to credit volatility, the impact of foreign bank presence is uncertain. It was mentioned before that from the theoretical literature the final effect of foreign financial institutions on macroeconomic volatility depends on the type of shocks hitting the economy, which are in practice extremely difficult to isolate. An overall measure of foreign bank presence is then included in the conditional variance equation. If stabilizing effects predominate, then its coefficient should be negatively signed and statistically significant at conventional levels.

Regarding financial development and volatility, the literature points many channels by which the banking system may contribute to reduce macroeconomic volatility. Deeper financial markets may be more efficient in matching savers and investors, then allowing for a smoother absorption of shocks (Aghion et al. 1999). Moreover, the financial sector may also facilitate risk diversification, both at the micro and macroeconomic level, which would in turn reduce volatility. Another channel runs through the positive link between asymmetric information and output volatility. As far as the level of financial development is related to the ability of economies to generate and process information,

---

7 De Haas and Van Lelyveld (2004) and Kraft (2002) find similar results for eastern European countries, while Demirgüç-Kunt et al. (1998), Levine (1999) and Martinez Peria et al. (2005) find a similar conclusion studding cross-country experiences.

8 Apart from maintaining or increasing lending in crisis periods, another way to gain market share consists of acquiring local private banks. See Cull and Martinez Peria (2007).

a negative relationship between financial market development and volatility would be expected.[9]

As with foreign bank presence, banking concentration may either increase or reduce credit volatility. On one side, higher concentration levels may imply less competition and higher profits. The resulting increase in banking franchise values boosts their incentives to make good loans, thus avoiding investing in more speculative assets, and rendering the system more stable.[10] On the other side, more powerful institutions have the ability to charge higher interest rates to their customers, encouraging risk taking behaviour, and leading to greater vulnerabilities in the system (Boyd and De Nicoló 2005). Also, more concentrated banking systems are likely to induce moral hazard behaviour because concentrated banks are 'too important to fail' (Mishkin 1999). All in all, risk taking behaviour increases, which results in more vulnerable banks.[11]

Morgan and Strahan (2003) have already tested the impact of foreign bank entry, financial development and banking concentration on economic volatility, measured by the squared and absolute value deviation of actual from expected GDP and investment growth.[12] For the full sample, they find tentative evidence of a positive link between foreign bank presence and economic volatility. Concentration is not statistically significative, and financial development is found to increase economic volatility, which is puzzling both on theoretical and empirical grounds. Nevertheless, when they focus on a sample of Latin American countries, they obtain a negative coefficient in the GDP regression for both banking integration and financial development, although they are not statistically significant. Denizer et al. (2002) find evidence that financial deepness helps to reduce economic volatility, using a panel of 70 developed and developing countries starting in the mid-1950s. They find that financial development reduces per capita consumption, investment and income volatility growth. In the same direction, by studying bank-specific data on lending behaviour by domestically- and foreign-owned banks in Argentina and Mexico, Dages et al. (2000) find that foreign banks generally had higher loan growth rates than their domestically-owned counterparts, with lower volatility of lending, contributing to lower overall volatility of credit. Interaction terms between the crisis variables and foreign and state-owned banks are also included in the variance equation, to determine if these types of banks act as volatility absorbers or amplifiers.

---

9 See Bernanke and Gertler (1989, 1990), Greenwald and Stiglitz (1993) and Kiyotaki and Moore (1997) for theoretical models formalizing the abovementioned hypothesis. See Gertler and Gilchrist (1994) and Kashyap and Stein (1995, 2000) for empirical evidence using microeconomic data.

10 See Allen and Gale (2004) and Boot and Greenbaum (1993) for theoretical models on banking concentration and credit stability. Empirical work was conducted by Paroush (1995), Benston et al. (1995), and more recently Beck et al. (2005). On the relationship between franchise value (i.e., the present value of the stream of profits that a firm is expected to earn), and risk see Keeley (1990), Demsetz et al. (1996), Hellman et al. (2000), and Bergstresser (2001).

11 Empirical evidence around these lines is presented by Boyd and Graham (1991, 1996), De Nicoló and Kwast (2002), and de Nicoló et al. (2004).

12 Expected GDP and investment growth is computed from a regression on time and fixed effects, banking integration, as well as a set of control variables. The sample consists of a panel of nearly 100 countries and spans the period 1990-1997.

5 Panel unit root tests

Financial sector foreign direct investment in Latin America is a relatively new phenomenon, putting a burden on data availability and invalidating the use of time series techniques applied to individual countries. The usage of panel data that pools together information for different cross-sectional units increases the amount of information and the power of econometric estimations. Nevertheless, the usual concerns about nonsense spurious regressions and misleading statistical inferences still arise when using potential non-stationary panels, in which the time dimension exceeds by far the number of cross-sectional units. Indeed, checking the panel unit-root properties of the variables will be an ineluctable step in disentangling the effect that foreign banks may have on credit volatility in our sample of Latin American countries.

The number of tests to detect the presence of unit-roots in panels has been growing rapidly over the recent past. First generation techniques have ignored the possibility that unobserved common factors can affect the cross-section units simultaneously. This possibility can easily arise in macroeconomic applications that use country or regional data, substantially biasing the estimated coefficients and distorting the size of the test statistics. In particular, Pesaran (2007) shows that in the presence of high cross-sectional dependence, the bias is such that the empirical size is higher than the nominal size.[13] Recognizing this deficiency, a second group of panel tests have been proposed to successfully address this issue (see Breitung and Pesaran 2006, for a literature review). Nevertheless, the unwarranted application of these techniques does not go without difficulties. If panel unit root tests that allow for cross-section dependence are used when this is not high, tests might result in a loss of power.

Therefore, before deriving any inference on the statistical properties of the data it is necessary to establish whether or not the variables in the panel are subject to a significant degree of error cross-section dependence. This can be achieved by conducting first generation panel unit root tests, obtaining the residuals of each equation, and then computing the Lagrange Multiplier (LM) test suggested by Breusch and Pagan (1980). The LM statistics tests the null hypothesis of zero cross-equation error correlations, and is based on the average of the squared pair-wise sample correlation of the residuals ($\hat{\rho}_{i,j}$).[14] The test is given by:

$$CD_{lm} = T\sum_{i=1}^{N-1}\sum_{j=i+1}^{N}\hat{\rho}_{i,j}^2$$

Under the null hypothesis of no cross section dependence, the statistics converges to a chi-squared distribution with $N(N-1)/2$ degrees of freedom. This test has been shown to

---

13 The extent of over rejection increases with the degree of cross-sectional dependence, and with N and T.

14 Specifically, $\hat{\rho}_{i,j} = \hat{\rho}_{j,i} = \sum_{t=1}^{T} e_{i,t}e_{j,t} \Big/ \left(\sum_{t=1}^{T} e^2_{i,t}\right)^{1/2}\left(\sum_{t=1}^{T} e^2_{j,t}\right)^{1/2}$, where $e_{i,t}$ is the Ordinary Least Squares (OLS) estimate of the error equation used to test for unit roots for each i separately.

be especially suitable for cases in which N is sufficiently small relative to T, as in our case.[15]

Choi's (2001) test is used to determine the existence of panel unit roots, assuming that the individual time series are cross-sectionally independent. While the test can be constructed by applying any unit root test to the individual series, here we estimate standard ADF regressions for each cross-section unit and combine the *p-values* associated with each lagged dependent variable so as to form the following statistics:

$$Z = \frac{1}{\sqrt{N}} \sum_{i=1}^{N} \Phi^{-1}\left(p_i\right)$$

where Φ is the standard normal cumulative distribution function, and p_i is the asymptotic *p-value* for the ADF unit root test of cross section i. The null hypothesis is that all times series have a unit-root, while under the alternative hypothesis some of the variables are stationary. Choi (2001) shows that under the null hypothesis, Z converges to a standard normal distribution. A main advantage is its improved finite sample power over other traditional techniques, like Levin et al. (2002) and Im et al. (2003).

The panel unit root test that allows for cross-sectional dependence is constructed around the same lines as before, the only difference being the ADF regression used to obtain the *p-values* and the fact that the Z statistics does not converge any more to a standard normal distribution, even for large T and N. In such a case, stochastically simulated critical values have to be used.[16] Pesaran (2007) proposed an easy way to deal with the problem of cross-sectional correlation arising from an unobserved common factor, which consists of simply augmenting the standard ADF regression with cross-section averages of lagged levels and first-differences of the individual series. In particular, the Z test will be constructed now using the OLS p-values associated to coefficient β_i in the following p^{th} order regression:

$$\Delta y_{it} = \alpha_{mi} d_{mt} + \beta_i y_{i,t-1} + \varsigma_i \bar{y}_{t-1} + \sum_{j=0}^{p_i} \delta_{ij} \Delta \bar{y}_{t-j} + \sum_{j=1}^{p_i} \rho_{ij} \Delta y_{i,t-j} + e_{it}$$

where $\bar{y}_t$ is the cross-section mean of y_{it} given by $\bar{y}_t = N^{-1} \sum_{i=1}^{N} y_{it}$, and m = 1,2,3, with $d_{1t} = \{\phi\}$, $d_{2t} = \{1\}$ and $d_{3t} = \{1, t\}$.[17] Simulations performed by Pesaran show that the cross-sectionally augmented version of Choi's test has quite satisfactory power and size, even for small values of T and N.

---

15 Pesaran (2004) has proposed another test for cross section dependence, using the level rather than the squared values of the pair-wise correlation coefficients. By using Pesaran's test, we cannot rule out the possibility that positive and negative correlations compensate each other, leading to the erroneous inference that no cross-sectional dependence is present in the data.

16 These depend on the sample size (T and N), and on the deterministic components included in the regressions.

17 The augmentation order p is selected here on account of the Schwartz Information Criterion applied to each cross section ADF equation, without the cross-section variables $\bar{y}_{t-1}$ and $\Delta \bar{y}_{t-j}$, $j = 1,\ldots,p_i$.

Table 2 shows the cross-sectional dependence and panel unit root test results for the abovementioned set of variables over the period 1995:1-2001:4. Recognizing that potential biases may arise when including different deterministic components in the estimating unit root equations, we have adopted the following strategy. Country-variables that do not exhibit a trend were centred, and no deterministics at all were included in the analysis, which allows for testing the null of a random walk without drift process against the alternative of the variable being level stationarity. When country-variables do exhibit a trend, models with a constant, and with and without a linear time trend were estimated. Only if the null hypothesis is not rejected in both cases, then we can be confident that a unit root is present in the data. Indeed, column 1 contains variable's names, column 2 indicates the cases in which some of the equations include a constant only, while column 3 presents those in which the same set of equations was extended to include a linear trend.

The test of cross-section dependence performed on the residuals of the ADF estimations that do not contain averages of the dependent variable are reported in column 4. May be not surprisingly, significant dependence across countries was found in macroeconomic variables like real GDP, the fiscal balance, real private lending and real exchange rates. In contrast, most of the banking variables are free of common factors that may make them to be highly correlated. The last column of Table 2 presents the results of the *Z* statistics, with or without cross section dependence according to the results obtained by the CD_{lm} statistics presented before.[18] For almost all the variables it is possible to reject the unit root null, while for real GDP the test gives conflicting results. Somewhat expectedly for trendy variables, the unit root null is not rejected when only a constant is included in the equation while it is rejected when a trend is added.[19] Individual unit root tests results for the U.S. real GDP and the Federal Funds Rate (not reported) do not allow rejecting the presence of a unit root in the series.

Based on these results, variables are transformed accordingly. In particular, when the unit root is rejected and variables are expressed as shares (like foreign bank presence, concentration, etc.), the level difference with respect to the HP trend is considered. And when they are not, the percentage deviation with respect to the trend is used instead. Of course, when a unit root is present, the variables are differentiated.[20] Finally, in the case of the real GDP in which results are not conclusive, estimations are carried out under both types of transformation, to ensure the robustness of the econometric results.

---

18 Accordingly, when the CD_{lm} statistics does not reject the null, the unit root test that excludes cross-section averages was used.

19 This means the test has difficulties in discerning whether the series is trend stationary or random walk with drift.

20 Panel unit root tests (not reported) performed on the transformed variables reject the unit root null in all cases.

6 Estimation results

6.1 The econometric methodology

This section presents the steps followed to estimate the model, selecting among a set of variables those significant at conventional statistical levels, while a more detailed analysis of the role played by interaction terms in the mean equation is presented in Annex I, following a particular to general strategy.[21] To choose the final specification for the model, we preliminary identify the presence of fixed effects in the mean and variance equation, we test for poolability of the data (i.e., that coefficients in the mean equation are the same across countries), and we identify the presence of ARCH effects in the conditional covariance equation.

Indeed, we begin by estimating the mean equation by OLS and testing for the presence of fixed effects using a Chow test, assuming that the data are poolable. In particular, we test for the null hypothesis that all the individual fixed dummy variables included in equation (1) are zero: $H_0 : \mu_1 = \mu_2 = \cdots = \mu_{N-1} = 0$. The Chow test is just an *F*-test for the joint significance of these dummies and is computed as:

$$F_0 = \frac{(RRSS - URSS)/(N-1)}{URSS/(NT-N-K)} \sim F_{N-1,N(T-1)-K}$$

where *N* and *T* are the cross and time series dimension respectively, *K* is the number of coefficients, excluding the dummy variables and the constant, *RRSS* is the residual sum of squares of the restricted model (i.e., the pooled OLS), and *URSS* is the residual sum of squares of the unrestricted model (i.e., the OLS model that includes the fixed effects dummy variables).

Once we have decided on the inclusion of fixed effects dummy variables, we proceeded to test for poolability of the data in the estimated equation. Again, this is performed using a Chow statistics to test the null hypothesis that all the coefficients are equal among the cross-sectional units. The statistic is defined as:

$$F_0 = \frac{(RRSS - URSS)/(N-1)K'}{URSS/N(T-K')} \sim F_{(N-1)K',N(T-K')}$$

where $K' = K + 1$. The *RRSS* is given by the OLS estimation performed on the pooled model (i.e., assuming homogenous coefficients), whereas the unrestricted residual sum of squares (*URSS*) is the sum of the residual squares performed on each separate country-specific OLS regression (i.e., assuming a different coefficient for each country equation).

The next step consisted in using the residuals of the previously estimated mean equation to test for the presence of fixed effects in the conditional variance. In particular, we ran a regression of the form:

21 Results presented in Annex I are also used as a guide to choose the most general model estimated in this section.

$$u_{it}^2 = \beta_0 + \mu_i + \sum_{j=1}^{p} \alpha_j u_{it-j}^2$$

and tested for the joint significance of μ_i using the same Chow test as before, for a given value p. Once we have decided on the inclusion of fixed effects in the conditional variance, we proceeded to test for the presence of ARCH effects (i.e., the significance of α_j's). The null hypothesis is then that no ARCH effects are present in the model.

Once the model's specification has been determined, equations (1)-(4) were estimated jointly using maximum likelihood techniques, including different sets of explanatory variables in the variance equation. To avoid potential endogeneity problems, lagged domestic variables are introduced in both the mean and variance equations. Indeed, we measure the impact of previous values of the variables on subsequent credit behaviour. The exceptions are the currency and banking crisis dummies and the foreign variables (U.S. GDP and the Federal Funds Rate), which are included in contemporaneous form.

6.2 The determinants of credit behaviour: the mean equation

Model (1) in Table 3 presents the results for the most general estimation performed on the mean equation, while model (2) excludes the insignificant variables at conventional statistical levels.[22] In both estimations, the null hypothesis of absence of fixed effects in the mean equation is not rejected, suggesting that fixed effects are not present in the data. The same results hold true for the null of poolability, suggesting that coefficients are significantly equal across countries, justifying the standard homogeneity assumption. Models (3)-(10) present the same estimation as model (2), the only difference being the inclusion of explanatory variables in the variance equation (Table 4).

As reported in that Table, tests of fixed effects in the conditional variance do allow rejecting the null hypothesis for both models (1) and (2). This is due to the dummy for Mexico which is highly significant, while the rest are not statistically different from zero. Absence of ARCH effects of order one is rejected at the 10 per cent level, while absence of ARCH(4) effects is rejected even at the 1 per cent level in model (2). To keep the specifications as parsimonious as possible, all the models are estimated including only an ARCH(1) effect in the conditional variance equation, while testing whether higher order ARCH effects are present in the residuals.[23]

Results for the mean equation show that the autoregressive coefficient is highly significant suggesting persistence in the dependent variable.[24] Persistence in credit behaviour has also been found by Detragiache et al. (2006) for a panel of 89 low income and lower middle income countries over the second half of the 1990s. Our estimations

---

22 Giving that macro variables are very likely to be correlated, then reducing individual significance, a larger 10 per cent level was used as a threshold instead of the traditional 5 per cent.

23 Including a dummy for Mexico in this equation makes the algorithm not to converge. The dummy was then not included in the estimations.

24 Model (2) was re-estimated adding an autoregressive component up to the order four, giving that quarterly data is being used in the econometric analysis. Results (not reported) show that only the first autoregressive coefficient is statistically significant.

also point to a high degree of lending procyclicality, a finding that has also been found by Barajas and Steiner (2002) and Arena et al. (2007) for Latin American countries.[25] Results presented in Annex I also point that foreign banks do not contribute to amplify credit cycles in the region, given that the interaction between domestic GDP and foreign bank presence is not significant. This finding is consistent with Micco and Panizza (2006).

As previously mentioned, push factors like the U.S. GDP and the Federal Funds Rate are included to account for the buoyancy of international financial markets. While having the expected signs, they are found insignificant at standard statistical levels in the econometric estimations. This indicates that mainly pull factors played a role in shaping credit behaviour in the second half of the 1990s in our sample of Latin American countries.[26] Interaction terms between these variables and foreign bank presence are not significant, giving no credit to the hypothesis that foreign banks may exacerbate external shocks (Annex I).

Two pull variables which are not significant in the econometric estimation are the fiscal balance and the spread between the lending and borrowing rates. It was expected that healthier public finances would lead to a crowd-in effect, then boosting domestic loans. And higher intermediation costs would make loans more expensive, then reducing the stock of credit in the economy. As with the push variables, even when the coefficients have the expected signs, they are not significant at conventional statistical levels.

As hypothesized, the lagged degree of financial development is highly significant, pointing that more developed financial systems today (proxying for a lower degree of financial imperfections) help to foster real private credit in subsequent periods. As expected, the real exchange rate is also significant, showing that depreciations reduce credit in domestic economies, because it makes more expensive the repayment of loans in foreign currency. This finding is similar to that of Arena et al. (2007) for the same region.

Banking variables including foreign and public bank presence and the degree of concentration in the sector do not seem to have a role to play in affecting the level of loans granted by the system. As aforementioned, empirical evidence is inconclusive up to the impact of foreign banks on credit dynamism. While foreign banks seem to have been more active in the granting of loans than their domestic counterparts in Latin America over the second half of the 1990s (Crystal et al. 2002), it seems that they restrict access to private commercial credit once attention is focused on low income and lower middle income countries (Detragiache et al. 2006). These authors also observe that the negative effect disappears once mid- and high-income countries are included in the sample, a finding which is consistent with the fact that a larger foreign bank presence is associated with more, rather than less, private credit in higher income economies. In contrast to our findings, the same authors find that state banks lead to a slowdown in credit growth, a finding that persists even after including wealthier countries in the estimation sample.

---

25 Similar results are found when the domestic GDP growth is used instead of the HP filtered series.

26 Since Spanish owned banks account for almost half of total foreign bank lending in the region, the American GDP was replaced by the Spanish GDP, but results remained unchanged.

Both banking and currency crisis are expected to have a negative effect on credit behaviour. During banking crisis generally involving bank runs, deposits are depleted putting a burden on the granting of loans. Currency crisis may trigger balance sheet effects, leading the banking sector to experience bankruptcy problems. As mentioned before, they also tend to be associated with general losses of confidence in the system and a retrenching in deposits. Only the currency crisis dummy was found to have the expected negative sign while banking crisis does not have a statistically distinguishable effect on credit behaviour.[27] This might be due to the fact that banking crises tend to coincide with deterioration in economic fundamentals making their impact indistinguishable from other cyclical downturns. Regarding interaction terms, foreign banks do not seem to behave differently than national institutions, both in banking and currency crisis (Annex I). Nevertheless, government owned banks do seem to have a stabilizing role on credit during banking crisis. A similar result reported by Micco and Panizza (2006) states that public-owned banks may play a useful credit-smoothing role, because their lending is less responsive to macroeconomic shocks than the lending of private banks.

6.3 The impact of foreign bank presence on credit volatility

The estimation of the variance equation is presented in Table 4, where different sets of explanatory variables susceptible of having a burden on credit volatility are included in the analysis. This is a main difference with previous work on the issue of disentangling the impact of foreign banks and financial development on macroeconomic volatility, since we modelled together the first and second conditional moment of the data, instead of using two-step estimators which are known to be inefficient.

Indeed, in Models (1) and (2) of Table 4, the variance equation only includes an ARCH(1) process, significant at the 10 per cent level. Model (3) also includes a dummy variable for banking crisis, which proves to be highly significant. Indeed, while periods of banking crisis do not seem to have a statistically significant effect on the level of credit, they are characterized by a heightened volatility of credit to the private sector. Next, we test for a differential behaviour of foreign and state-owned banks during financial stress periods, by including, in turn, interaction terms between the dummies for banking crisis and foreign and public bank presence. If either one or the other exacerbates credit volatility, the coefficient should be positive and statistically significant at conventional levels. In the present case, there does not seem to be such differential behaviour. In the case of state-owned banks, the conclusion seems to support the idea that they have a stabilizing effect during stress periods but only on the level of credit (i.e., the first moment of the data). While banking crisis increase credit volatility, this does not seem to be true for currency crisis, as shown by Model (6).

The revision of the theoretical literature presented at the beginning of the paper has shown that there is no definite answer to the question of whether or not foreign banks raise or reduce credit volatility. Everything depends on the types of shocks hitting the economy, which are in practice extremely difficult to isolate. That is why in empirical work only an aggregate measure of foreign bank presence is included in the estimations.

---

[27] Note that the banking crisis dummy variable was not included in the general estimation performed in Table 3, since it was found to be insignificant in spite of having an expected negative sign in the exercise reported in Annex I.

If stabilizing effects predominate, then the coefficient should be negatively signed and statistically significant at conventional levels. The opposite is of course true when foreign banks increase credit volatility. As gauged by the empirical evidence presented under Model (7) in Table 4, foreign banks do seem to have contributed to reduce real credit volatility in our sample of Latin American countries over the period 1995-2001. The coefficient for foreign banks is negative and statistically significant at the 5 per cent level, which is consistent with Dages et al. (2000) for Argentina and Mexico, who find that foreign banks exhibit lower volatility of lending than their domestically-owned counterparts, contributing to lower overall credit volatility.

Of course, it can be argued that foreign bank presence could just be capturing the degree of development of the financial sector, or a more concentrated banking system. Financial foreign direct investment may help develop the banking sector because foreign banks tend to have higher average loan growth rates, higher average provisioning expenses and greater loss-absorption capacity (Dages et al. 2000; Crystal et al. 2002). As mentioned before, theoretical literature abounds on how a deeper financial sector is expected to lead to a more stable macroeconomy. Also, foreign bank entry often consists on the buying and merging of local banks, which may result in more concentrated banking systems.

Evidence reported under Model (8) shows that deeper banking systems indeed result in lower credit volatility, a finding coherent with Denizer et al. (2002) for other macroeconomic variables, like GDP, consumption and investment.[28] Including concentration among the explanatory variables in the conditional variance does not modify the previous findings, both in terms of signs and statistical significance, but shows that concentration has no role to play in shaping credit volatility patterns. This result is in line with Morgan and Strahan (2003), who found that banking concentration does not affect GDP and investment growth volatility in a sample of Latin American countries over the period 1990-1997.

In a last specification, state-owned banks were included in the econometric estimation, but proved not to be significant. Indeed, the final specification retained for the conditional variance equation is Model (8), in which banking crisis increase real credit volatility, while foreign bank entry and banking development reduce such volatility.[29] The diagnostic tests performed on this and the other models show that we can be confident about the specification of the econometric equations. We cannot reject the null hypothesis that the standardized residuals are normal, as stated by equation (4). Also, we test for the correct specification of the mean and variance equations by testing for the presence of autocorrelation and conditional heteroscedasticity in the standardized residuals. Indeed, we never reject the hypothesis of absence of low and high orders of ARCH effects, and while autocorrelation of order four seems to be present in models (1)

---

28 Note that, due to the positive correlation between foreign bank presence and banking sector development, the coefficient for foreign bank penetration is slightly reduced in absolute terms, while remaining significant at conventional statistical levels.

29 To rule out the possibility of reverse causality between foreign bank presence and credit volatility, the conditional variance estimated using model (2) was included as a regressor in an equation having foreign bank presence as the dependent variable. Credit volatility proved to be insignificant at conventional statistical levels.

to (7), it disappears once financial development is included among the regressors of the conditional variance equation.

7 Conclusion

Foreign bank entry into developing countries has exploded since the 1990s, favoured by the liberalization of external sectors and the embracement of a series of market-friendly policy reforms, including deregulation and privatization of the banking sector. One region that has been very active in attracting foreign direct investment into the banking industry has been Latin America. This unprecedented internationalization of the banking sector has prompted a debate on the potential consequences for the recipient countries, in terms of efficiency and diversification of the banking industry, the quality of the regulatory environment, and competition and access to banking services by small and medium size enterprises, among a broader set of topics studied by the literature.

Against this background, the present paper has investigated the impact of foreign bank presence on real credit volatility in a panel of eight Latin American countries, using quarterly data over the period 1995:1-2001:4. We have tried to disentangle the effect of foreign banks on credit volatility by using ARCH techniques to model jointly the first and second conditional moment of real domestic credit. Indeed, the conditional volatility equation is extended to include the degree of development of the banking sector and the internationalization of the banking system, among other regressors. To the knowledge of the authors, this is the first time that such tools are used to analyse the impact of foreign bank presence on macroeconomic volatility.

The theoretical literature examining the link between foreign direct investment in the banking sector and macroeconomic stability is rather limited and does not give a clear answer to this issue. Everything depends on the type of shock hitting the economy, and indeed, the overall impact of banking integration on volatility is then an empirical question. But apart from the theoretical arguments, there are other potential explanations of why foreign banks may contribute to credit stability. Foreign banks are typically well diversified institutions with access to a broader set of liquidity sources than domestic banks. They may also allow for a faster recapitalization of local banks after a crisis, and they may have superior risk management systems and better screening devices, then improving the quality of their assets. Finally, they have the potential of avoiding capital flight in the case of domestic shocks, since people may prefer to redirect deposits towards foreign owned institutions, instead of withdrawing the money out of the system altogether. This contributes to higher funding and lending stability.

The main findings of the paper regarding credit volatility are as follows. First and perhaps not surprisingly, banking crisis steep up real credit volatility. Public and foreign banks do not have a discernible effect (positive or negative) during these stressful periods. Second, evidence reported in this paper shows that deeper banking systems result in lower credit volatility, a finding that is coherent with Denizer et al. (2002) for volatility in other macroeconomic variables. Finally, stabilizing effects predominate in such a way that foreign bank presence reduces credit volatility in our panel of eight Latin American countries over the period 1995-2001.

ANNEX I
ASSESSING THE IMPACT OF INTERACTION TERMS ON CREDIT DYNAMICS

This Annex presents a particular to general econometric analysis to evaluate the importance of interaction terms in shaping the dynamics of the level of real credit (mean equation). As such, interaction terms between public/foreign banks, and banking/currency crises, and between foreign banks and the Federal Funds Rate and the American and domestic GDP are included in the analysis.

Estimation results of applying this procedure to the panel of eight Latin American countries are presented in Tables A.1 and A.2. As in the main text, for each model we test for the presence of fixed effects in the mean and variance equation, we test for poolability of the data (i.e., that coefficients in the mean equation are the same across countries), and we identify the presence of ARCH effects in the conditional covariance equation. Empirical evidence shows that we cannot reject the null hypothesis of absence of fixed effects and poolability of the data in the mean equations for all the estimated models. And while we can reject the null of absence of fixed effects in the conditional variance equation, this is due to the high significance of the dummy for Mexico.[30] Moreover, as we cannot reject the presence of ARCH effects, the conditional variance will be modelled allowing for an autoregressive heteroscedastic process of order one. Estimations are then performed using quasi maximum-likelihood techniques. Variables are introduced one by one in the mean equation and kept in the next round of estimations only when they are significant in the previous step. This allows keeping the model as parsimonious as possible, which is of paramount importance to guarantee convergence of the estimating algorithm.

The most parsimonious model that we estimate includes only an autoregressive term (Model 1). As it proves to be highly significant, it is kept in the subsequent estimations. Note that, while this term always keeps a high degree of significance throughout all the models, the absolute value of the coefficient is reduced as we start including extra explanatory variables. This is due to the fact that other regressors start capturing part of the credit dynamics that was before completely imputed to the autoregressive term.

The next model includes domestic GDP to capture procyclicality in credit behaviour. This variable also proves to be highly significant and positively signed across all the estimated models. In Model (3) we test for the impact that foreign bank presence may have in amplifying or reducing procyclicality in lending behaviour. If the hypothesis that international banking institutions exacerbate procyclicality is not rejected, then an interaction term between domestic GDP and foreign bank presence should be positively signed and statistically significant at conventional levels. Empirical evidence presented in Table A.1 shows that foreign banks cannot be blamed of having amplified credit cycles in the region, a finding that is in line with that of Micco and Panizza (2006) and Arena et al. (2007).

30 As mentioned in the text, when a dummy for this country was introduced in the variance equation the estimated algorithm did not reach convergence. Indeed, it was decided to exclude such dummy variable.

In models (4)-(7) we include external or push factors that may have a burden on domestic lending patterns. In particular, we include the U.S. GDP and the Federal Funds Rate, as well as their interactions with foreign banks, to test the hypothesis that these institutions amplify foreign shocks. It is found that, while the U.S. GDP has a positive impact on credit, neither the Federal Funds Rate nor the interaction terms are statistically significant. The finding of a positive coefficient for the American GDP is consistent with Goldberg (2002) who observes that U.S. bank claims to Latin American countries increase as the U.S. economy grows. It should be noted however, that while the U.S. GDP is kept in the estimations because it is significant in this round, it will lose its explanatory power in subsequent models.

Next, we include the fiscal balance, the spread between the lending and deposit rates, the degree of financial development, and the bilateral real exchange rate with the U.S. (models 8-11). From this set of variables, only the last two appear to be significant and were consequently kept in subsequent estimations. As noted in the main text, lagged degrees of financial development help boosting real private credit in the future. Also, increases in the real exchange rate (i.e. depreciations) have a negative impact on the level of credit, since they make more expensive the repayment of loans in foreign currency. Similar results for the region were found by Arena et al. (2007).

In models (12)-(14) of Table A.2 we added banking variables to the econometric regressions: public and foreign bank presence, and the concentration of the banking system. None of these proved to be significant. Empirical evidence to date is inconclusive up to the impact of foreign banks on credit dynamism. Our results coincide with those of Detragiache et al. (2006) who found that foreign banks have no discernible effect on the level of credit granted by the system, once their sample of countries is expanded to include mid- and high-income countries.

The next six models (15-20) include the occurrence of banking and currency crisis and their interactions with foreign and state-owned banks. While having the expected negative sign, the banking crisis dummy is not statistically significant. As explained in the text, this may be due to the fact that the impact of crises might be captured by movements in other macroeconomic variables. Concerning the interaction with foreign and state-owned banks, only the last one is positive and statistically significant. This means that state-owned banks help stabilizing credit levels during banking crisis periods, including a currency crisis dummy among the regressors gives a negative and statistically significant coefficient, while foreign and state-owned banks do not seem to exacerbate or buffer the impact of such events on credit behaviour.

References

Aghion, P., A. Banerjee, and T. Piketty (1999). 'Dualism and Macroeconomic Volatility'. *Journal of Economics*, 114: 1359–97.

Allen, F., and D. Gale (2004). 'Competition and Financial Stability'. *Journal of Money, Credit and Banking*, 36 (3), 453–80.

Arena, M., C. Reinhart, and F. Vázquez (2007). 'The Lending Channel in Emerging Economies: Are Foreign Banks Different?'. *Working Paper* 48. Washington, DC: IMF.

Barajas, A., and R. Steiner (2002). 'Credit Stagnation in Latin America'. *Working Paper* 53. Washington, DC: IMF.

Barajas, A., R. Steiner, and N. Salazar (2000). 'Foreign Investment in Colombia's Financial Sector'. In S. Claessens and M. Jansen (eds), *The Internationalization of Financial Services: Issues and Lessons for Developing Countries*. Boston, MA.

Beck, T., A. Demirgüç-Kunt, and M. S. Martinez Peria (2005). 'Reaching Out: Access to and Use of Banking Services Across Countries'. Unpublished Manuscript. Washington, DC: World Bank.

Benston, G. J., W. C. Hunter, and L. D. Wall (1995). 'Motivations for Bank Mergers and Acquisitions: Enhancing the Deposit Insurance Put Option Versus Earnings Diversification'. *Journal of Money, Credit, and Banking*, 27 (3): 777–88.

Bernanke, B., and M. Gertler (1989). 'Agency Costs, Net Worth, and Business Fluctuations'. *American Economic Review*, 79 (1): 14–31, March.

Bernanke, B., and M. Gertler (1990). 'Financial Fragility and Economic Performance'. *The Quarterly Journal of Economics*, 105 (1): 87–114, February.

Berger, A., and T. Hannan (1989). 'The Price-Concentration Relationship in Banking'. *Review of Economics and Statistics*, 71: 291–99.

Berger, A., L. Klapper, and G. Undell (2001). 'The Ability of Banks to Lend to Informationally Opaque Small Businesses'. *Journal of Banking and Finance*, 25: 2127–67.

Bergstresser, D. (2001). 'Market Concentration and Loan Portfolios in Commercial Banking'. Unpublished manuscript. Cambridge, MA: MIT Department of Economics.

Boot, A., and S. Greenbaum (1993). 'Bank Regulation, Reputation and Rents: Theory and Policy Implications. In C. Mayer and X. Vives (eds), *Capital Markets and Financial Intermediation*. Cambridge: Cambridge University Press, pp. 262–85.

Boyd, J. H., and G. De Nicoló (2005). 'The Theory of Bank Risk-Taking and Competition Revisited'. *Journal of Finance*, 60 (3): 1329–43.

Boyd, J. H., and S. L. Graham (1991). 'Investigating the Banking Consolidation Trend'. *Federal Reserve Bank of Minneapolis Quarterly Review*, Spring Issue, 1–15.

Boyd, J. H., and S. L. Graham (1996). 'Consolidation in U.S. Banking: Implications for Efficiency and Risk'. *Working Paper* 572, Federal Reserve Bank of Minneapolis, Minneapolis, MN.

Breitung, J., and M. H. Pesaran (2006). ‘Unit Roots and Cointegration in Panels’. In L. Matyas and P. Sevestre (eds), *The Econometrics of Panel Data: Fundamentals and Recent Developments in Theory and Practice*. Kluwer Academic Publishers.

Breusch, T. S., and A. R. Pagan (1980). ‘The Lagrange Multiplier Test and Its Applications to Model Specification in Econometrics’. *Review of Economic Studies*, 47 (1): 239–53.

Buch, C. (2000). ‘Why Do Banks Go Abroad? Evidence from German Data’. *Financial Markets, Institutions and Instruments*, 9 (1): 33–67.

Caprio, G., D. Klingebiel, L. Laeven, and G. Noguera (2005). ‘Banking Crisis Database’. In P. Honohan and L. Laeven (eds), *Systemic Financial Crises*, Cambridge University Press, Cambridge.

Carstens, A., D. Hardy, and C. Pazarbasioglu (2004). ‘Avoiding Banking Crises in Latin America’, *Finance & Development*, September Issue, 30–33.

Choi, I. (2001). ‘Unit Root Tests for Panel Data’. *Journal of International Money and Banking*, 20: 249–72.

Claessens, S., and M. Jansen (2000). ‘The Internationalization of Financial Services: Issues and Lessons for Developing Countries’. In S. Claessens and M. Jansen (eds), *The Internationalization of Financial Services: Issues and Lessons for Developing Countries*, Kluwer Academic Press, Boston, MA.

Claessens, S., A. Demirgüç-Kunt, and H. Huizinga (2001). ‘How does Foreign Entry Affect the Domestic Banking Market?’. *Journal of Banking and Finance*, 25 (5): 891–911.

Corvoisier, S., and R. Gropp (2002). ‘Bank Concentration and Retail Interest Rates’. *Journal of Banking and Finance*, 26 (11): 2155–89.

Crystal, J., B. G. Dages, and L. Goldberg (2001). ‘Does Foreign Ownership Contribute to Sounder Banks? The Latin American Experience’. In R. Litan, P. Masson and M. Pomerleano (eds), *Open Doors: Foreign Participation in Financial Systems in Developing Countries*, Brookings Institution Press, Washington, DC.

Crystal, J., B. G. Dages, and L. Goldberg (2002). ‘Has Foreign Bank Entry Led to Sounder Banks in Latin America?’. *Current Issues in Economics and Finance*, 8 (1), Federal Reserve Bank of New York, NY.

Cull, R., and M. Martínez Pería (2007). ‘Foreign Bank Participation and Crises in Developing Countries’. *Working Paper*, No. 4128, World Bank, Washington, DC.

Dages, B. G., L. S. Goldberg, and D. Kinney (2000). ‘Foreign and Domestic Bank Participation in Emerging Markets: Lessons from Mexico and Argentina’. *Economic Policy Review*, September Issue, 17–36, Federal Reserve Bank of New York, NY.

Dahl, D., and R. Shrieves (1999). ‘The Extension of International Credit by U.S. Banks: A Disaggregated Analysis, 1988-1994’. *Journal of International Money and Finance*, 18: 153–167.

De Haas, R., and I. Van Lelyveld (2004). ‘Foreign Bank Penetration and Private Sector Credit in Central and Eastern Europe’. *Journal of Emerging Market Finance*, 3 (2): 125–51.

De Haas, R., and I. Van Lelyveld (2006). 'Foreign Banks and Credit Stability in Central and Eastern Europe: A Panel Data Analysis'. *Journal of Banking and Finance,* 30 (7): 1927–52.

De Nicoló, G., and M. L. Kwast (2002). 'Systemic Risk and Financial Consolidation: Are They Related?'. *Journal of Banking and Finance*, 26: 861–80.

De Nicoló, G., P. Bartholomew, J. Zaman, and M. Zephirin (2004). 'Bank Consolidation, Internationalization, and Conglomeration: Trends and Implications for Financial Risk'. *Financial Markets, Institutions and Instruments*, 13 (4): 173–217.

Demirgüç-Kunt, A., L. Laeven, and R. Levine (2003). 'The Impact of Bank Regulations, Concentration, and Institutions on Bank Margins'. *Policy Research Working Paper*, 3030, World Bank, Washington, DC.

Demirgüç-Kunt, A., R. Levine, and H. Min (1998). 'Foreign Banks: Efficiency, Stability and Growth'. *The Implications of Globalization of World Financial Markets*, Bank of Korea, Seoul.

Demsetz, R., M. Saidenberg, and P. Strahan (1996). 'Banks with Something to Lose: The Disciplinary Role of Franchise Value'. *Economic Policy Review*, October Issue, Federal Reserve Bank of New York, NY.

Denizer, C., M. Iyigun, and A. Owen (2002). 'Finance and Macroeconomic Volatility'. *Contributions to Macroeconomics*, 2 (1), Berkeley Electronic Press, Berkeley, CA.

Detragiache, E., and P. Gupta (2006). 'Foreign Banks in Emerging Market Crises: Evidence from Malaysia'. *Journal of Financial Stability*, 2 (3): 217–42.

Detragiache, E., T. Tressel, and P. Gupta (2006). 'Foreign Banks in Poor Countries: Theory and Evidence'. *Working Paper*, 18, International Monetary Fund, Washington, DC.

Dornbusch, R. (1983). 'Real Interest Rates, Home Goods, and Optimal External Borrowing'. *Journal of Political Economy*, 91 (1): 141–53.

Evanoff, D., and E. Ors (2002). 'Local Market Consolidation and Bank Productive Efficiency'. *Working Paper*, 25, Federal Reserve Bank of Chicago, Chicago, IL.

Frankel, J. A., and A. K. Rose (1996). 'Currency Crashes in Emerging Markets: Empirical Indicators'. *Discussion Paper*, 1349, Centre for Economic Policy Research, London, UK.

Galindo, A., A. Micco, and A. Powell (2005). 'Loyal Lenders or Fickle Financiers: Foreign Banks in Latin America'. *Working Paper*, 529, Inter-American Development Bank, Washington, DC.

Gertler, M., and S. Gilchrist (1994). 'Monetary Policy, Business Cycles, and the Behavior of Small Manufacturing Firms'. *Quarterly Journal of Economics*, 109, 309–340.

Goldberg, L. S. (2002). 'When is U.S. Bank Lending to Emerging Market Volatile?'. *Working Paper*, 8209. Cambridge, MA: National Bureau of Economic Research.

Goldberg, L. S. (2007). 'Financial Sector FDI and Host Countries: New and Old Lessons'. *Economic Policy Review*, 13 (1): 1–17.

Greenwald, B., and J. Stiglitz (1993). ‘Monetary Policy and the Theory of the Risk-averse Bank’. Working Papers in Applied Economic Theory 93-04, Federal Reserve Bank of San Francisco.

Hannan, T. (1991). ‘Bank Commercial Loan Markets and the Role of Market Structure: Evidence from Surveys of Commercial Lending’. *Journal of Banking and Finance*, 15: 133–49.

Hellman, T., K. Murdock, and J. Stiglitz (2000). ‘Liberalization, Moral Hazard in Banking, and Prudential Regulation: Are Capital Requirements Enough?’. *American Economic Review*, 90 (1): 147–65.

Holmström, B., and J. Tirole (1997). ‘Financial Intermediation, Loanable Funds, and the Real Sector’. *Quarterly Journal of Economics*, 112 (3): 663–91.

Horvath, E., K. Mérõ, and B. Zsámboki (2002). ‘Studies on the Procyclical Behaviour of Banks’. *Occasional Papers*, 10, National Bank of Hungary, Budapest.

Im, K., M. H. Pesaran, and Y. Shin (2003). ‘Testing for Unit Roots in Heterogeneous Panels’. *Journal of Econometrics*, 115: 53–74.

Jenneau, S., and M. Micu (2002). ‘Determinants of International Bank Lending to Emerging Market Countries’. *Working Paper*, 112, Bank of International Settlements, Basel.

Kashyap, A. K., and J. C. Stein (1995). ‘The Impact of Monetary Policy on Bank Balance Sheets’. Carnegie-Rochester Conference Series on Public Policy, 42: 151–95.

Kashyap, A. K., and J. C. Stein (2000). ‘What Do a Million Observations on Banks Say about the Transmission of Monetary Policy?’. *American Economic Review*, 90 (3): 407–28.

Keeley, M. (1990). ‘Deposit Insurance, Risk and Market Power in Banking’. *American Economic Review*, 80 (5): 1183–200.

Kiyotaki, N., and J. Moore (1997). ‘Credit Cycles’. *Journal of Political Economy*, 105 (2): 211–48.

Kraft, E. (2002). ‘Characteristics and Behaviour of Foreign Banks During and After the Banking Crisis’. Presented at the 5th Conference on Financial Sector Reform in Central and Eastern Europe: The Impact of Foreign Bank Entry, Tallinn.

Levin, A., J. C. Lin, and C.-S. Chu (2002). ‘Unit Root Tests in Panel Data: Asymptotic and Finite-Sample Properties’. *Journal of Econometrics*, 108: 1–24.

Levine, R. (1997). ‘Financial Development and Economic Growth: Views and Agenda’. *Journal of Economic Literature*, 35: 688–726.

Levine, R. (1999). ‘Foreign Bank Entry and Capital Control Liberalization: Effects on Growth and Stability’. Working Paper, Carlson School of Management, University of Minnesota, Minneapolis, MN.

Martinez Peria, M. S., A. Powell, and I. Vladkova (2005). ‘Banking on Foreigners: The Behaviour of International Bank Lending to Latin America, 1985-2000’. *IMF Staff Papers*, 52 (3). Washington, DC: IMF.

Micco, A., and U. Panizza (2006). ‘Bank Ownership and Lending Behaviour’. *Economics Letters*, 93 (2): 248–54.

Micco, A., U. Panizza, and M. Yañez (2007). ‘Bank Ownership and Performance: Does Politics Matter?’. *Journal of Banking and Finance*, 31 (1): 219–41.

Mishkin, F. (1999). ‘Financial Consolidation: Dangers and Opportunities’. *Journal of Banking and Finance*, 23: 675–91.

Morgan, D., and P. Strahan (2003). ‘Foreign Bank Entry and Business Volatility: Evidence from U.S. States and Other Countries’. *Working Paper*, 9710, National Bureau of Economic Research, Cambridge, MA.

Morgan, D., B. Rime, and P. Strahan (2004). ‘Bank Integration and State Business Volatility’. *Quarterly Journal of Economics*, 119 (4): 1555–85.

Moshiriam, F. (2001). ‘International Investment in Financial Services’. *Journal of Banking and Finance*, 25: 317–37.

Paroush, J. (1995). ‘The Effect of Merger and Acquisition Activity on the Safety and Soundness of a Banking System’. *Review of Industrial Organization*, 10 (1), February.

Peek, J., and E. Rosengren (1997). ‘The International Transmission of Financial Shocks: The Case of Japan’. *American Economic Review*, 87: 495–505.

Peek, J., and E. Rosengren (2000a). ‘Implication of the Globalisation of the Banking Sector: The Latin American Experience’. *New England Economic Review*, September-October Issue, 45–62.

Peek, J., and E. Rosengren (2000b). ‘Collateral Damage: Effects of the Japanese Bank Crisis on the United States’. *American Economic Review*, 90 (1): 30–45.

Pesaran, M. H. (2004). ‘General Diagnostic Tests for Cross Section Dependence in Panels’. *Working Paper*, 0435. Cambridge: Cambridge University.

Pesaran, M. H. (2007). ‘A Simple Panel Unit Root Test in the Presence of Cross Section Dependence’. *Journal of Applied Econometrics*, 22 (2): 265–312.

Pyle, D. (1971). ‘On the Theory of Financial Intermediation’. *Journal of Finance*, 26 (3): 737–47.

Table 1: Currency and banking crisis in Latin America (1995:1-2001:4)

Country	Currency crisis	Banking crisis
Argentina		1995:1-1995:4
		2001:1-2001:4
Bolivia		1995:1-1995:4
Brazil	1999:1	1995:1-1996:4
	2001:2	
Colombia	1999:1	1998:1-1998:4
Costa Rica		1995:1-1998:1
Mexico	1995:1	1995:1-1996:4
	1995:4	

Source: Author's calculations, Carstens et al. (2004) and Caprio et al. (2005).

Table 2: Cross sectional dependence and panel unit root tests

Variable	Constant[1]	Constant plus trend[2]	CD_{lm}[3]	*Z* statistics[4]
Real GDP	X		121.5***	1.5
Real GDP		X	115.9***	-5.9***
Fiscal balance	X		42.6**	-2.5**
Fiscal balance		X	45.3**	-4.1***
Financial development	X		27.4	-6.4***
Financial development		X	29.8	-7.1***
Real private lending	X		63.8***	-3.7***
Real private lending		X	66.6***	-6.4***
Banking spread	X		34.1	-2.9***
Banking spread		X	27.6	-5.8***
Real exchange rate	X		64.6***	-3.0**
Real exchange rate		X	70.9***	-8.1***
Foreign bank presence	X		29.1	-1.3*
Foreign bank presence		X	29.0	-4.1***
Public banks presence	X		17.7	-2.9***
Public banks presence		X	21.6	-3.8***
Banking concentration	X		27.8	-1.5*
Banking concentration		X	30.2	-3.0***

Note

1. Some variables contain a constant term. The remaining do not contain deterministics and are cantered.

2. Some variables contain a constant term plus a trend. The remaining do not contain deterministics and are cantered.

3. Breusch and Pagan's test for the null hypothesis of zero cross-sectional dependence. (***), (**) and (*) denote, respectively, statistical significance at the 1, 5 and 10% levels.

4. Choi's panel unit root test, with and without cross sectional dependence according to the *CDlm* test result. The null is rejected at a significance level α when $Z < c_{z\alpha}$, where $c_{z\alpha}$ is the lower tail of the standard normal distribution or the simulated critical vale, for the cases without and with cross section dependence, respectively. The optimal number of lags included in each cross section ADF equation was determined using the Schwarz Information Criterion (SIC).

Source: Author's calculations.

Table 3: The determinants of credit dynamics in Latin America (1995:1-2001:4)
(dependent variable: HP adjusted real credit to the private sector)[1]

Mean equation	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
AR(1)	0.26	0.47***	0.49***	0.36**	0.49***	0.48***	0.45***	0.47***	0.46***	0.47***
	[1.30]	[3.47]	[4.02]	[2.43]	[4.07]	[3.98]	[3.62]	[3.72]	[3.70]	[3.74]
Real GDP (t-1)	0.88***	0.68***	0.74***	0.79***	0.72***	0.70***	0.84***	0.77***	0.77***	0.77***
	[3.56]	[3.11]	[3.28]	[3.81]	[3.21]	[3.15]	[3.96]	[3.81]	[3.85]	[3.78]
US GDP (t)	0.37									
	[0.53]									
Fed rate (t)	-1.44									
	[-1.64]									
Fiscal balance (t-1)	0.80									
	[0.75]									
Spread (t-1)	-2.63									
	[-0.99]									
Financial dev. (t-1)	1.28**	1.17**	0.83*	1.24**	0.84**	0.89**	1.13**	1.30***	1.30***	1.30***
	[2.06]	[2.36]	[1.89]	[2.67]	[1.94]	[2.05]	[2.67]	[3.55]	[3.62]	[3.55]
RER (t-1)[2]	-0.25**	-0.20**	-0.19**	-0.20**	-0.18**	-0.20**	-0.19**	-0.20**	-0.20**	-0.20**
	[-2.44]	[-2.10]	[-2.40]	[-2.76]	[-2.39]	[-2.63]	[-2.12]	[-2.30]	[-2.38]	[-2.29]
Public banks (t-1)	-0.39									
	[-1.22]									
Foreign banks (t-1)	0.23									
	[0.65]									
Banking conc. (t-1)	0.55									
	[1.61]									
BC*PB (t)[3]	1.70***	1.18**	1.17*	1.70**	1.18*	1.21*	1.54*	1.51**	1.53**	1.49**
	[3.56]	[2.72]	[1.79]	[2.64]	[1.70]	[1.86]	[1.99]	[2.11]	[2.19]	[2.10]
Currency crisis (t)	-5.45*	-4.59**	-4.99	-5.16	-4.95	-5.22***	-5.47*	-5.71**	-5.59**	-5.76**
	[-1.76]	[-2.03]	[-1.61]	[-1.63]	[-1.60]	[-4.15]	[-1.92]	[-2.18]	[-2.15]	[-2.17]
Chow test for absence of FE in mean equation[4]										
F test	1.68	1.56								
(p-value)	*0.15*	*0.18*								
Chow test for poolability of the data[5]										
F test	1.62	2.64								
(p-value)	*0.95*	*0.99*								

1. Robust *t-statistics* are reported between brackets. (*), (**) and (***) denote statistical significance at the 10%, 5% and 1% levels, respectively.

2. RER = real exchange rate.

3. BC = banking crisis; PB = public banks.

4. The null hypothesis is absence of fixed effects in a mean equation estimated by OLS.

5. The null hypothesis is that the data are poolable in a mean equation estimated by OLS.

Source: Author's calculations.

Table 4: The determinants of credit dynamics in Latin America (1995:1-2001:4)
(dependent variable: credit volatility)[1]

Variance equation	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Constant	9.64***	10.89***	2.01***	2.07***	2.01***	2.03***	2.03***	2.06***	2.08***	2.06***
	[5.72]	[6.59]	[12.07]	[12.03]	[12.04]	[12.01]	[11.01]	[12.08]	[11.55]	[11.80]
ARCH(1)	0.28*	0.21*	0.14	0.10	0.15	0.14	0.13	0.04	0.03	0.04
	[1.87]	[1.71]	[1.08]	[0.84]	[1.11]	[1.07]	[1.22]	[0.75]	[0.52]	[0.68]
BC (t)			1.18***	1.07***	1.17***	1.17***	1.22***	1.06***	1.02**	1.06***
			[3.26]	[2.88]	[3.14]	[3.19]	[3.03]	[2.86]	[2.72]	[2.85]
BC*FB (t)[2]				-0.33						
				[-1.40]						
BC*Public banks (t)					0.07					
					[0.32]					
Currency crisis (t)						-2.91				
						[-0.45]				
Foreign banks (t-1)							-0.24**	-0.20**	-0.20**	-0.19**
							[-2.49]	[-2.22]	[-2.14]	[-2.10]
Financial dev. (t-1)								-0.26**	-0.25**	-0.26**
								[-2.07]	[-2.05]	[-2.08]
Banking conc. (t-1)									-0.03	
									[-0.34]	
Public banks (t-1)										0.02
										[0.20]
Chow test for absence of FE in conditional variance equation[3]										
F test	2.47	3.51								
(p-value)	*0.04*	*0.01*								
F test for ARCH(1) effects in conditional variance[4]										
F test	13.25	2.76								
(p-value)	*0.00*	*0.099*								
F test for ARCH(4) effects in conditional variance[4]										
F test	10.21	6.33								
(p-value)	*0.00*	*0.00*								
Diagnostics										
Log-likelihood	-338.5	-363.2	-356.3	-355.2	-356.2	-354.3	-353.3	-349.4	-349.3	-349.4
Normality test (KS) (*p-value*)[5]	0.72	0.50	0.89	1.00	0.88	0.98	0.93	0.98	0.98	0.97
BG(1) *(p-value)*[6]	0.84	0.59	0.47	0.70	0.53	0.45	0.70	0.14	0.18	0.15
BG(4) *(p-value)*	0.00	0.02	0.07	0.06	0.06	0.08	0.01	0.15	0.16	0.16
ARCH(1) *(p-value)*[7]	0.29	0.51	0.54	0.45	0.53	0.45	0.70	1.00	0.98	0.98
ARCH(4) *(p-value)*	0.61	0.82	0.98	0.95	0.97	0.93	0.97	0.97	0.96	0.97

1. Robust *t-statistics* are reported between brackets. (*), (**) and (***) denote statistical significance at the 10%, 5% and 1% levels, respectively.

2. BC = banking crisis; FB = foreign banks.

3. The null hypothesis is absence of fixed effects in the variance equation. The model considered is an *ARCH(1)*.

4. The null hypothesis is the absence of *ARCH* effects.

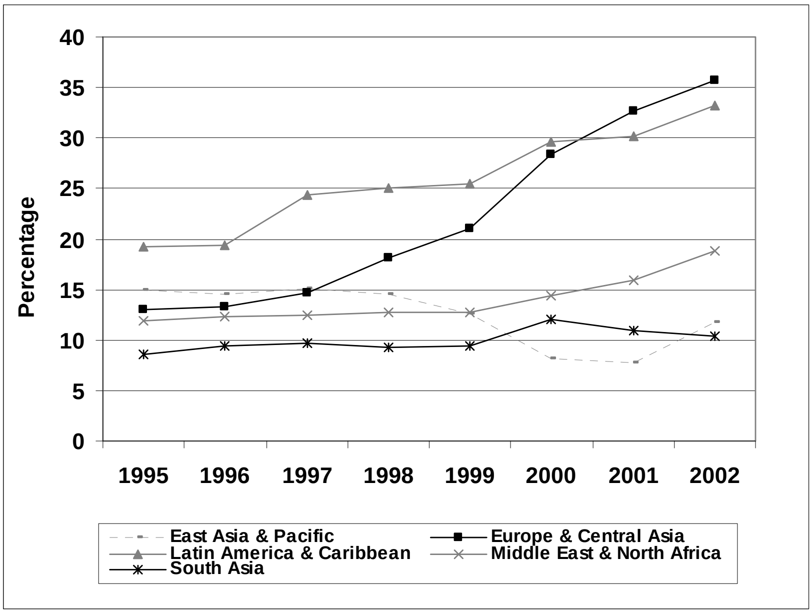
5. Kolmogorov-Smirnov normality test computed on the standardized residuals.

6. Breusch-Godfrey test for serial correlation computed on the standardized residuals.

7. Auto-regressive conditional heteroscedasticity test computed on the standardized residuals.

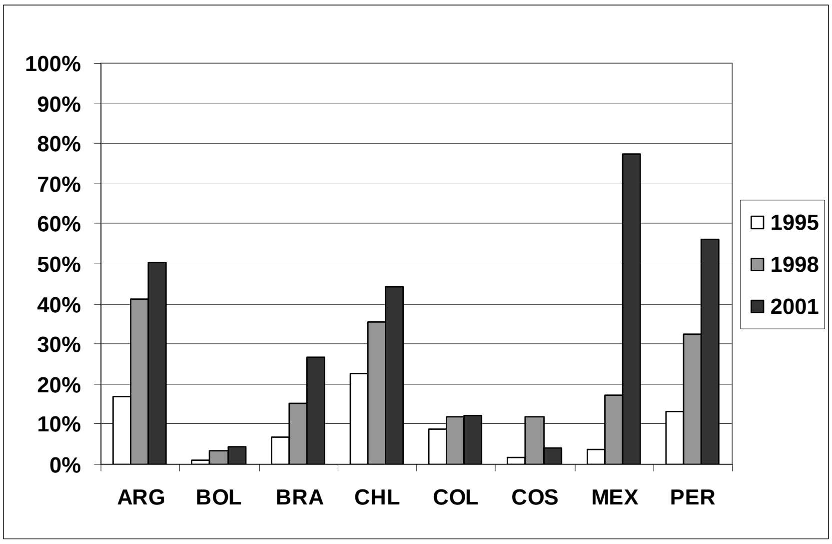
Source: Author's calculations.

Figure 1: Foreign bank participation in emerging countries, by region
(share of total banking assets)



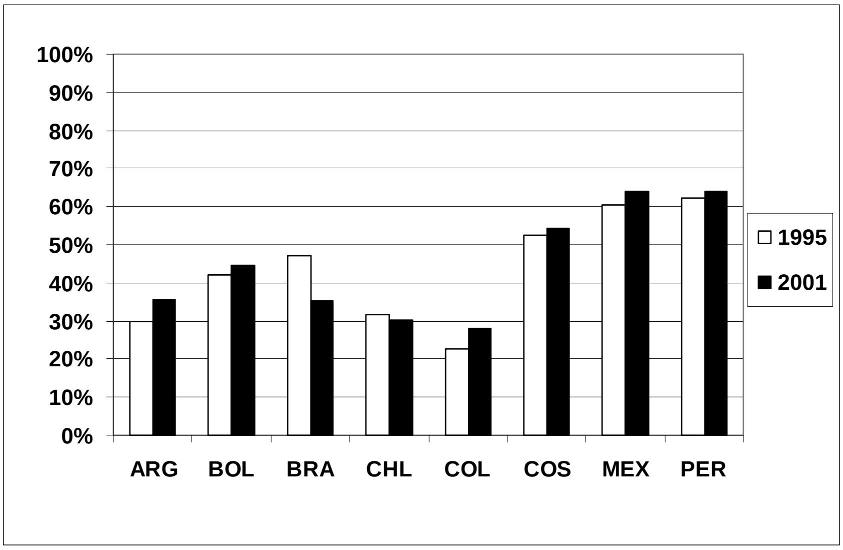
Source: Author’s calculations.

Figure 2: Foreign banks in Latin America
(share over total credit)



Source: Author’s calculations.

Figure 3: Banking concentration
(share top 3 banks on total credits)



Source: Author's calculations.

Table A.1: The determinants of credit dynamics in Latin America (1995:1-2001:4) (dependent variable: HP adjusted real credit to the private sector)[1]

Mean equation	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
AR(1)	0.80***	0.74***	0.75***	0.76***	0.76***	0.74***	0.74***	0.75***	0.74***	0.59***
	[23.94]	[18.14]	[17.89]	[18.36]	[18.66]	[18.10]	[18.00]	[18.56]	[17.39]	[7.39]
Real GDP (t-1)		0.55***	0.55***	0.47***	0.43***	0.48***	0.48***	0.49***	0.48***	0.58***
		[3.71]	[3.62]	[3.36]	[3.08]	[3.16]	[3.21]	[3.33]	[3.07]	[3.93]
GDP * FB (t-1)			0.02							
			[0.57]							
US GDP (t)				0.81*	0.86*	0.80*	0.85*	0.80*	0.77*	1.09**
				[1.80]	[1.90]	[1.75]	[1.89]	[1.76]	[1.66]	[2.42]
US GDP * FB (t)					0.05					
					[0.48]					
Fed rate (t)						0.34	0.40			
						[0.47]	[0.54]			
Fed rate * FB (t)							-0.07			
							[-0.37]			
Fiscal bal. (t-1)								-0.39		
								[-0.73]		
Spread (t-1)									0.32	
									[0.20]	
Fin. dev. (t-1)										1.16***
										[3.27]
Variance equation										
Constant	8.78***	11.47***	11.46***	10.56***	10.22***	10.61***	10.46***	10.64***	10.97***	10.41***
	[8.52]	[10.11]	[10.03]	[9.89]	[9.44]	[9.88]	[9.73]	[9.84]	[9.61]	[9.86]
ARCH(1)	1.10***	0.29***	0.29***	0.37***	0.41***	0.36***	0.38***	0.37***	0.36***	0.30***
	[7.62]	[2.65]	[2.65]	[2.82]	[2.84]	[2.79]	[2.83]	[2.77]	[2.74]	[2.61]
Chow test for absence of FE in mean equation[2]										
F test	0.40	0.46	0.43	0.44	0.43	0.44	0.51	0.36	0.60	0.34
(p-value)	0.90	0.86	0.88	0.87	0.88	0.88	0.82	0.92	0.75	0.93
Chow test for poolability of the data[3]										
F test	1.63	1.58	1.36	1.26	1.12	1.15	1.00	1.15	1.40	1.17
(p-value)	0.93	0.94	0.88	0.81	0.69	0.73	0.51	0.72	0.92	0.75
Chow test for absence of FE in conditional variance equation[4]										
F test	3.36	2.88	2.86	2.96	2.96	2.98	2.98	3.73	2.69	3.06
(p-value)	0.00	0.01	0.01	0.01	0.01	0.01	0.01	0.00	0.01	0.00
F test for ARCH(1) effects in conditional variance[5]										
F test	3.13	4.85	5.49	8.45	8.54	9.57	8.62	0.97	11.70	7.05
(p-value)	0.08	0.03	0.02	0.00	0.00	0.00	0.00	0.33	0.00	0.01
Diagnostics										
Log-likelihood	-648.7	-586.5	-586.5	-584.3	-584.1	-584.2	-584.0	-557.3	-565.4	-569.0
Normality test[6]	0.07	0.20	0.25	0.40	0.34	0.36	0.38	0.31	0.43	0.28
BG(1) (p-value)[7]	0.17	0.40	0.42	0.34	0.29	0.35	0.33	0.37	0.33	0.43
BG(4) (p-value)	0.33	0.19	0.19	0.18	0.18	0.19	0.20	0.25	0.16	0.08
ARCH(1) (p-value)[8]	0.25	0.15	0.15	0.13	0.16	0.12	0.14	0.12	0.17	0.88
ARCH(4) (p-value)	0.61	0.28	0.27	0.22	0.26	0.24	0.24	0.22	0.27	0.25

1. Robust t-statistics are reported between brackets. (*), (**) and (***) denote statistical significance at the 10%, 5% and 1% levels, respectively.
2. The null hypothesis is absence of fixed effects in a mean equation estimated by OLS.
3. The null hypothesis is that the data are poolable in a mean equation estimated by OLS.
4. The null hypothesis is absence of fixed effects in the variance equation. The model considered is an ARCH(1).
5. The null hypothesis is the absence of ARCH effects.
6. Kolmogorov-Smirnov normality test computed on the standardized residuals.
7. Breusch-Godfrey test for serial correlation computed on the standardized residuals.
8. Auto-regressive conditional heteroscedasticity test computed on the standardized residuals.

Source: Author's calculations.

Table A.2: The determinants of credit dynamics in Latin America (1995:1-2001:4) (dependent variable: HP adjusted real credit to the private sector)

Mean equation	(11)	(12)	(13)	(14)	(15)	(16)	(17)	(18)	(19)	(20)
AR (1)	0.45***	0.46***	0.45***	0.43***	0.49***	0.52***	0.49***	0.50***	0.50***	0.51***
	[4.63]	[3.59]	[4.63]	[4.23]	[5.15]	[5.44]	[4.02]	[4.07]	[4.18]	[4.31]
Real GDP (t-1)	0.70***	0.71***	0.71***	0.70***	0.62***	0.58***	0.69***	0.65***	0.64***	0.65***
	[4.85]	[3.90]	[4.91]	[4.87]	[4.07]	[3.80]	[3.28]	[3.13]	[3.11]	[3.19]
US GDP (t)	0.97***	0.67	0.98***	0.97***	1.04***	1.02	0.78	0.69	0.68	0.76
	[2.06]	[1.04]	[2.06]	[2.03]	[2.22]	[2.18]	[1.31]	[1.21]	[1.20]	[1.33]
Fin. dev. (t-1)	1.52***	1.03***	1.54***	1.60***	1.52***	1.42***	1.05***	1.16***	1.15***	1.10***
	[4.32]	[2.51]	[4.47]	[4.83]	[4.33]	[3.72]	[2.59]	[2.56]	[2.59]	[2.46]
RER (t-1)	-0.22***	-0.16**	-0.22***	-0.23***	-0.20***	-0.20***	-0.08	-0.18*	-0.19*	-0.16
	[-3.33]	[-1.94]	[-3.35]	[-3.51]	[-2.87]	[-2.71]	[-0.93]	[-1.78]	[-1.81]	[-1.54]
PB (t-1)		-0.28								
		[-1.07]								
FB (t-1)			-0.05							
			[-0.40]							
Bank conc. (t-1)				-0.18						
				[-1.32]						
BC (t)					-0.62	-0.51	-1.23			
					[-0.72]	[-0.57]	[-1.10]			
BC * FB (t)						0.26				
						[0.58]				
BC * PB (t)							1.21***	1.14***	1.13***	1.13***
							[2.89]	[2.62]	[2.59]	[2.61]
Currency crisis (t)								-4.36**	-4.49	-3.80*
								[-2.05]	[-0.85]	[-1.88]
CC * FB (t)									0.18	
									[0.10]	
CC * PB (t)										0.52
										[0.43]
Variance equation										
Constant	10.36***	11.51***	10.33***	10.44***	10.28***	10.18***	10.85***	10.67***	10.61***	10.4***8
	[9.63]	[7.21]	[9.59]	[9.62]	[9.43]	[9.39]	[5.74]	[6.21]	[6.16]	[6.05]
ARCH(1)	0.25***	0.28**	0.25***	0.24***	0.26***	0.27***	0.23	0.22*	0.22*	0.23*
	[2.55]	[2.03]	[2.55]	[2.46]	[2.54]	[2.56]	[1.52]	[1.75]	[1.77]	[1.76]
Chow test for absence of FE in mean equation										
F test	0.34	0.76	0.34	0.33	0.47	0.52	1.41	1.60	1.55	1.61
(p-value)	0.94	0.58	0.93	0.94	0.86	0.82	0.23	0.16	0.18	0.16
Chow test for poolability of the data										
F test	1.44	1.43	1.55	1.23	2.15	2.62	46.63	2.33	1.92	1.92
(p-value)	0.94	0.91	0.98	0.83	1.00	1.00	1.00	1.00	0.99	0.99
Chow test for absence of FE in conditional variance equation										
F test	2.92	6.59	2.87	2.91	3.12	2.80	3.46	3.65	3.62	3.65
(p-value)	0.01	0.00	0.01	0.01	0.00	0.01	0.01	0.00	0.00	0.00
F test for ARCH effects in conditional variance										
F test	12.73	8.90	11.88	17.63	0.11	0.51	0.00	3.24	3.23	3.25
(p-value)	0.00	0.00	0.00	0.00	0.74	0.48	0.95	0.07	0.07	0.07
Diagnostics										
Log-likelihood	-564.4	-384.6	-564.2	-564.0	-542.8	-543.63	-364.4	-362.5	-362.4	-362.1
Normality test	0.32	0.52	0.35	0.41	0.31	0.39	0.72	0.51	0.54	0.62
BG(1) (p-value)	0.43	0.71	0.43	0.46	0.47	0.46	0.54	0.63	0.60	0.62
BG(4) (p-value)	0.02	0.84	0.02	0.01	0.00	0.01	0.24	0.03	0.04	0.05
ARCH(1) (p-value)	0.72	0.27	0.69	0.61	0.84	0.85	0.74	0.51	0.50	0.49
ARCH(4) (p-value)	0.42	0.49	0.44	0.47	0.33	0.30	0.83	0.78	0.76	0.86

Source: Author’s calculations.