

Tropical Bubbles

Asset Prices in Latin America, 1980–2001

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The authors examine if observed asset prices in Latin America depart significantly from fundamentals-determined levels. These departures, or bubbles, are found to be equally determined by both country-specific and common external variables, contrary to previous studies that found that local factors were predominant in asset price determination in Latin America.



Summary findings

Herrera and Perry test for the existence of asset price bubbles in Latin America in 1980–2001, focusing mainly on stock prices. Based on unit root and cointegration tests, they find that they cannot reject the hypothesis of bubbles. They arrive at the same conclusion using Froot and Obstfeld's intrinsic bubbles model.

To examine empirical regularities of these bubble episodes in the region, the authors identify periods of significant stock price overvaluation. They quantify the relative importance of different factors that determine the probability of bubble occurrence, focusing on the contrast between the country-specific variables and the common external factors. They include as country-specific variables both the level and the volatility of domestic credit growth, the volatility of asset returns, the capital flows to each country, and the terms of trade. As common external variables, they consider the degree of asset overvaluation in the U.S. stock and real estate markets and the term spread of U.S. Treasury securities. To quantitatively assess the relative importance of each factor, they estimate a logit model for a panel of five Latin American countries from 1985 to 2001.

In general, the authors find that the marginal probabilities of common and country-specific variables are of roughly the same order of magnitude. This finding contrasts with those of previous studies that real asset returns in Latin America are dominated by local factors.

Finally, the authors explore the main channels through which asset prices affect real economic activity, with the most important being the balance sheet effect and its impact on bank lending. They show how the allocation of bank lending across different sectors responded sensitively to real estate prices during the boom years in countries that experienced banking crises. Thus asset price bubbles have long-lasting effects in the financial sector and, through this channel, on growth. Another channel through which asset prices—particularly stock market prices—affect long-run growth is through their effect on investment. The authors find a strong positive association between stock prices and investment and a negative effect of stock price volatility on investment. An additional motive for the central bank to monitor asset prices is the general coincidence of the crash episodes identified by the authors with currency crises in the region in the past two decades.

This paper—a product of the Office of the Chief Economist, Latin America and the Caribbean Region—is part of a larger effort to understand asset pricing in the region, and to examine the relationship between asset prices and other variables, such as domestic credit and capital flows. Copies of the paper are available free from the World Bank, 1818 H Street NW, Washington, DC 20433. Please contact Ruth Izquierdo, room I8-012, telephone 202-458-4161, fax 202-522-7528, email address rizquierdo@worldbank.org. Policy Research Working Papers are also posted on the Web at <http://econ.worldbank.org>. The authors may be contacted at sherrera@worldbank.org or gperry@worldbank.org. November 2001. (58 pages)

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Tropical Bubbles: Asset Prices in Latin America, 1980-2001¹

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I. INTRODUCTION AND SUMMARY

The past decade began in a festive mood in Latin America, with policymakers welcoming the future that had finally arrived with its rewards: growth, rising asset prices, rising investment-to-GDP ratios, strong capital inflows, and booming domestic credit. The decade ended, however, on a gloomy note: stagnant growth, falling investment ratios, stagnant or falling asset prices, sluggish domestic credit, and several crippled banking systems. How do we account for this transition, and can we explain the interactions among all these variables? In particular, what is the relationship between the credit boom (external and domestic) and the behavior of asset prices? Is the bust related to the boom? Is the behavior of asset prices governed by fundamentals, or do asset prices in Latin America reflect the presence of bubbles, as in other emerging market economies (Sarno and Taylor, 1998)? To address these issues, we divide the present paper into three main chapters after this introduction.

Chapter II describes the evolution of asset prices in Latin America, summarizes the statistical tests for the existence of bubbles, and identifies chronologically the periods of significant asset overvaluation, as well as the crash episodes. This analysis is based mostly on data for stock market prices, and, to a lesser extent (because of limited data availability), on real estate prices. To identify periods of asset overvaluation we used Froot and Obstfeld's (1991) intrinsic bubbles model as a benchmark, given its tractability and parsimony. This chapter concludes with a generalized rejection of the no-bubbles hypothesis and with the identification of significant stock overvaluation episodes that were synchronized across the region only in the early 1990s. Most bubbles were followed by stock market crashes, with the resulting prices being lower than those before the bubble formed. Crash episodes also tended to lead or coincide with currency crises documented in several papers on this related topic (Kaminsky, Lizondo and Reinhart, 1997).

Chapter III identifies empirical regularities that characterize the significant overvaluation periods. We quantify the relative importance of different factors that determine the probability of bubble occurrence, focusing on the contrast between the country-specific variables and the common external factors. We included in the country-specific variables (a) both the level and the volatility of domestic credit growth, (b) the volatility of asset returns, (c) capital flows to each country, and (d) the terms of trade. As common external variables, we considered (a) the degree of asset overvaluation in the stock market and the real estate market in the U.S., and (b) the term spread of U.S. Treasury securities. To quantitatively assess the relative importance of each factor, we estimated a Logit model for a panel of five Latin American countries from 1985 to 2001. In general, we found that the marginal probabilities of both common and country-specific variables were roughly of the same order of magnitude. This contrasts with previous studies that found that real asset returns in Latin America were dominated by local factors (Harvey, 1995; Hargis and Maloney, 1997). The single most important determinant of bubbles was the U.S. term spread as it predicts future real interest rates and future economic activity in the

U.S. Domestic credit growth was the second most important factor. Other relevant factors were the volatility of credit growth and the volatility of real asset returns, as predicted by recent models that explain bubbles based on financial intermediary behavior (Allen and Gale, 2000).

Chapter IV explores some empirical channels through which asset prices affect real economic activity. One of the main channels through which asset prices may affect activity is the balance-sheet effect and bank lending. We investigate and present evidence about how the allocation of bank lending across different sectors responded sensitively to real estate prices during the boom years in countries that experienced banking crises. Thus, asset price bubbles have long-lasting consequences in the financial sector, and, through this channel, on growth. Another channel through which asset prices—in particular stock market prices—affect long-run growth is their effect on investment. We found (a) a strong positive association between stock prices and investment, and (b) a negative effect of stock price volatility on investment. Regarding stabilization policy, stock prices in Latin America also provide information about future consumption and investment. To the extent that stock prices incorporate information about future aggregate demand, the central bank may benefit from monitoring them as a means of anticipating the future course of inflation. An additional motive for the central bank to monitor asset prices is the general coincidence of the crash episodes identified in this paper with currency crises experienced in the region in the last two decades.

Chapter V draws conclusions and discusses some policy implications.

II. DO TROPICAL BUBBLES EXIST? WHAT DO THEY LOOK LIKE?

A. Evolution of asset prices in Latin America²

The evolution of stock prices in Latin America (LAC) since the 1980s (Graph 1) can be divided in two sub-periods:

1. During the 1980s, prices (in real terms) showed no clear trend, and there were some spikes in Argentina, Brazil, and (to a lesser extent) Mexico. The price spike episodes of the 1980s were preceded by high and rising inflation.
2. During the 1990s, there was a substantial price increase in the early years that was stabilized, and even reversed in the latter part of the decade. Only in Brazil did the rising trend continue throughout the whole decade.

Another feature of stock prices in LAC during this period was their recurrent crashes, particularly in Argentina, Brazil, and Mexico. In Chile and Colombia, these sudden falls were less frequent.

² A full description of the variables used in this paper and their sources is presented in Appendix A.

Returns in stock markets across the region (Graph 2) were positively and highly correlated (table 1). The highest correlation coefficients are between Chile and Mexico (0.92), Colombia and Mexico (0.87), Chile and Colombia (0.86), Argentina and Chile (0.82), Mexico and Argentina (0.81). Across time, however, these correlations seem to have changed. computing the correlation coefficients with a 24-month rolling window, we observe that, for the biggest three countries, these are stationary; however, there is a structural change in the generating processes in the early 1990s (Graph 3).³

Finally, there is a positive association between stock prices and dividends (normalized) in most countries (Graph 4). Both variables tend to move together, although there are periods when they diverge. Colombia seems to be the only case in which the series are clearly divergent.

Comparing these results with those from East Asian countries—Indonesia, Korea, Malaysia, and Thailand—we find that, for most countries, correlations of stock returns are around 0.50 or lower. The only exception is the correlation between Malaysia and Thailand of 0.86 (Table A1 in Appendix A). Thus, stock return correlations are lower in Asia than in LAC. As for change in these East Asian correlations over time, there is no clear evidence of structural breaks in the correlations as was the case in the big three LAC countries (Graph A1 in Appendix A.)

Regarding real estate prices (Graph 5), we constructed proxies by using the rent component of the CPI, deflated by the overall index. We were able to build this proxy for Argentina, Colombia, and Mexico, which were the only countries where we had data availability since the 1980s. In the case of Mexico, there is a strong correlation between this proxy and the index of the value of urban land in Mexico City constructed by Banco de Mexico (Guerra, 1997). A boom in real estate prices occurred in Mexico in 1988–91, in Argentina in 1990–93, and in Colombia in 1993–96. We also present real estate prices in two affluent neighborhoods of Santiago de Chile in 1975–82,⁴ as reported by Conley and Maloney (1995).

Graphs 6a and 6b show the evolution of returns on real estate investment in Mexico and the U.S., as captured by our proxies.⁵ For the whole sample period, there is no correlation (Graph 6a); however, there is a correlation of 0.70 after 1995 (Graph 6b).

B. Testing for Bubbles: The Latin American Case

1. Overview of bubble testing

³ The tests for stationarity and structural change were those of Perron (1994). For the correlation between Brazil and Mexico, the structural change is in 1991:11, while that for Argentina and Mexico is in 1991:06. The Argentina-Brazil break occurs at an earlier date—1985:06.

⁴ The data were extracted weekly from *El Mercurio*. The authors thank Bill Maloney for making the data available.

⁵ For the U.S. we consider the return index of real estate investment trusts (REITS) as computed by the NAREIT. For Mexico, it is the y-o-y growth rate of the housing price index we constructed.

Theoretic and empirical work on bubbles is vast. We will briefly summarize the more relevant work for our purposes, given that well-known surveys have already been done: Camerer (1989) performed a general survey of both the theoretical and empirical literature, while Flood and Hodrick (1990) concentrated on the empirical branch. A more recent survey of both approaches can be found in Campbell (2000).

In general, a bubble (B_t) is defined as the difference between the fundamentals-determined price (P^{PV}) and the observed price (P_t). In the case of stocks, the fundamentals price can be expressed as the sum of discounted expected future cash flows—or dividends—to the holder of the asset.⁶

$$P_t = P_t^{PV} + B_t \quad (1)$$

The bubble term, B , if it exists, is expected to grow at the real rate of interest.⁷

The bulk of the bubble-testing literature falls into three types of tests. The first type examines the relationship between (a) the observed price and (b) the present-value price or the fundamentals used to forecast it. For example, tests of the bubble hypothesis in exchange rates examine the existence of a long-run equilibrium (cointegrating) relationship among exchange rates, money supplies, and prices (Meese, 1986; Chin and Meese, 1995; and Mark, 1995). Tests of the bubble hypothesis in stock markets examine the existence of equilibrium (cointegrating) relationships between prices and dividends (Campbell and Shiller, 1987 and Campbell, Lo, and McKinlay, 1997). A second type of bubble test compares the volatilities of observed prices and the present-value prices (Shiller, 1981; Le Roy and Porter, 1981; and West, 1988). The third type of test is more elaborate and indirect; it estimates a reduced-form price equation by two alternative methods and verifies whether the parameter values are the same.

The first method is a simple (instrumental variables) projection of prices onto the information set of the fundamentals forecasting. The second method is a simultaneous estimation of two equations: The first is the fundamental forecasting process, and the second is a reduced-form equation that assumes no bubbles and imposes cross-equation restrictions as in Hansen-Sargent (1981). If certain parameters are estimated as being the same by both methods, then the no-bubble hypothesis cannot be rejected. This approach, pioneered by West (1987) for stock market prices, has been used by Casella (1989) for price level bubbles and by Meese (1986) for exchange rate bubbles.

Testing for bubbles is plagued with problems and obstacles. At a theoretical level, reconciling rational behavior with the existence of bubbles is not trivial, and the conditions under which they can emerge are quite restrictive (Santos and Woodford,

⁶ $P_t^{PV} = \sum_{s=t}^{\infty} e^{-r(s-t+1)} E_t[D_s]$ is one solution to the equation $P_t = e^{-r} E_t[D_t + P_{t+1}]$. It is the

particular solution attained after imposing a transversality condition.

⁷ The bubble term is generally described as $B_{t+1} = B_t(1+r) + b_{t+1}$, where b_{t+1} is the innovation in the bubble at time $t+1$ (Flood and Hodrick, 1990).

1997). This is an ongoing topic of debate in several papers presented in this conference and in the academic literature (Shiller, 2001). To reconcile rational behavior with bubbles in asset prices, researchers shifted to some type of market imperfection to explain this apparently less-than-rational outcome. From this perspective, the Allen-Gale (2000) model is of particular interest, because it links asset price bubbles to financial intermediary behavior. In this model, bubbles⁸ arise because of an informational asymmetry between the borrowers and the bank: given the limited liability of borrowers, they bid up the price of assets that are in fixed supply, with the cost of the excessive risk being borne by the intermediary. In this setting, the volatility of real returns affects the probability (and magnitude) of bubbles. The magnitude and uncertainty of credit expansion is also an important determinant of bubbles in this model. Other models (Krugman, 1998) also link bubbles to financial sector developments, but, in most cases, the results are based on implicit guarantees that lead to risk-shifting. The Allen-Gale model does not require these conditions.

Other problems with bubble tests arise from the empirical side. First, the procedures are generally a joint test of a fundamentals-forecasting equation and the transversality condition. Therefore, the rejection of “no bubbles” implies a rejection of the underlying model and the transversality condition (Meese, 1986). Related to the issue of the underlying model, Flood and Garber (1980) pointed out, in their seminal paper on testing for bubbles, how omitted variables may explain the behavior of prices that the analyst incorrectly concludes have dynamics independent from fundamentals, biasing the test towards rejection of “no bubbles.” The finding that bubble behavior is observationally equivalent to that of expected future regime changes (Hamilton and Whiteman, 1985 and Flood and Hodrick, 1986) points in the same direction. Hence, the inability to reject the “no bubble” hypothesis can be consistent with a non-bubble present value price whose fundamentals-generating processes are expected to change in the future.

2 Tests for the asset price bubble hypothesis in Latin America

Bubble tests applied to the Latin American case are very scarce. Among the most recent treatments of the topic is an application of a test to Argentina’s exchange rate, price level, and money stock (Hall, et.al., 1999). The paper rejects the no-bubble hypothesis in the series for a time period that matches our own findings for stock prices, described in the next section. A second paper describes the specific topic of bubbles in housing prices in affluent neighborhoods of Santiago de Chile in the late 1970s and early 1980s (Conely and Maloney, 1995). Nonetheless, this paper falls short of any empirical testing of the hypothesis, since that was not the paper’s objective. Two papers on the related topic of predictability of asset returns in emerging markets (Harvey, 1995 and Hargis and Maloney, 1997) reach opposite conclusions on the relationship between asset prices and

⁸ In the Allen-Gale model, the fundamentals-determined price is defined as the price that would prevail in the absence of risk-shifting, i.e., in a perfect information setting and with no limited liability in the debt contract.

fundamentals: (a) the Harvey paper concludes that standard asset pricing models produce large errors, fail to price assets correctly, and are unable to account for time variation in expected returns; (b) on the other hand, the Hargis and Maloney paper concludes that equity markets in Latin America (and East Asia) do reflect expected cash flows and correctly incorporate domestic and global shocks.

With the caveats derived from the previous section, we choose the tests with the simplest structure as proposed by Campbell, Lo, and McKinlay (1997), and those used to test for bubbles in other emerging market economies (Sarno and Taylor, 1999). The general idea is to verify or reject the existence of a stable (non-explosive) relationship among stock prices, dividends, and returns.

The equation that establishes the basis for the tests (see Appendix B for a derivation) is:

$$d_t - p_t = -\frac{k}{1-\rho} + E_t \left[\sum_{j=0}^{\infty} \rho^j (-\Delta d_{t+1+j} + r_{t+1+j}) \right] \quad (2)$$

where d_t =log dividends

p_t = log prices

r_t = return

Given the accounting identity nature of the above equation,⁹ if prices go up, either dividends go up, or expected future returns go down to maintain the dividend-to-price ratio stationary. Hence, the tests are oriented toward examining the stationary (or explosive) behavior of the log dividend-price ratio and the existence of a stable relationship among dividends, prices, and returns.

Accordingly, we perform two types of tests. First, we check for unit roots in the log dividend-price ratio and in the real return series. If dividends follow an I(1) process, their difference is stationary. Then, the return series must be of the same order of integration as the dividend-price ratio. The idea is that, if the series have unit roots, the “no bubble” hypothesis is rejected. Second, we check for a cointegrating relationship between the log dividend-price ratio and returns. If a stable (equilibrium) relationship is rejected, then the “no bubble” hypothesis is also rejected.

a. Unit root tests

Standard Augmented Dickey-Fuller tests (ADF) for series with bubbles may result in accepting stationarity, even though the series are in fact explosive, because the recurrent collapse of the bubble may resemble a mean reversing process (Evans, 1991). Results for the ADF (unit root) tests are in table B1 of Appendix B. For most countries, total return and (log) dividend yields have a unit root. The standard ADF tests reject the unit root hypothesis in Argentina (both series) and Mexico (dividend yield), although the results

⁹ This point is emphasized by Campbell (2000), noting that this equation is derived from an identity, solved forward imposing the restriction of a transversality condition and taking expected values.

are not robust to the inclusion of a deterministic time trend. Hence, for most countries, the bias toward accepting stationarity due to periodically collapsing bubbles was not a problem, except for the possible exceptions of Argentina and Mexico.¹⁰

To correct the potential bias in the standard unit root (and cointegration) tests, Taylor and Peel (1998) propose a residuals-augmented least squares (RALS), briefly described in Appendix B. The main idea is to introduce into the standard ADF tests an auxiliary term that mitigates the skew and kurtosis originating from the collapsing prices.

Results for the RALS unit root test are shown in Table B.3 (Appendix B), and summarized in table 2. These results are similar to those obtained with the ADF test. Total returns and log dividend yield are non-stationary series in most countries (Brazil, Chile, Colombia, and Peru). However, the unit root hypothesis is rejected for stock returns in Argentina and for the dividend yield in Mexico.¹¹ Hence, in most countries we are unable to reject the unit root hypothesis, implying bubble behavior in stock prices during the sample period. We proceed to verify these conclusions with the cointegration tests.

b. Cointegration tests

To verify (or reject) the hypothesis of the existence of a long-run relationship between the log dividend-price ratio and real returns, we use three alternative tests: the standard Johansen cointegration test, the RALS cointegration test, and the autoregressive distributed lag (ARDL) method. The ARDL has the advantage of circumventing the issue of pre-testing for unit roots, given that the tests are valid regardless of the order of integration of the variables (Pesaran, et. al., 1999).

Table 2 summarizes the results.¹² Both the standard Johansen and the RALS cointegration tests reject the hypothesis of a long-run relationship between total returns and the log of dividend yield in all the countries.¹³ Finally, the ARDL method rejects the existence of a long-run relationship, except for the case of Argentina, in the particular case in which the regression includes time trend.

Thus, the main results and conclusions of this section can be summarized as follow:

- Returns are non-stationary when performing ordinary ADF tests. This result is confirmed by the Taylor-Peel RALS test, with the possible exception of Argentina, when a deterministic time trend is included.

¹⁰ The rejection of the unit root hypothesis in both Argentina and Mexico depends on whether the trend or the constant is included. In Argentina when a constant, but no trend, is included, the unit root cannot be rejected. This is simply noted to show the ambiguity of the results concerning the “no bubble” hypothesis in these two countries.

¹¹ Argentina’s result is sensitive to the inclusion of a constant and a time trend in the regression of the null hypothesis.

¹² See tables B2 to B5 in appendix B for detailed results.

¹³ The Johansen test (table 2) shows two cointegrating vectors for Argentina and Mexico. This problem (given that the test is for cointegration among two variables) is not present in the RALS test.

- Log dividend yield is non-stationary with ADF and the RALS tests, with the possible exception of Mexico, when a deterministic time trend is included in the regression.
- Log dividend-yield and returns are not cointegrated according to the standard Johansen tests, and, using the ARDL method, we reject the hypothesis of a long-run relationship between these variables. Argentina is an exception in the case in which the null includes a time trend.

All of the above tests lead to the general rejection of the hypothesis of no-bubbles in stock prices in this group of Latin American countries between 1980 and 2001, with some ambiguity in Argentina's case.

C. When have bubbles occurred, and how persistent are they?

The previous section rejected the “no bubble” hypothesis for stock prices in Latin America. This section aims at identifying periods when observed prices differed significantly from the fundamentals-determined price, based on Froot and Obstfeld's (1991) intrinsic bubbles model. This model is useful to explain persistent over- or under-valuation and has the advantage of being parsimonious. In addition, as a direct application, the authors separate the present value component from the bubble component, which is our objective. In its simplest version, bubbles depend on fundamentals (dividends) only. In more elaborate versions, bubbles can be made to be functions of time.

Froot and Obstfeld set up the basic model of the stock price-dividend relationship¹⁴ and arrive at the following expression for stock prices described in equation (1):

$$P_t = P_t^{PV} + B_t$$

where P_t , the stock price, is equated to P_t^{PV} , the present discounted value of expected future dividend payments,¹⁵ and B_t , the bubble term (if it exists), is expected to grow at the real interest rate.¹⁶

To find a simple, closed-form solution for P_t^{PV} , some assumption has to be made about future dividend growth or the dividend-generating process. Froot and Obstfeld show that if the log dividend process is generated by

$$d_{t+1} = \mu + d_t + \xi_{t+1} \tag{3}$$

¹⁴ $P_t = e^{-r} E_t [D_t + P_{t+1}]$

¹⁵ $P_t^{PV} = \sum_{s=t}^{\infty} e^{-r(s-t+1)} E_t [D_s]$

¹⁶ $B_t = e^{-r} E_t [B_{t+1}]$

where μ is the trend growth in dividends and ξ_t is a $(0, \sigma^2)$ normal random variable, then the present value price is proportional to dividends:

$$P_t^{PV} = \kappa D_t \quad (4)$$

where $\kappa = (e^r - e^{\mu + \sigma/2})^{-1}$

r = return on stocks over the whole sample period, and

μ, σ are, respectively, the trend growth rate of $\log(\text{dividends})$ and the standard deviation of the residuals of an AR(1) process describing the dividends (equation 3).¹⁷

Since, for each country, we have r and can obtain estimates of μ, σ from a regression of equation (3), we calibrate the value of κ .¹⁸ Hence, the bubble component can be approximated by the difference between the observed price and the P^{PV} .¹⁹ Graph 7 shows the (normalized) observed prices and the present-value prices for Argentina, Brazil, Chile, and Mexico. This simple approximation shows that during the 1990s there were persistent deviations of the observed price from the estimated present-value price. Argentina is the only exception, but there were significant, although very transitory deviations in the 1980s.

Froot and Obstfeld estimate a more complicated version of their model, where bubbles depend not only on dividends, but also on time,²⁰ of the following type:²¹

$$P_t/D_t = c_0 + c_1 D_t^{\lambda-1} + c_2 D_t^{\lambda-1} + g X_t \quad (5)$$

¹⁷ Sellin (1998) also uses this assumption on the dividend process to obtain a simple solution for the present-value price.

¹⁸ Appendix C, Table C1 shows coefficients for Latin American countries and the U.S.

¹⁹ An important assumption is that dividends at time t are known before setting the price. This way, any "news" effect is captured immediately by the present-value price. Empirically, however, because of timing issues, (distributed) dividends not necessarily reflect all the news, we might have a problem of identifying as a bubble something that is simply the price responding to news that the dividend series did not incorporate. This methodological problem translates into the fact that prices should contain useful information to forecast future dividends. To examine the extent of this problem we ran Granger causality tests between the (differenced) log prices and log dividends up to twelve lags, and only in Argentina did we find robust significant evidence of prices Granger-causing dividends. In the other four countries we reject the hypothesis that prices contain useful information beyond that in current dividends. Froot and Obstfeld reject causality from prices to dividends in the case of the U.S.

²⁰ Their simplest model, where bubbles depend on fundamentals (dividends) only, is of this type:

$B(D_t) = cD_t^\lambda$ where λ is the positive root of the dynamic system $\lambda^2 \sigma^2 / 2 + \lambda \mu - r = 0$ and c is a constant. By adding the present value price and the bubble term we get $P(D_t) = P_t^{PV} + B(D_t) = kD_t + cD_t^\lambda$ which is a price with the bubble driven by fundamentals.

²¹ This is the model reported in Froot and Obstfeld's Table 4, augmented with the negative root part as indicated in their footnote 28.

where P = prices

D = dividends

X = a linear term on Dividends or a time trend. We included a quadratic time trend also.

λ, λ^1 are the positive and negative roots, respectively of the dynamic system

$\lambda^2 \sigma^2 / 2 + \lambda \mu - r = 0$ described above. The estimation of the nonlinear equation was done restricting the values of these roots.

This model, besides allowing the quantification of the divergence of observed values from the predicted values, is useful to verify the existence of bubbles. The null hypothesis of no bubbles implies that $c_1 = 0$, while the alternative implies $c_1 > 0$. The estimations (Appendix C) lead to rejection of the “no bubble” hypothesis in all the cases except Argentina. In all cases (Graph 8), deviations of the observed price from the predicted one are persistent, frequent, and significant in size. Argentina is the only country that does not have a bubble after 1992. These results are similar to those of the simpler model. The timing of the bubbles tends to coincide in both approximations.²² In Argentina, there is evidence of bubbles during the periods of hyperinflation from 1985 to 1989. Hall, Psaradakis, and Sola (1999) identify bubbles for money supply, exchange rate, and prices in 1985 and 1989, years for which we also find bubbles in the stock market.

Following the same procedure, we estimate the intrinsic bubble model for the U.S. and obtain interesting results (Graph 9). First, we cannot reject the hypothesis of bubble behavior, similar to Froot and Obstfeld’s result and to the more recent ones of Bonds and Cummins (2000). Second, the deviations of the observed price from the predicted one show two periods of significant stock price overvaluation. One begins in 1986 and ends in 1987 (the crash of 1987). The second one occurs in the second half of the 1990s. The significant deviation of late 1997 was (temporarily) corrected in 1998, while the overvaluation in 2000 had not been corrected by the end of the year.²³ These results are very similar to Reinhart’s (1998) quantification of overvalued stock prices in U.S.²⁴

Having quantified the over/under valuation of assets in the Latin American countries and the U.S.,²⁵ we examine the correlation of these measures across time. For the period

²² The residuals are the difference between the observed price and the predicted price. Since the predicted price already has a bubble component, this approximation underestimates the size of the bubble and might lead us to overlook a bubble episode when it actually existed. This residual can be considered as the innovation in the bubble.

²³ The data come from Robert Shiller’s book, *Irrational Exuberance*, and can be found on his webpage. He has dividend data until December 2000, and hence the model was estimated until that date.

²⁴ Reinhart uses a cointegration approach to verify the equilibrium relationship between the price-to-earnings ratio and fundamentals for the period 1980–96. His results of departures from the predicted model are summarized in Figure 1 of his paper.

²⁵ We tried to quantify for the East Asian countries the degree of departure from the benchmark cases, and we did not obtain any significant departures after 1995. This is a strange result, given that we reject, as Sarno and Taylor do, the no-bubbles hypothesis. The only explanation seems to be that all the bubbles happened before 1995. Future versions of this paper will present this, together with the crashes in East Asian countries.

1980–2001 (Table 3) we note very low correlation coefficients (recall that stock returns were highly correlated). A second striking feature is the negative sign between deviations from the model in the U.S. and in the Latin countries, except in Mexico, where it is negligible. These correlations are significantly lower than those found among industrialized nations for a similar time period (IMF, 1998).²⁶ For example, the correlation coefficients were of the following order: U.S.-Canada, 0.58; U.S.-U.K. 0.65; U.S.-Germany, 0.44; Germany-France, 0.49.

Restricting the sample period to 1995–2001, we note that correlation coefficients increase, with the most notable changes being the rise in the Mexico-U.S. coefficient to 0.40, and the negative coefficient of Brazil-U.S., -0.27 (Table 4). The negative correlation between the Brazilian data and other Latin American countries is also found by others (Harvey, 1995), although for a different sample period.

Based on this section's results, we build a dummy variable for bubble periods, which will be used in following sections. The dummy takes the unit value whenever there is a significant departure of stock prices from their present-value price or from the price predicted by the intrinsic bubble model.²⁷ Leaving the negative deviations aside,²⁸ we labeled these as bubble episodes (Table 5), and will use these in subsequent sections.

Regarding real estate prices, apart from noting periods of rapid growth, we are unable to identify departures of observed prices from the fundamentals, because we are using one of the most important fundamentals, rent, to build our price proxy.²⁹

D. Most (but not all) bubbles end in crashes, and there are crashes without bubbles.

To describe the evolution of stock prices and determine crash periods, we follow Patel and Sarkar (1998) in the construction of an auxiliary variable defined as the ratio of the stock market price to the maximum value of the series up to that time period. This variable, referred to as CMAX, is bound to take values between zero and one. Patel and Sarkar define a crash period for Latin America (treated as a whole region) when the price index falls more than 35% relative to the historical maximum. One of the advantages of this variable is its wide use among equity market practitioners, which facilitates the interpretation of some results. Another advantage is that it provides a date for the onset of the crash and a date for the trough, i.e., the point at which the price index reaches its minimum level during the crisis. Finally, we have a benchmark with which our results can be compared.

²⁶ In the 1998 *World Economic Outlook*, the IMF computed deviations of actual prices from predicted stock prices and calculated the correlations between 1985 and 1999.

²⁷ Appendix C presents the details.

²⁸ The literature has generally ruled out the existence of negative bubbles, because they would imply negative stock prices (Diba and Grossman, 1988 and Froot and Rogoff, 1991). However, Allen and Gale (2000b) consider the case where stock prices fall below their fundamental values.

²⁹ In the case of the U.S., we consider deviations of the proxy from a Hodrick Prescott trend in a later part of the paper.

Once the CMAX variable was computed (see Appendix C) we endogenously determined the country-specific thresholds such that, if the CMAX fell below this level, the episode was catalogued as a crash. The endogenous threshold selection was done by means of the Self-Exciting Threshold Autoregression (SETAR) mechanism (Potter, 1995).³⁰ In this fashion, events were classified as crashes if the CMAX fell below the following threshold levels: Argentina, 31%; Brazil, 63%; Chile, 79%; Colombia, 87%; and Mexico, 70%.³¹

As an alternative to working with the CMAX variable, we used the monthly percentage changes of the stock price indexes in real terms and determined endogenously the country-specific threshold using the SETAR method. The monthly percentage falls exceeding the following threshold levels were considered crash episodes: Argentina, 34%; Brazil, 22%; Chile, 11%; Colombia, 11%; and Mexico, 11%.

Having determined crash periods with the two methods, we combine this information with the bubble periods for Argentina, Brazil, Chile, Colombia, and Mexico. Graph 10 and Table 5 show results for bubbles and crashes identified by each of the methodologies. Crash type 1 identifies crash periods with the CMAX, while crash type 2 relates to the crashes identified with the monthly percentage changes of the stock prices.

We observe that there are more crashes than bubbles. While there are 22 bubbles in the 1980–2001 period, there are between 24 and 41 crashes depending on the method of identification. Bubbles and crashes have similar average durations: bubbles persist for 8 months, while crashes last 10 months.³² The average price increase during bubbles is 173%³³, while the average price fall in crashes is 181 %.

Most bubbles end in crashes: 14 out of 22 bubbles (64%) end up bursting. In these types of bubble episodes, which are clearly associated with crash episodes, generally the stock market price resulting after the crash is lower than that prevailing before the bubble; in 12 out of the 14 episodes, prices were lower after the crash than before the bubble.

We also note that there are crashes not preceded by bubbles. Most of these episodes are related to foreign crises. Brazil, Chile, Colombia, and Mexico show crashes not preceded by bubbles in the periods of the Asian crises (1997:10) and the Russian crisis (1998:08). The U.S. stock market crash in 1987 is also reflected in crashes in Brazil, Chile, and Mexico.

³⁰ Patel and Sankar worked in nominal figures, while we use real stock prices. The variable is regressed against its own lags, each time with a different dummy “threshold” variable. The threshold selected corresponds to the regression with a lower likelihood ratio or Akaike information criterion. See Appendix C for a detailed explanation and tables of the CMAX variables for each country.

³¹ In the U.S., the threshold CMAX was 94%.

³² The crash duration refers to the time elapsing between the point when the threshold is surpassed and the moment the CMAX reaches a minimum.

³³ This percentage increase is computed by comparing the prices at the beginning and end of the bubble. However, the price increase is larger if you consider the percentage change between the maximum price during the bubble and the beginning price. This happens because the maximum price is reached before the bubble ends.

Similarly, most countries register a crash in the first or second quarter of 1994, when there was a shift in U.S. monetary policy: Between February and June the Federal Reserve funds rate was increased 4 times, signaling a clear change in policy direction. According to these results, the Tequila crisis can be thought of as beginning with the bursting of the bubble in the Mexican stock market by the second quarter of 1994 and the intervention of two banks in September 1994, before the well-known episodes that followed. Later, in the first months of 1995, after the devaluation of the peso, there is evidence of a second crash in Mexico. In Brazil, Chile, and Colombia, there are also price crashes in the first quarter of 1994, clearly associated with the same external event.

Finally, we wish to note that most of the crash periods lead or coincide with currency crises documented elsewhere (Kaminsky, Lizondo and Reinhart, 1996) (Appendix C), and this explains why including the stock market price improves the performance of early warning systems of currency crisis (Herrera and Garcia, 1999).

III. DETERMINANTS OF BUBBLES

This chapter seeks to identify the main determinants of the probability of a bubble occurrence, in the vein of the international business cycle theory that allows for domestic stock returns to be affected by foreign economic activity and foreign financial variables, in addition to domestic factors (Canova-De Nicolo, 1995). Initially, we describe graphically some empirical features of the data to elucidate the statistical results presented in the latter part of the chapter.

A. Stock prices, capital flows, domestic credit growth, and terms of trade: Stylized facts and motivating evidence

Graphic analysis (Graphs 11 to 16) shows the evolution of stock prices (in real terms), capital flows, and domestic credit in the biggest LAC countries. Stock prices show a non-stationary behavior, with a significant change in level in the early 1990s. To have a better idea of the timing of these changes, we used Zivot and Andrews (1992) and Vogelsang and Perron (1998) tests for structural breaks. Additionally, we tried Self-Exciting Threshold Autoregression (SETAR) methods to determine approximate break-points in the series. Different methods yield different results (Table 6), but in Argentina, Chile, Colombia, and Mexico, we determine a structural break in the stock price series in mid-1991.

Capital flows are mean-reverting processes with a change in mean. This change took place in 1991 in Argentina, Chile, and Mexico. Domestic credit growth is also stationary, with structural breaks taking place in the early 1990s, except in Mexico, where it happened in 1988.

There are two other interesting empirical regularities. First, the relationship between terms-of-trade positive shocks and the occurrence of bubbles (Graph 17) seems to have been important in Argentina, Chile, Brazil, and Colombia. Second, the realization of the relationship between domestic credit growth and the bubble episodes (Graph 18) seems to take place with a long lag. Almost all of the bubble episodes were preceded by a cycle of credit boom, but in Mexico the bubble surges once the acceleration has ceased.

B. Determinants of bubbles in Latin America

We wish to examine whether there is some statistical support for the empirical regularities described graphically in the previous section. With statistical analysis, we can gauge the relative importance of the different factors affecting the divergence of asset prices from their fundamentals-determined prices.

Our empirical exercise follows the thrust of the international business cycle theory that analyzes the relationship between (a) stock returns in one country and (b) foreign indicators of economic activity and foreign financial variables (Canova-De Nicolo, 1995). Empirically, this approach has been used to explain asset returns in emerging markets (Harvey, 1995 and Hargis-Maloney, 1997), based on information sets that contain both local and global variables.

Among the country-specific factors, we included domestic credit expansion, volatility of credit growth, and volatility of asset returns. One could justify the inclusion of these variables based on the Allen-Gale bubbles model. Other country-specific variables were capital flows and terms-of-trade shocks,³⁴ though they are “external.” The common external factors were the degree of overvaluation in asset prices (stocks and real estate)³⁵ in the U.S.A. and the spread between 10-year bonds and 3-month Treasury bills.³⁶ This last variable is included to incorporate agents’ expectations of future real interest rates in the U.S., or of the future course of economic activity (Estrella and Mishkin, 1997; Fama, 1990). If future output is expected to be high, consumption-smoothing agents will borrow against it and bid up interest rates (Chen, 1991). Alternatively, if the term spread is a good predictor of inflation, a rise in the spread anticipates future falls in real interest rates, as has been shown (Fama, 1990). Both effects lead to a positive expected sign in the coefficient of the term spread.

³⁴ Terms of trade were introduced in deviations from a Hodrick-Prescott trend.

³⁵ As a proxy of stock overvaluation in the U.S., we took the residuals from the intrinsic bubble model. As a proxy of real estate overvaluation, we considered deviations from a Hodrick-Prescott trend of the returns of equity real estate investment trusts (REIT).

³⁶ The spread between 10-year bonds and 3-month bills contains information about future changes in interest rates (Campbell and Shiller, 1987 and Mishkin, 1988) or future changes in economic activity (Estrella and Hardouvelis, 1991 and Chen, 1991). On the use of stock market prices and a proxy for real estate simultaneously, several studies have found a negative correlation between stock returns and real estate returns in the U.S. (Ibbotson and Siegel, 1984; Hartzell, 1986; and Worzala and Vandell, 1993). For this reason, some analysts recommend real estate investment as a hedging instrument for stock investments (NAREIT, 2000). We did not include real interest rates as an explanatory variable, since they are used in the construction of the bubble indicator to estimate the present-value price (to calibrate the value of K).

Pooling monthly data for five countries since 1985, we estimated a Logit model³⁷ (Table 7) with satisfactory results.³⁸ All of the domestic variables were significant and had the expected sign. Volatility of domestic credit and volatility of asset returns affect positively the probability of bubble occurrence, as predicted by the Allen-Gale model. So do terms-of-trade (positive) shocks and capital flows. Domestic credit growth has a negative sign 12 months prior to the bubble, but the sign is positive 18 months prior to the bubble, reflecting the non-linear effect of this variable.

Regarding the external variables, a striking result was the opposite sign of the overvaluation proxies for stock markets and real estate. When stock markets in the U.S. are overvalued, the probability of a bubble occurrence in Latin American stock markets declines. However, the probability is positively associated with overvaluation in real estate markets. The only way to reconcile this result is through the negative or low positive correlation between these asset classes in the U.S. The term structure spread is significant and has the expected sign: higher-term structure spreads are leading indicators of expanding economic activity or lower future real interest rates (Estrella and Mishkin, 1997 and Fama, 1990)

The marginal probabilities (Table 8) that tell us how the probability of a bubble changes with a unit change in each individual variable were calculated by estimating the marginal effect at each observation and then taking the sample average. The differences across countries arise because of the fixed-effects coefficient, and because the values of the explanatory variables are different for each country. The term spread in the U.S. has the largest marginal impact on bubble probability. This statistical result is confirmed by observing the correlation between the term spread and the residuals of the intrinsic bubble model (Graph 19), which is surprisingly close in the Mexican case. The second most influential variable is domestic credit growth; the volatility of real asset returns and terms-of-trade shocks are of similar importance.

Even though our results are not directly comparable to Harvey's (1995) or to Hargis and Maloney's (1997), given that they examine the impact of a set of variables on returns and we examine the impact of a set of variables on the probability of bubble occurrence, the general outcome reflecting the extent of the influence of global factors seems quite different. Part of the discrepancy can be explained by the sample period, since both papers examine data from 1976 to 1993, missing a great part of the economic developments of the 1990s which seem to dominate our results.

³⁷ We estimate both Logit and Probit models, and the results are almost identical (see Appendix D). However, we present the Logit exercise in the main text of this section, given the possibility that in some cases Probit models do not produce consistent estimators with fixed-effects panel methods (Hsiao, 1997 and Green, 1999).

³⁸ The model was estimated with instrumental variables to control for potential endogeneity of the domestic variables. Additionally, to avoid difficulties in the interpretation of the marginal probabilities, as well as to facilitate comparison across variables, we employed standardized variables (subtracting the mean and dividing by the standard deviation of the series) to circumvent the units problem. This does not alter the signs or significance of the results obtained with the raw data.

IV. ASSET PRICES, ECONOMIC ACTIVITY, AND GROWTH

A. Stock prices and economic activity: Consumption and investment

In LAC countries, we found a clear direction of causality running from stock market returns to consumption and investment (Table 9). The causality between stock returns and consumption can be justified on several grounds. First, there exists the wealth effect, according to which agents' consumption decisions depend on wealth, which in turn is affected by stock price movements. Given that stocks are not a widespread form of saving in LAC, this effect should not be important in explaining the observed statistical result. Second, stock prices might be signaling the future growth of the economy and, hence, the future growth of labor income (Ward, 1999), which will affect consumption. Third, as the perceived wealth of individuals changes positively, so does their creditworthiness, allowing agents to increase their indebtedness and/or to receive credit at a lower cost.

The observed causal relationship between investment and stock prices can be explained by Tobin's q theory: as stock prices change, then the ratio between the market valuation of existing capital and the cost of new capital (Tobin's q) changes. As stock prices rise, the cost of new capital falls relative to that of existing capital, and firms will acquire new capital (invest). A second explanation for the observed relationship between stock prices and investment can be found along the lines of the second explanation for consumption: stock prices convey information regarding the future growth of the economy and, therefore, affect investment through this channel. Finally, the balance sheet effect may also operate here: as asset prices change, so does firms' perceived creditworthiness and hence their access to credit.

Examining quarterly stock price and investment data for the 1990s (Graph 20), a clear positive relationship emerges. To incorporate stock prices into a longer-run perspective on the determinants of investment, we estimate a panel for five Latin American countries from 1980 to 1999 with SUR methods (Table 10). The dependent variable is the ratio of gross domestic investment to GDP while the explanatory variables were: foreign direct investment (Borensztein, De Gregorio, Lee, 1998), a dummy variable reflecting situations of state failure (from the Polity International database), government consumption expenditure, government capital expenditure, the log of stock prices, the volatility of stock returns, and lagged growth (Seven and Solimano, 1993 and Cardoso, 1993).³⁹ All

³⁹ The panel was estimated by Seemingly Unrelated Regressions (SUR) with instrumental variables. Appendix D shows that estimation with the original data does not alter the sign or significance of any of the results, and changes in the parameter values are relatively minor. The variables are all cointegrated according to tests proposed by Kao (1999) and Pedroni (1999), attenuating the autocorrelation of residuals problem. These tests require balanced panels; consequently, for this purpose only, we discarded Colombia from the sample, since it had stock price data only since 1984. (None of the results described in the text changes when Colombia is excluded from the sample.) The dynamics of investment should be explored by more elaborate methods, but the number of observations poses a problem. Notwithstanding this problem, we estimated the panel with variables expressed in changes and levels; this approach deals with serial correlation (Appendix D). The only significant variables turn out to be stock prices, the volatility of stock returns, the state failure variable, and government capital expenditure. With a larger set of countries and a longer time period, it can be shown (Herrera and Garcia, 2000) that government consumption and foreign

of the variables had the expected sign. We wish to highlight the positive sign of the stock prices and negative sign of the volatility variable. Hence, although bubbles may have a positive effect through increases in stock markets, prices generally end up lower than before the bubble started. Other negative effects on investment are transmitted by the increased volatility effect. Additionally, the adverse effect of bubbles on the financial sector lasts for many years, negatively affecting growth, given the relationship between financial intermediation and growth (Beck, Levine and Loayza, 2000).

A bubble episode may lead to overinvestment, which, combined with the effect on consumption, may lead to domestic undersaving. These two factors combined should produce larger current account deficits as asset prices tend to be overvalued. Graphical inspection (Appendix E) shows this is likely to be true in the United States and Mexico. However, other Latin American countries do not show any discernible relationship, and further work is required.

B. Housing prices and domestic credit

In the explanations of the observed relationship between asset prices and consumption and investment, the credit channel played a significant role. Here, we explore in more detail how asset price changes affect economic activity through the balance sheets of agents that change their creditworthiness as asset prices change. There exists a close relationship between housing prices and domestic credit growth in the three countries where we could get more complete series, namely Argentina, Colombia, and Mexico (Graph 21), and in Chile in the late 1970s and early 1980s before the banking crisis.⁴⁰ Granger causality between asset prices and domestic credit goes in both directions; however, when restricting the sample to the 1990s only, the causality is stronger (larger F statistics) from housing prices to domestic credit (Table F1 of Appendix F). Regarding the relationship between stock prices and domestic credit, we found evidence of Granger causality from the stock returns to domestic credit growth (Table F2 of Appendix F).⁴¹

These findings can be interpreted as suggestive evidence of balance-sheet effects operating as described by Bernanke, Gertler, and Gilchrist (1998). In their model, an agent's external finance is more expensive than internal financing, and the premium for external finance varies inversely with the net worth of potential borrowers. So, as asset prices fall (in this case, real estate prices) the net worth of borrowers falls, and banks are less willing to lend.

direct investment affect the investment ratio, with the signs and significance indicated in this descriptive exercise.

⁴⁰ Conley and Maloney (1995) graph date series for housing prices and consumer loans, but do not link their behavior relative to each other; they also do not link these data to the banking crisis that followed. They see the behavior of these series as separate elements of the euphoria prevailing in Chile during the period.

⁴¹ The causality direction changes when other transformations are used (i.e., month-to-month growth rates instead of year-to-year).

When asset prices rise, banks tend to lend more to the sectors affected by the bubble. Mexico's case is revealing: as real estate prices fell in the early 1980s, lending to the construction and housing sectors fell as a percentage of total bank lending. When real estate prices rose, the proportion of bank lending to these sectors doubled, reaching a peak in 1995, when real estate prices stopped growing and reversed the trend (Table 11). In Colombia, a similar relationship between real estate prices and credit to housing and construction is observed.

When the crash occurs, non-performing loans rise as borrowers become insolvent, and the ratio of non-performing loans is greater in the housing sector (Table 12). In Colombia, non-performing loans jumped from 6% at the end of 1997 to 10% in 1999 and to 13% at the end of 2000.

Argentina did not have a banking crisis of the magnitude experienced by Mexico or Colombia.⁴² In this country, credit to the construction sector did not increase during the rise of real estate prices. On the contrary, it fell (Table 11) and a similar trend occurred in housing mortgage credit. In the first quarter of 1991, credit to the construction sector was 6.7% of total credit; by the second quarter of 1994 it had fallen to 4.0%. Once the Tequila crisis hit in late 1994, the non-performing loans of this sector increased from 20% of total loans to the sector to 24%, which is much smaller than the Mexican case and slightly larger than in the Colombian case.

Another difference in banking practices between (a) Argentina on the one hand and (b) Colombia and Mexico on the other is revealed by the non-performing loans coverage ratio.⁴³ While Argentina's banking sector coverage was 55% in December 1994, it quickly rose to 68% in 1996. In Mexico, the coverage ratio was 49% in 1994, and it had risen to 55% by 1997. When the crisis began in 1998, Colombia had an extremely low 38% coverage ratio, which then increased to 46% by the end of 2000.

Finally, the Chilean experience, as derived from the Conley-Maloney data on housing prices, shows that the bubble was accompanied by rising domestic credit. In this case, the credit explosion took place in 1976–1979, prior to the rise in property prices.⁴⁴ However, the credit contraction did follow the fall in real estate prices, and the banking crisis in late 1981 followed the collapse in real estate prices. This is another example of the balance sheet effect operating. This bubble and the credit expansion occurred immediately before the banking crisis of the early 1980s.

⁴² World Bank staff estimates that the fiscal cost of the banking sector crisis in Mexico amounted to about 19% of GDP, and IMF staff estimates Colombia's banking sector rescue costs to be around 8% of GDP, while there were no fiscal costs in Argentina.

⁴³ The ratio of provisions for loan losses to non-performing loans.

⁴⁴ We were unable to find a uniform data series of credit for the non-financial private sector, and, therefore, we present only non-manipulated series. However, all the available series show the credit explosion in 1975–1979.

C. Bubbles have long-lasting consequences on poverty issues

Evidence from Mexico and Colombia (Appendix G) shows that, after the bubble bursts and domestic credit contracts, the proportion of people who own the house in which they live falls. This phenomenon tends to be more acute in the poorest quintiles. In 1995 in Colombia, 68% of households fully owned the houses in which they lived; by 1999 (after the crisis), this figure fell to 58%. In Mexico, the same statistic fell from 72% to 66% between 1995 and 2000. If middle-class and poor people lose their main asset during crises, their credit availability will be impaired in the future, given that they lose their only source of collateral. The fact that this phenomenon happens more often in middle- and lower-income families will affect income distribution inequality and, through this channel, will affect long-run growth (Barro, 2000)

V. POLICY IMPLICATIONS AND CONCLUSIONS

After documenting the existence of bubbles in Latin America, we examined their determinants and provided evidence of the negative effect of asset price bubbles on the financial sector, on investment, on income distribution, and, thence, indirectly through all these channels, on long-run growth. Therefore, the case for intervention to avoid bubbles seems warranted. However, there are several problems, beginning with the identification of bubbles, or asset price misalignments. The main difficulty arises from the observational equivalence of a bubble with the expected future changes in fundamentals. Circumventing this discussion, we turn to examine what authorities can do, based on our findings concerning bubble determinants.

Regarding common external (global) factors, there is little that policymakers in Latin America can do to influence their course. Similarly, there is little that can be done about country-specific, but external factors which we considered, such as capital flows and terms-of-trade shocks. There is ample literature about insulating the economy from capital flows volatility, and evidence is mixed regarding its effectiveness and costs (Dooley, 1996; Montiel, 1999). Regarding terms-of-trade shocks, our results show that there is still room for improvement on stabilization funds and saving mechanisms that buffer these recurrent shocks.

On the country-specific domestic factors the story is different. Given the difficulty of identifying the imminence of bubbles, authorities should permanently aim at smoothing the cyclical behavior of credit for the private sector. To this end, policymakers could use several tools. The first one would be the adoption of counter-cyclical provisioning of credit such that, during expansion phases, financial intermediaries create capital cushions, beyond the specific credit loss risks, to absorb potential losses during the crash. This generic provision would apply to all loans. Other types of restrictive policies could restrain the use of stocks and real estate as collateral during booms. Alternatively, by placing higher capital requirements for highly-leveraged customers, authorities can limit the potential for bubble generation and posterior damage. Finally, a widely used instrument in the past was the imposition of credit growth ceilings. The problem with

these restrictive policies, with the exception of the generic counter-cyclical provision, lies in the difficulty of monitoring, and, hence, they are subject to regulation arbitrage.

Another policy suggested to minimize the negative impact of the bursting of the bubble is based on the fact that most developing countries have inefficient secondary markets for family housing units. The functioning of this imperfect market is impaired even more during crisis, when negative shocks hit the economy, and banks are stuck with repossessed collateral, which they pretend to dump in the market to obtain liquidity. Because of this reason, analysts (Holmstrom and Tirole, 1998) have proposed some type of restraints on the disposition of real estate by commercial banks to prevent prices from falling further during stress situations. In this setting, both borrowers and banks are insured against price collapse, and consumers pay higher prices for real estate, but the insurance raises an ex-ante social surplus.

Rather than adopting policies that “throw sand into the wheels,” others propose the expansion of the number and variety of markets to stabilize prices. For example, Shiller (1993) argues for the creation of what he labeled macro markets. These are international markets for long-term claims on the incomes of countries or different occupational groups or markets for highly illiquid assets such as single-family homes. These are complex securities that are just being developed and are of very limited use. As an example of one of the few cases where these concepts have been used, Bulgaria recently issued GDP warrants as part of the Brady renegotiation.

Along this line, an option to develop real estate markets and diversify risk would be to use Real Estate Investment Trusts (REITs) which are passive portfolio managers of real estate properties (Herring and Wachter, 1999). These agents, whose shares are publicly traded, act as mutual funds that hold property and generate income and tax-free capital gains for individual investors.⁴⁵

Once a bubble is identified, attempting to puncture it is very risky. “Talking down” the bubble might be a difficult task, and, as time goes by, authorities will be forced to intervene directly. Recent experience in the U.S. shows that policy intervention is not exempt from costs. Another good example of the costs of attempting bubble-puncturing interventions is the Brazilian episode of the Real depreciation up to July of 2001. The depreciating trend of the Real (Graph 22) took the currency to around 2.50 per dollar by the third week of June 2001. Because this movement was interpreted as being driven by a speculative bubble, the central bank abandoned its no-intervention (in the foreign exchange market) policy, in effect since 1999, and sold dollars to “irrigate” the market, according to the central bank president. Selic interest rates were simultaneously raised by 150 basis points. The market welcomed the “Fraga moves” and the Real appreciated to 2.30 per dollar. Two weeks later, in the first week of July, the Real had fallen back to 2.55 per dollar, and the central bank lost approximately US\$2 billion of reserves that the IMF had disbursed in the last week of June (Graph 22). Additionally, on June 27, the central bank sold a total of \$4 billion Reales (approximately US\$1.7 billion) in one-year,

⁴⁵ The National Association of REITS (NAREIT) website presents detailed descriptions of the types of trusts that exist.

dollar-indexed bonds to stabilize the currency. Summing up the whole episode, the central bank's net debt in foreign currency increased by US\$6 billion, and the exchange rate was at about the same level as before the attempt to puncture the bubble.

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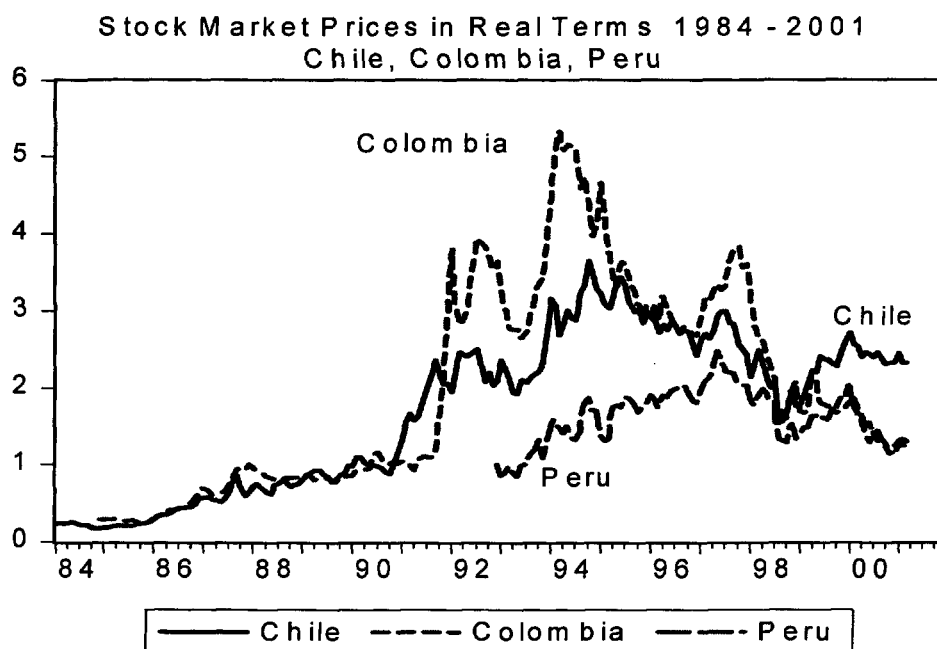
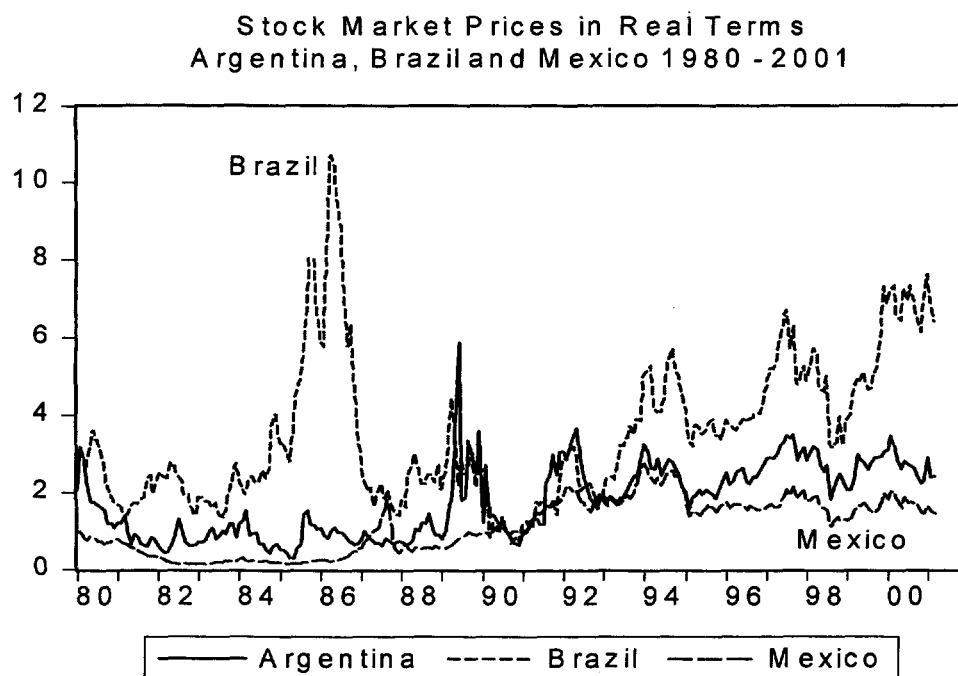
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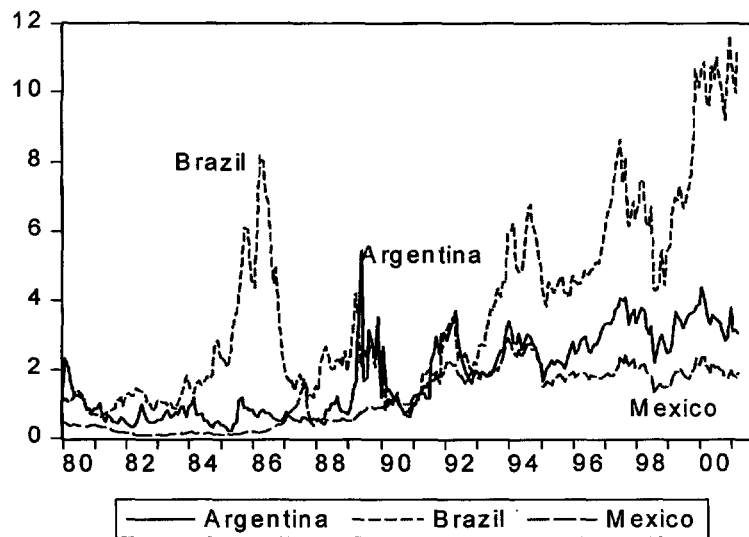
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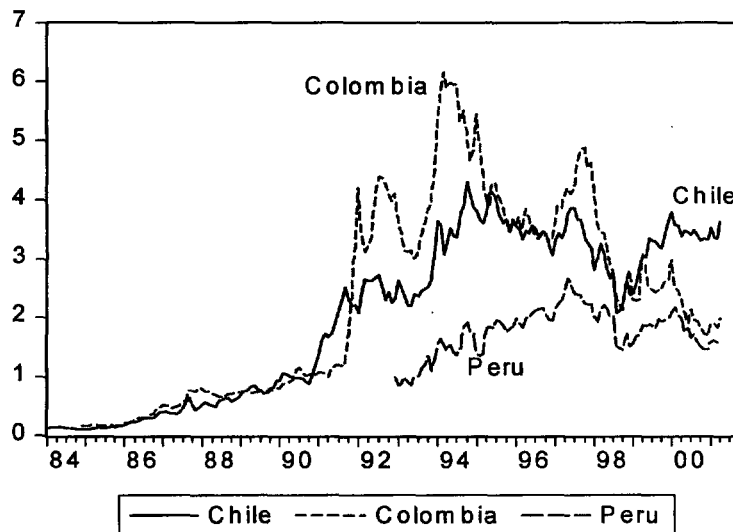
Graph 1

GRAPH 2

Stock Market Returns 1980-2001
Argentina, Brazil and Mexico

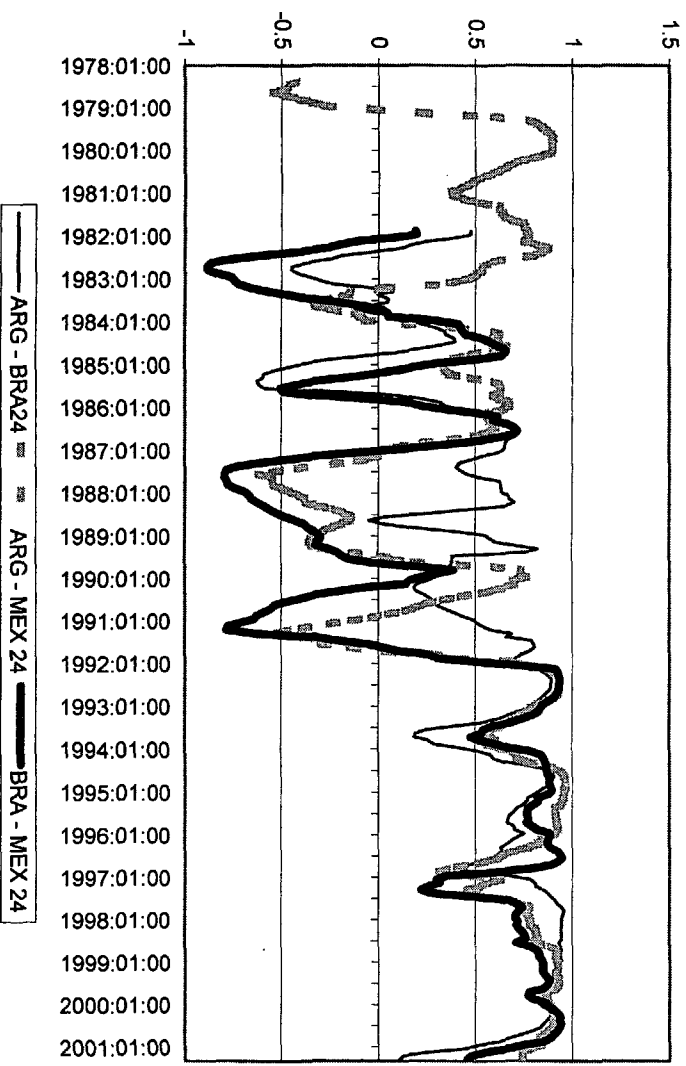


Stock Market Returns 1984 - 2001
Chile, Colombia and Peru

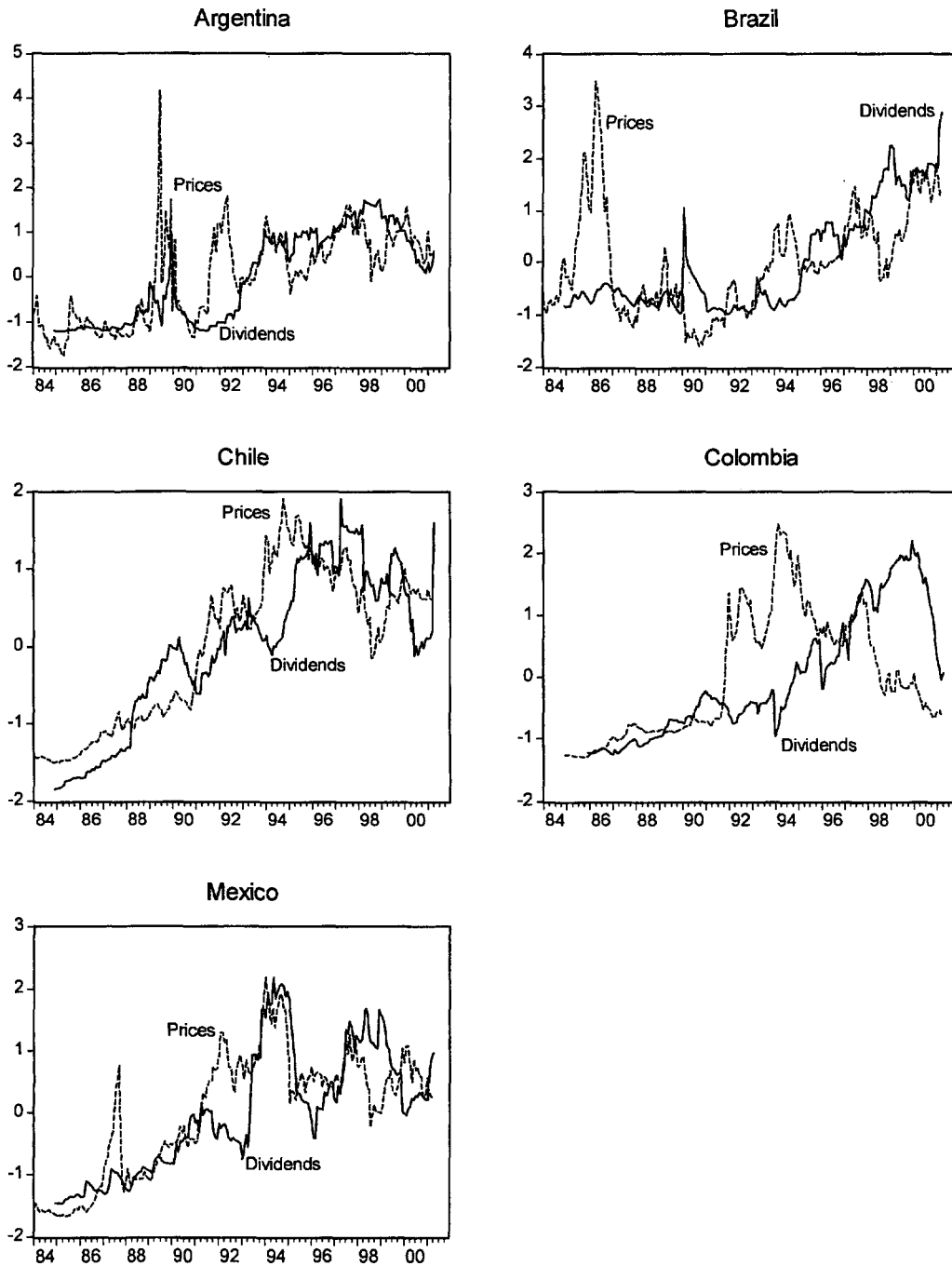


Graph 3

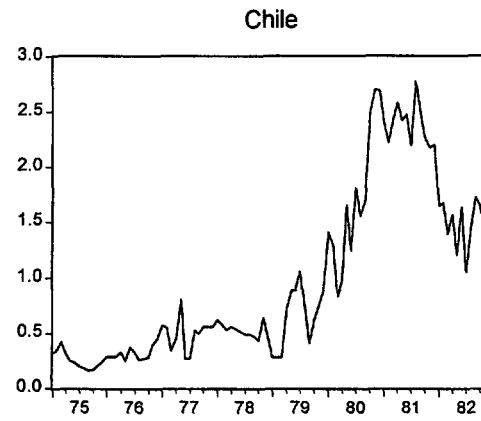
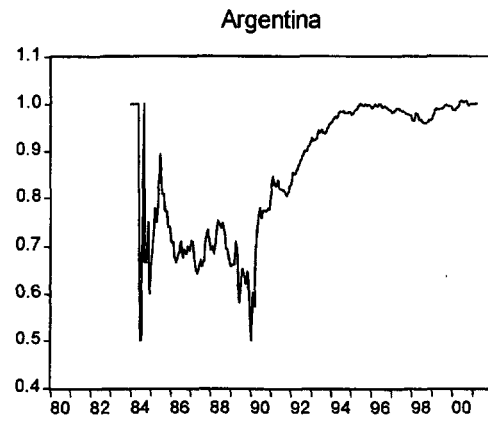
Correlation of Stock Market Returns: 1978-2001
Argentina, Brazil, and Mexico



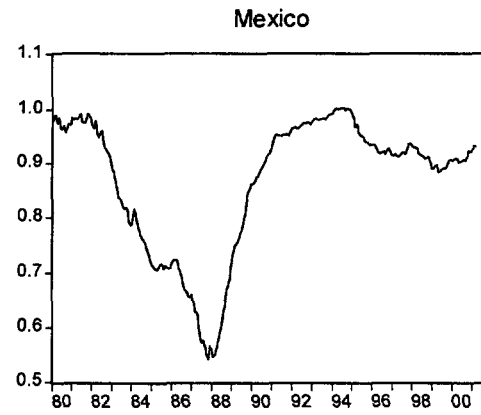
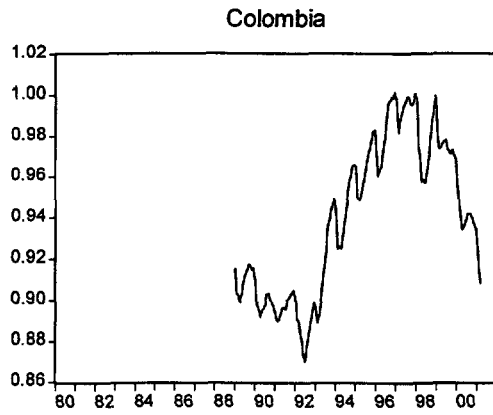
Graph 4
Stock Market Prices and Dividends in Latin America 1984 – 2001
(normalized data; prices dashed line, dividends solid line)



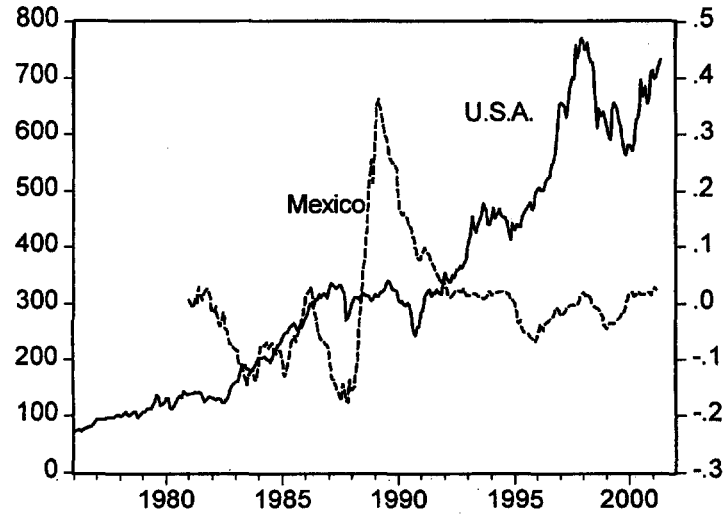
Graph 5
Real State Prices



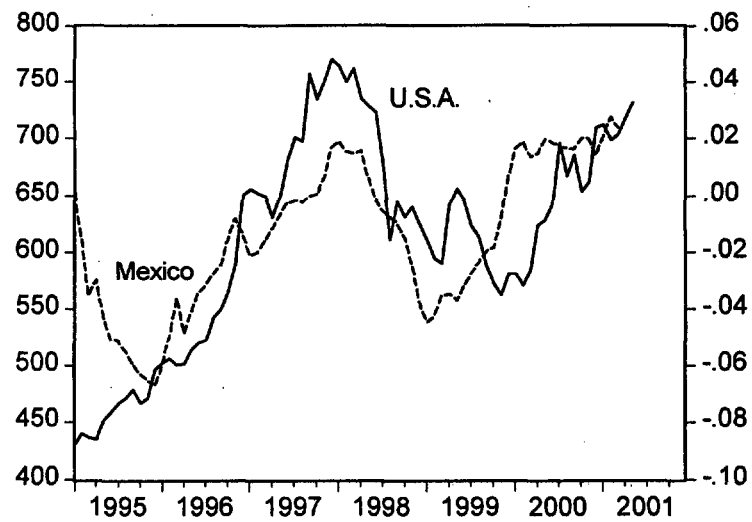
Source: Conley and Maloney (1995)



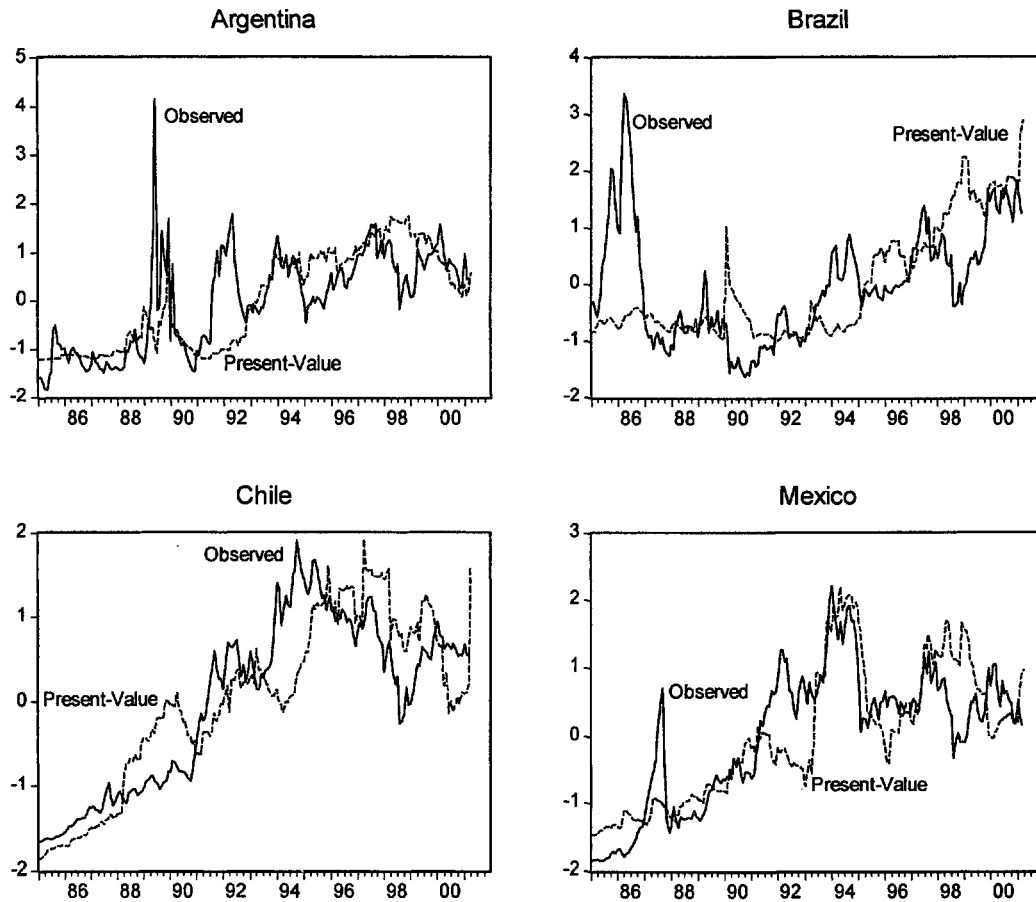
Graph 6a
Returns to Real State Investments in Mexico and U.S.A. 1976-2001



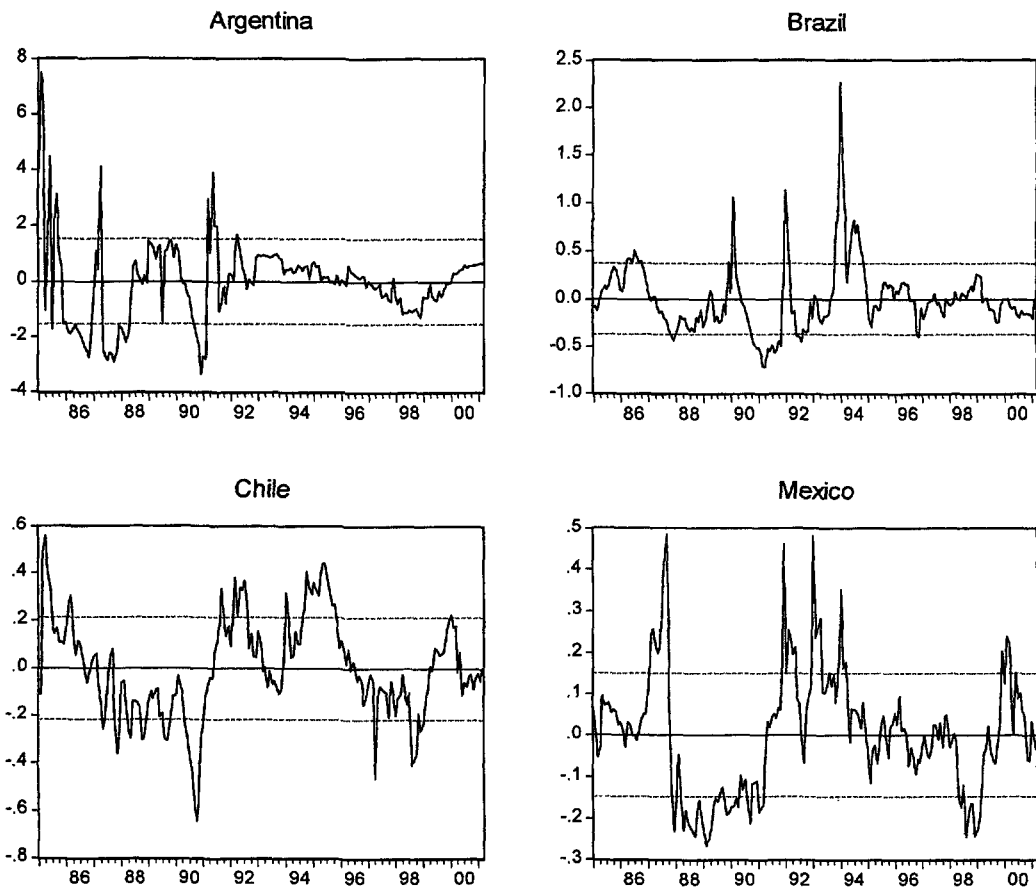
Graph 6b
Returns to Real State Investments in Mexico and U.S.A. 1995-2001



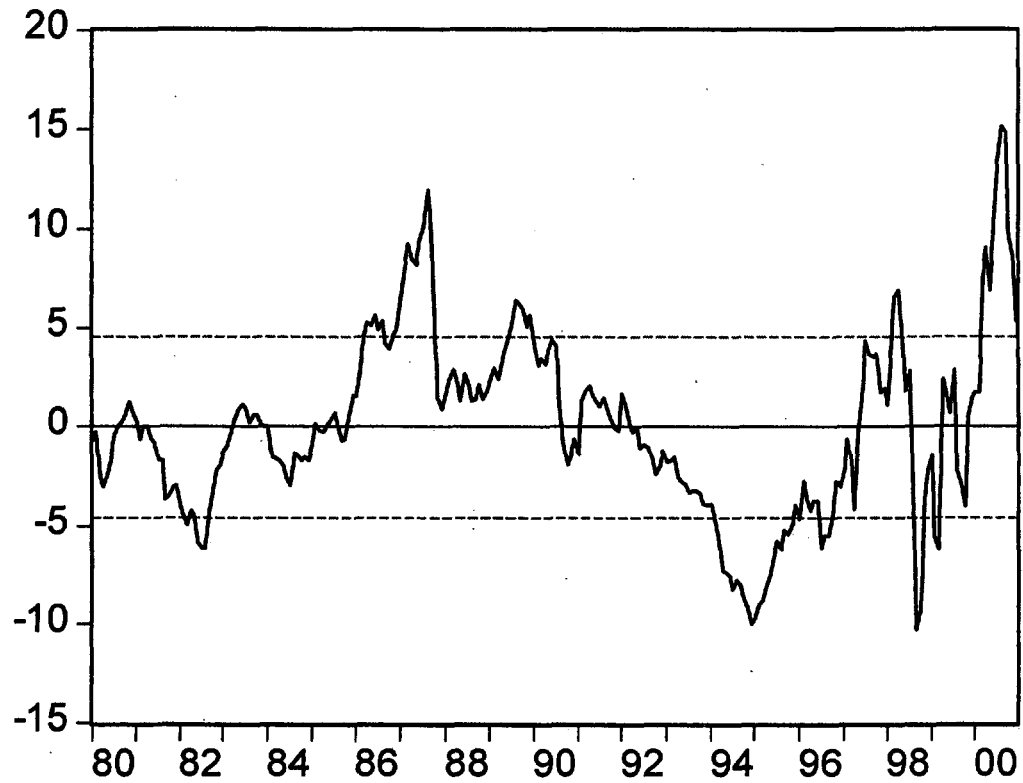
Graph 7
Divergence between Observed Prices and Present-Value Prices
(Solid line, observed prices; dotted line present-value prices)



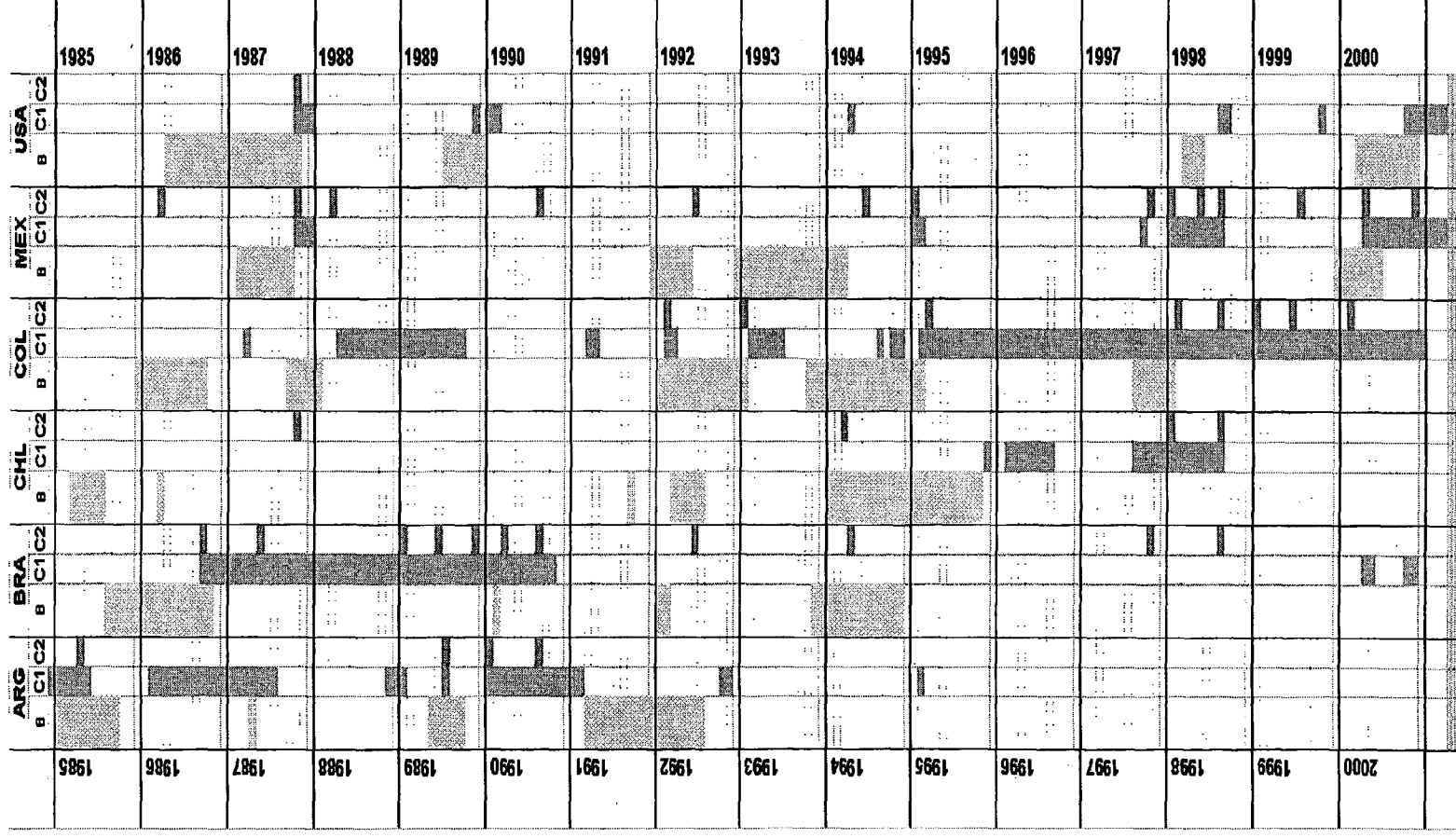
Graph 8
Deviations of the Observed Price from the Intrinsic Bubble Model Predicted Price
in several LAC countries



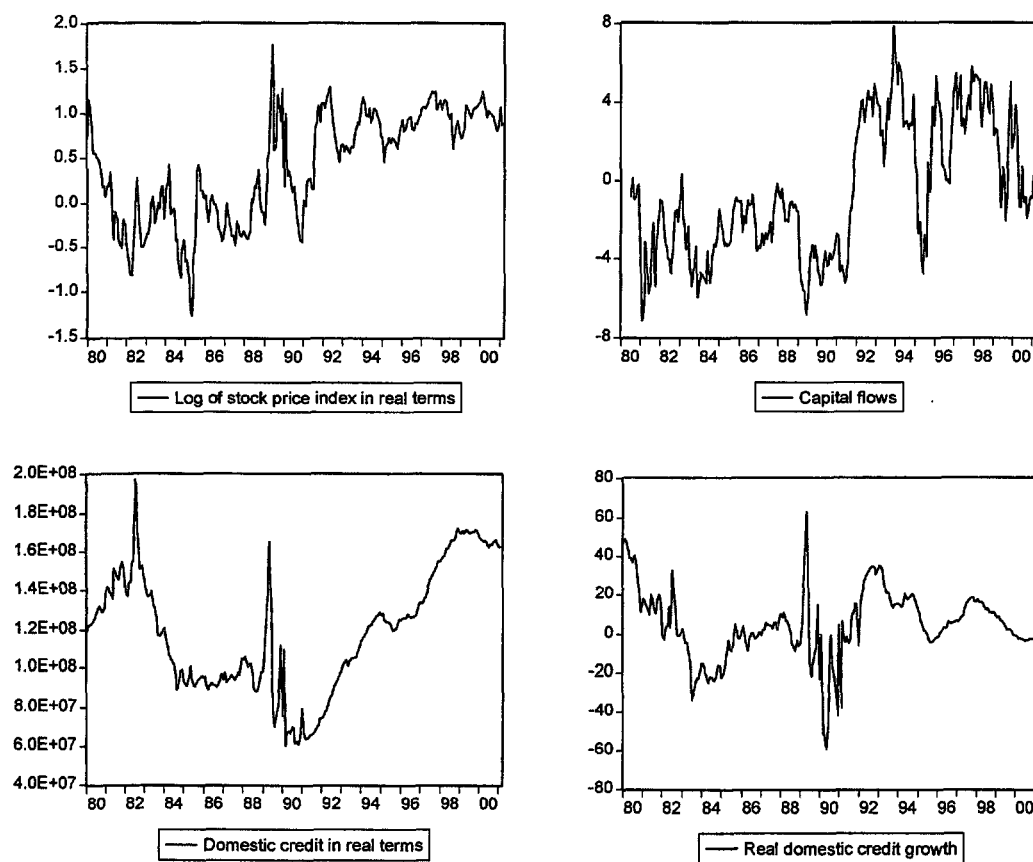
Graph 9
Deviations of the Observed Price from the Intrinsic Bubble Model Predicted Price
in the U.S.



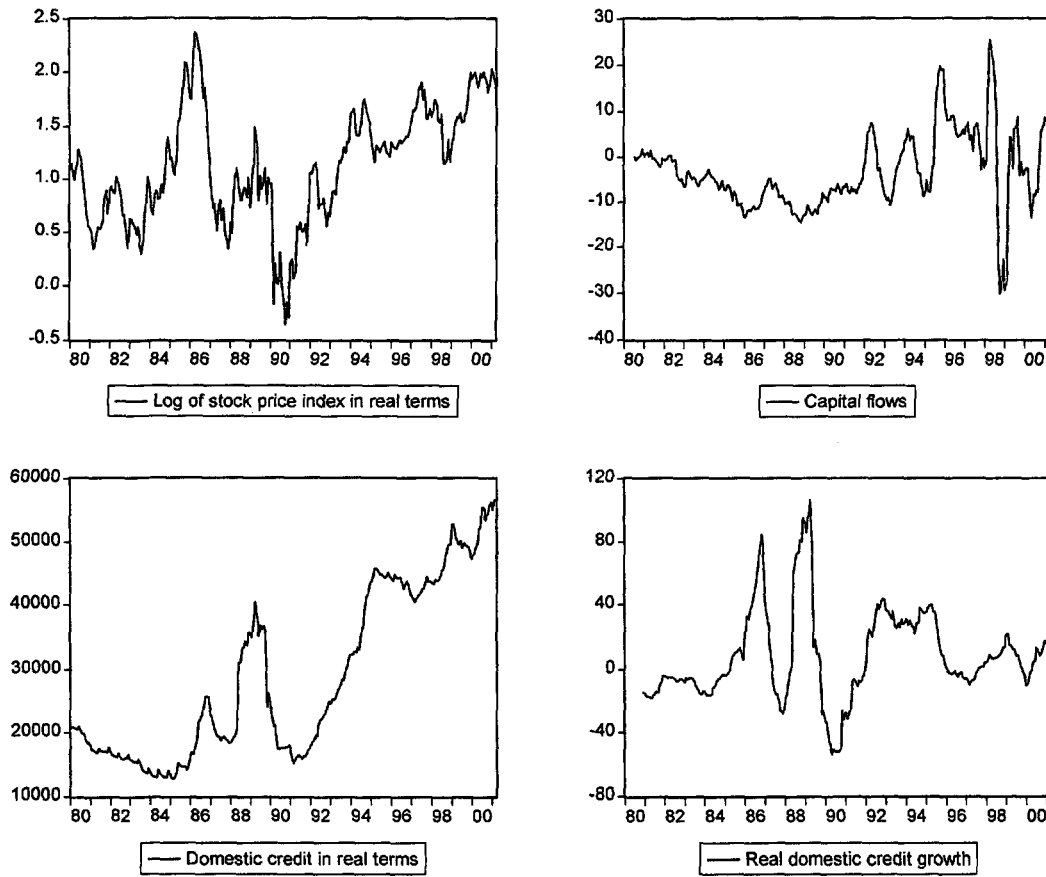
Graph 10
Timing of Bubbles (B) and Crashes (C1 and C2)



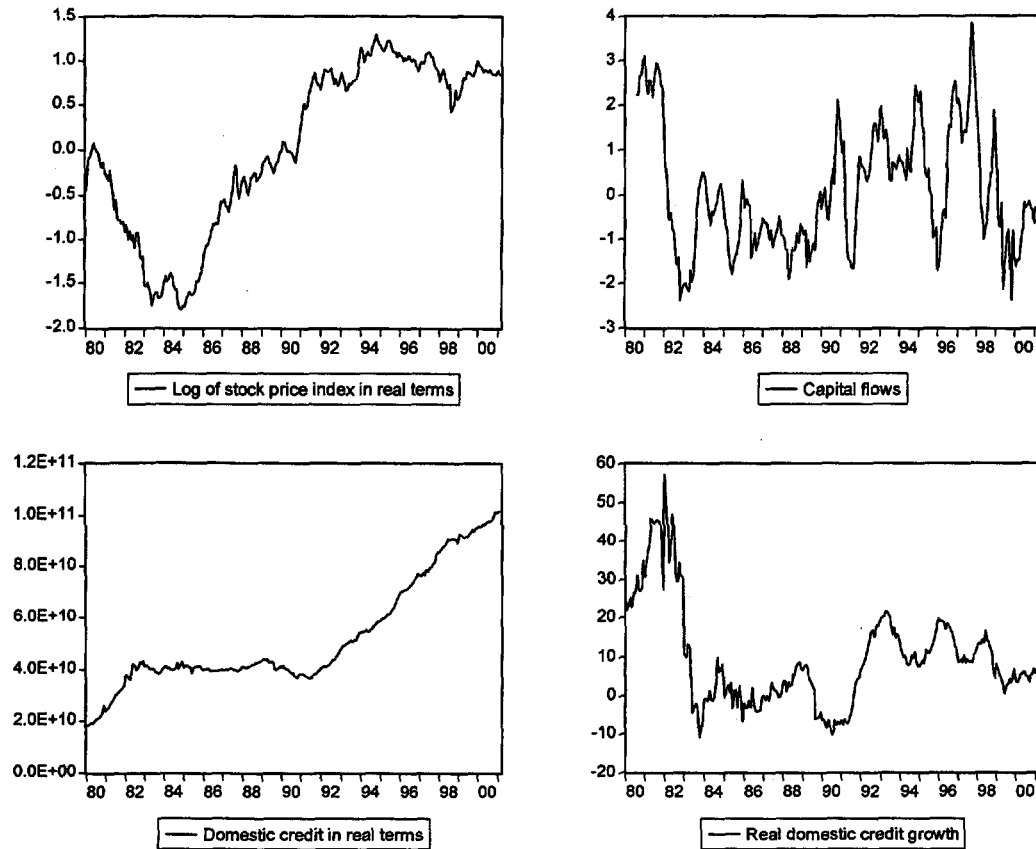
Graph 11
Stock Prices, Capital Flows and Domestic Credit. 1980 -2001
Argentina



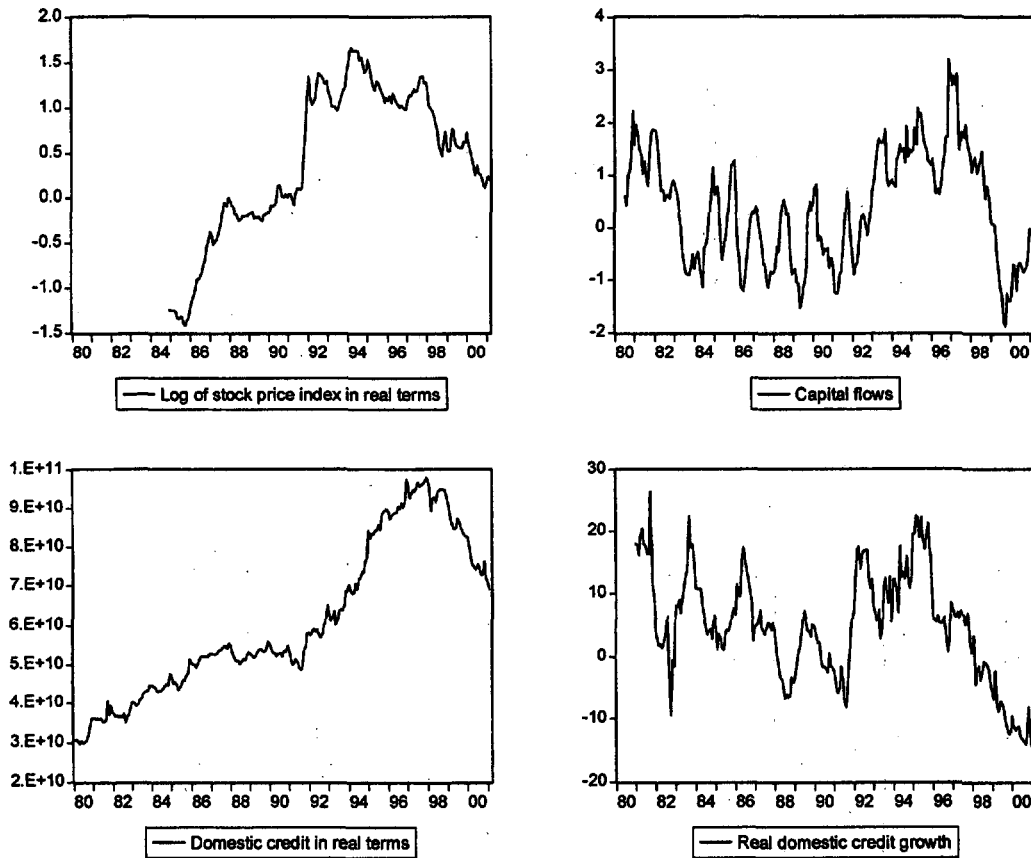
Graph 12
Stock Prices, Capital Flows and Domestic Credit. 1980 -2001
Brazil



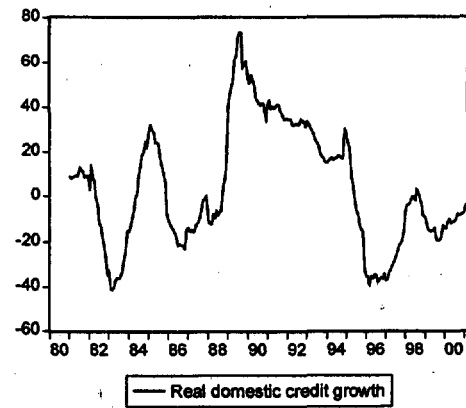
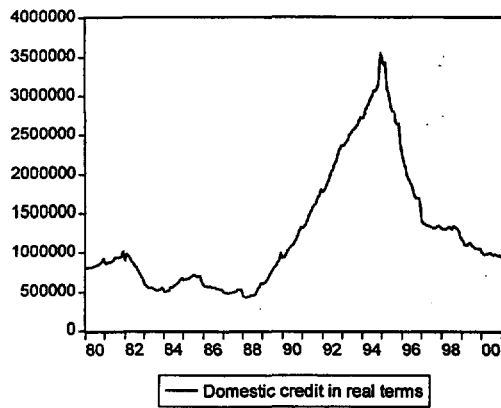
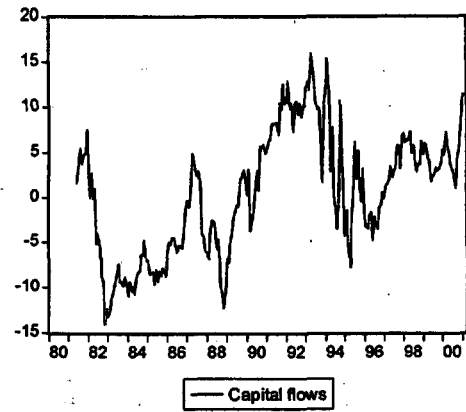
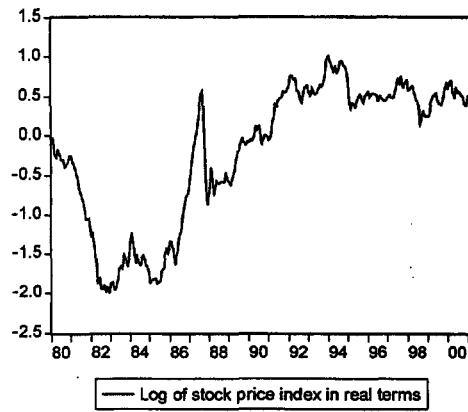
Graph 13
Stock Prices, Capital Flows and Domestic Credit. 1980 -2001
CHILE



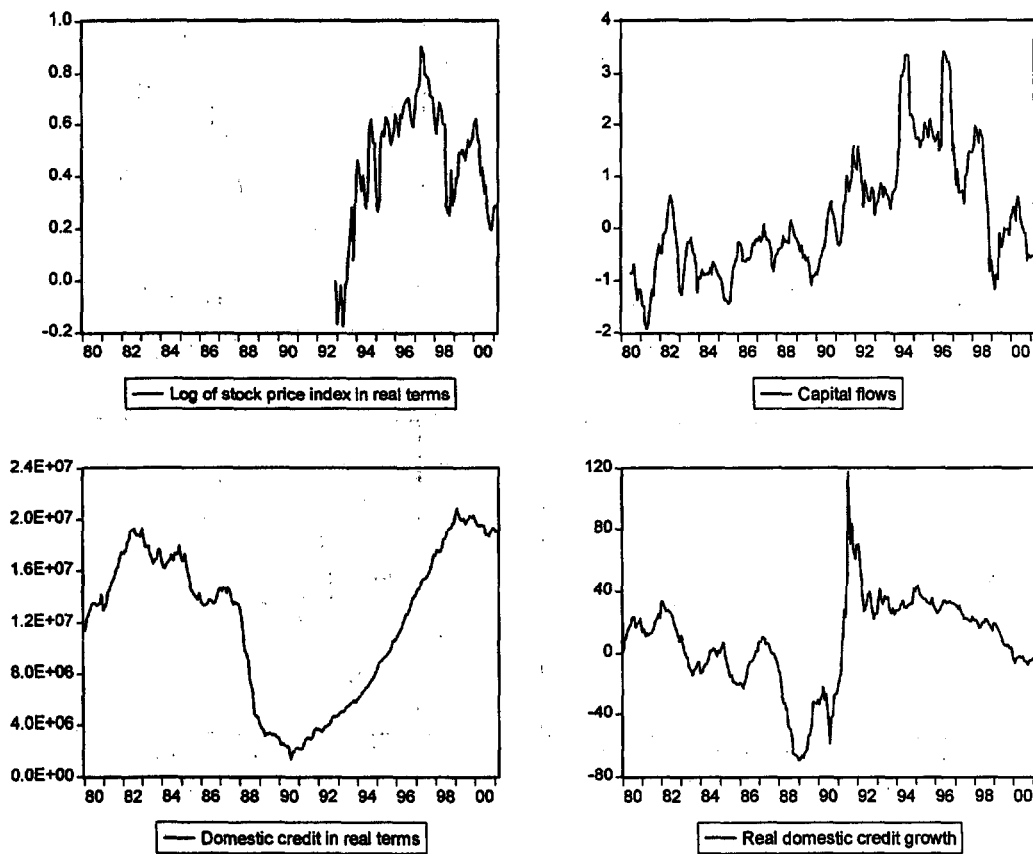
Graph 14
Stock Prices, Capital Flows and Domestic Credit. 1980 -2001
COLOMBIA



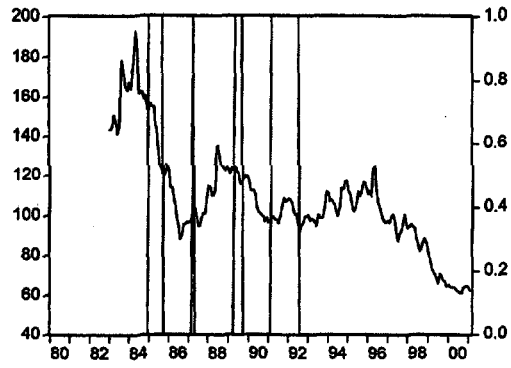
Graph 15
Stock Prices, Capital Flows and Domestic Credit, 1980 -2001
Mexico



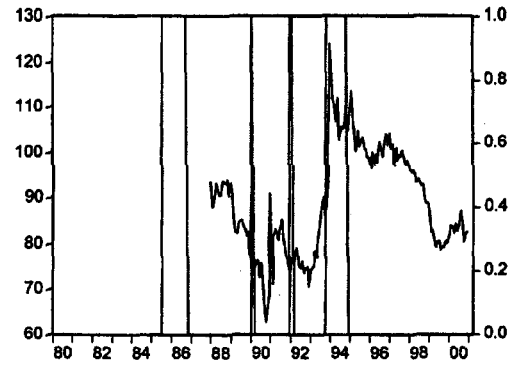
Graph 16
Stock Prices, Capital Flows and Domestic Credit 1980 - 2001
PERU



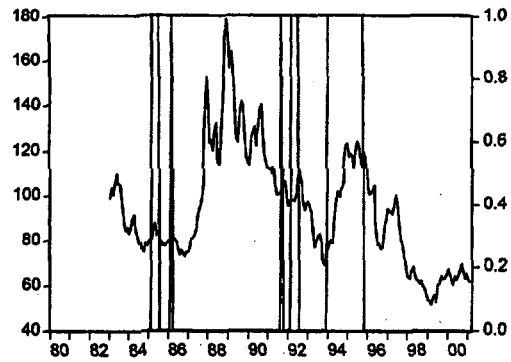
Graph 17
Positive Terms of Trade Shocks and Bubbles



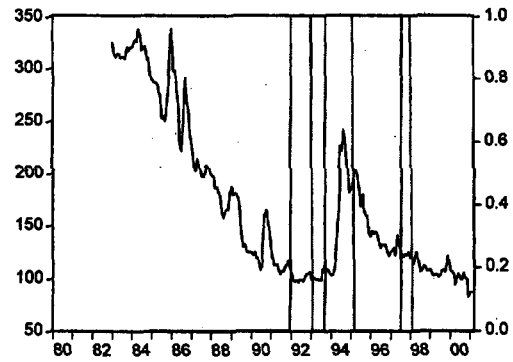
Argentina



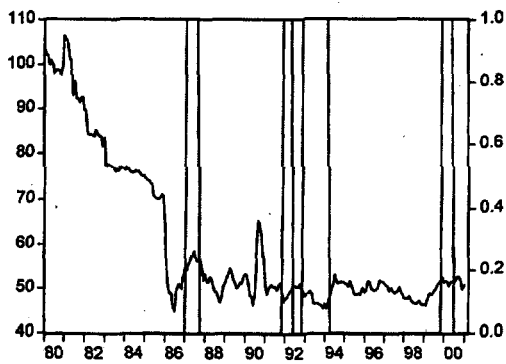
Brazil



Chile

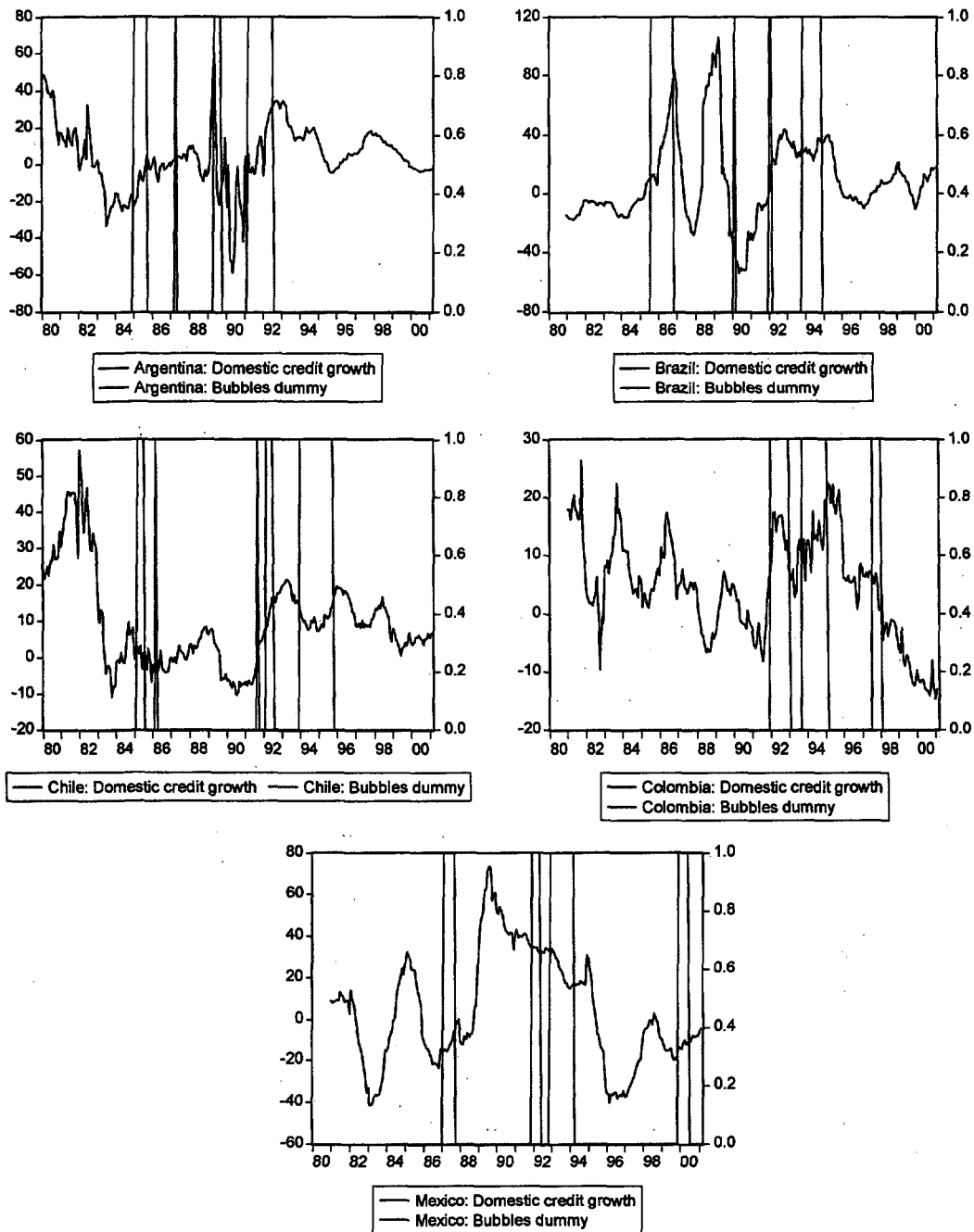


Colombia

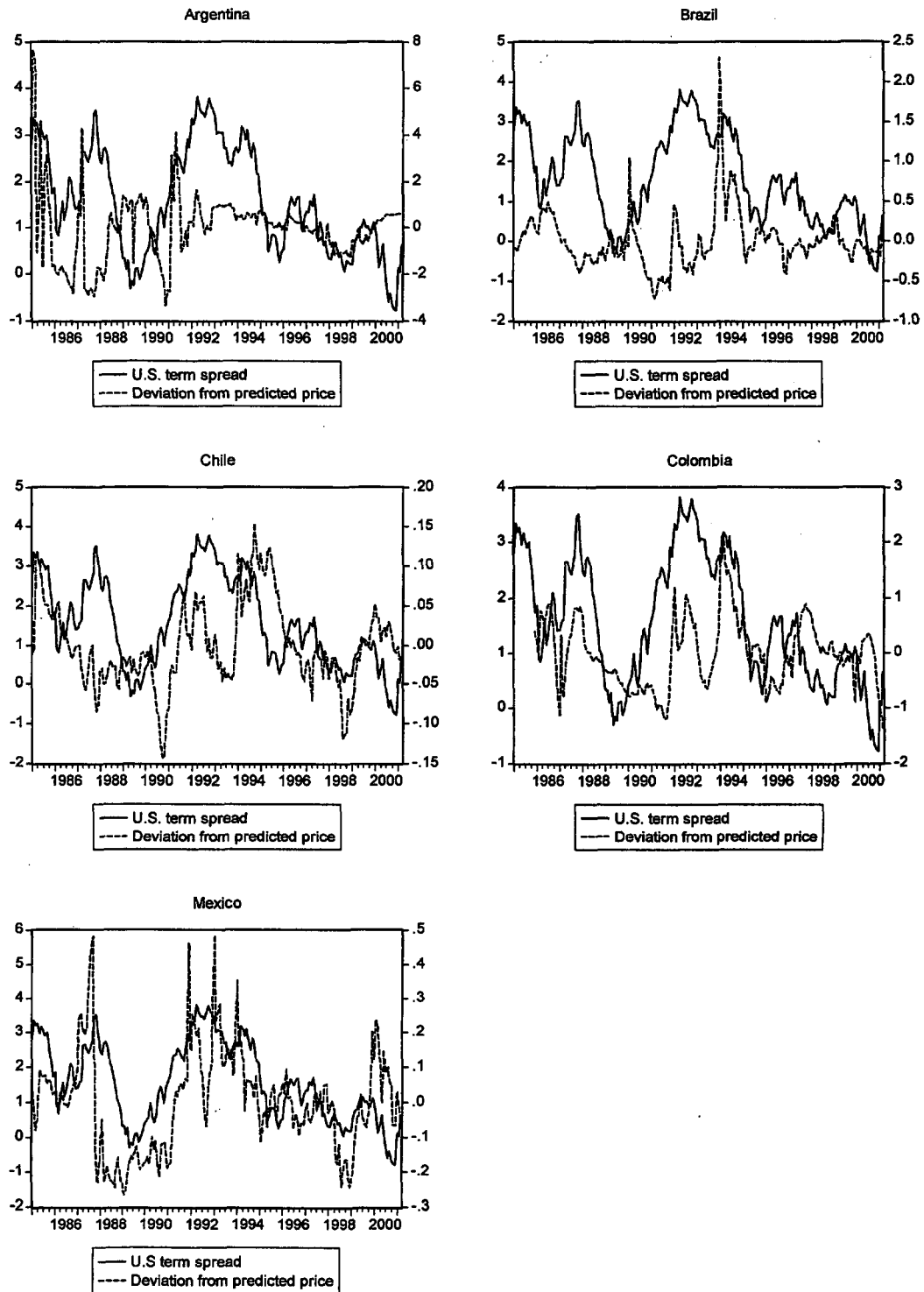


Mexico

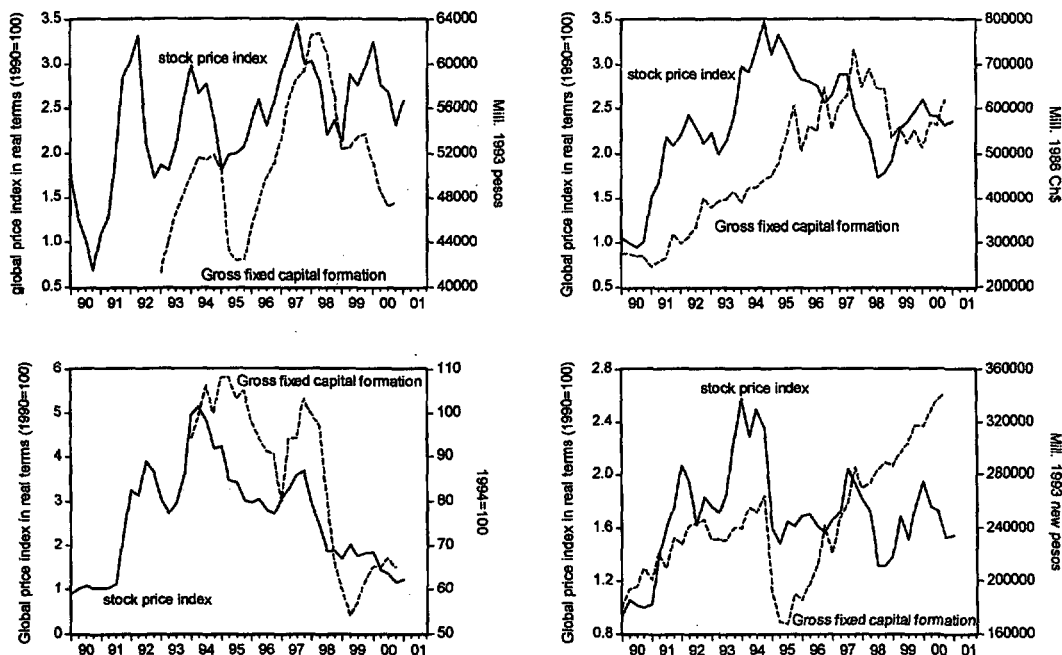
Graph 18
Real Domestic Credit Growth and Bubbles



Graph 19
U.S. TERM SPREAD AND STOCK PRICE OVERVALUATION IN LAC

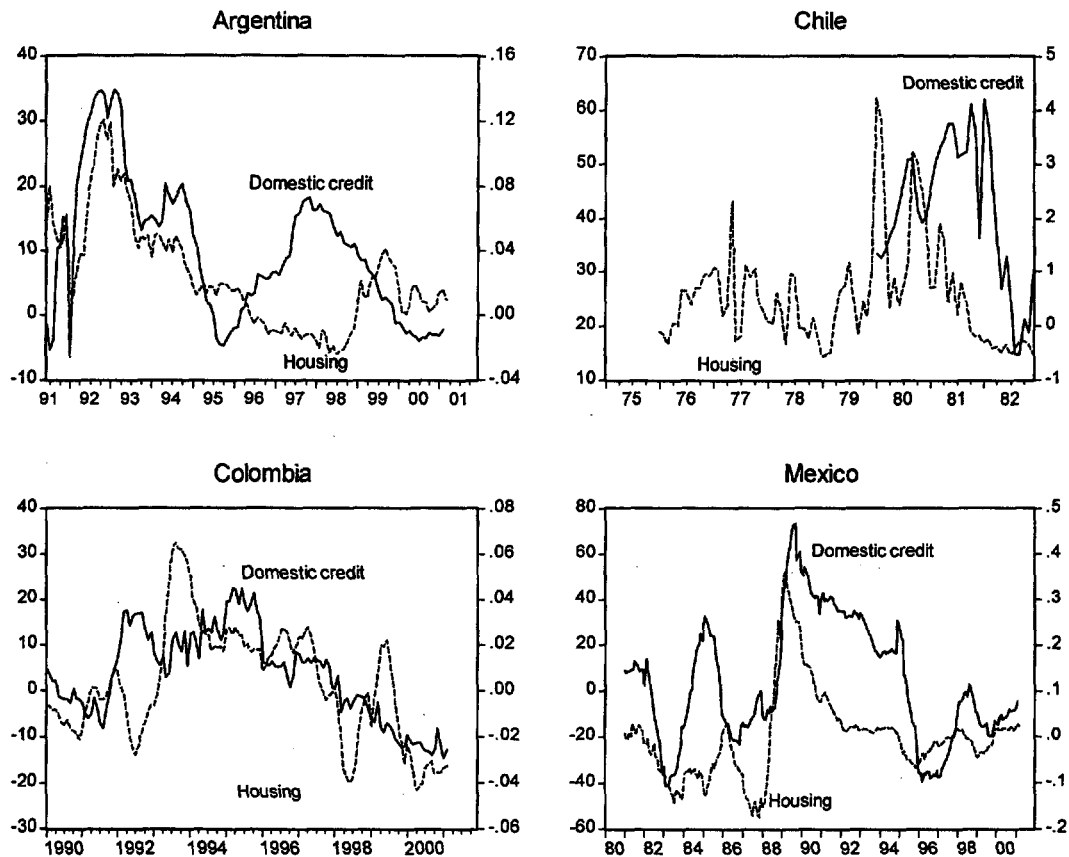


Graph 20
Stock Market Prices and Investment



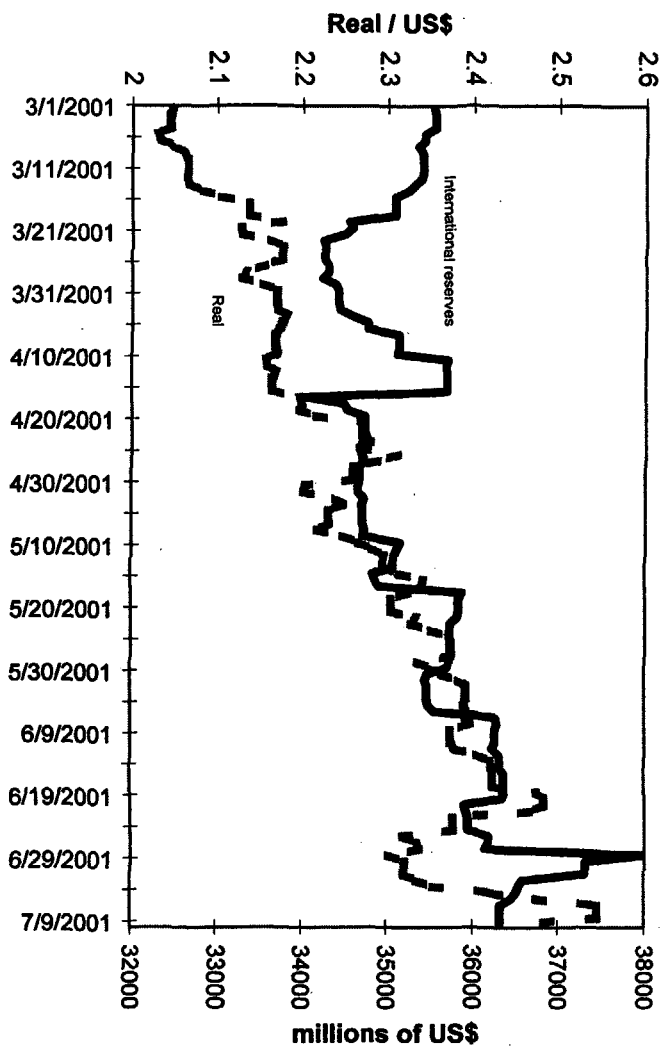
Sources: Stock market price indices are from EMDB database from S&P. Gross fixed capital formation: Argentina, Ministerio de Finanzas y Obras Publicas; Chile, Central Bank of Chile; Colombia, DANE; and Mexico, INEGI via HAVAR database.

Graph 21
RETURN ON HOUSING AND DOMESTIC CREDIT GROWTH IN LAC



Sources: Argentina: domestic credit to the private sector (line 32d..zf IMF IFS); Return on housing is the growth rate of the housing price index (1999=100) from INDEC. Chile: Credit from Central Bank of Chile; Housing prices from Maloney. Growth rate of average prices in constant unit prices in Lo Curro and La Dehesa. Colombia: Financial sector total net credit from revista del Banco de la Republica and Central Bank of Colombia webpage; Return on housing is the growth rate of the housing price index (Dec. 1998=100) from DANE. Mexico: domestic credit is lending by commercial banks to the non-bank private sector, the source is the central bank of Mexico; Return on housing is the growth rate of the housing price index (1994=100) from Banco de Mexico.

Graph 22
Brazil: Real, April - July 2001



Source: Bloomberg.

Table 1
Correlation Matrix of Stock Returns 1980 – 2001

	Argentina	Brazil	Mexico	Peru	Chile	Colombia
Argentina	1.00	0.66	0.81	0.68	0.82	0.65
Brazil	0.66	1.00	0.50	0.42	0.62	0.32
Mexico	0.81	0.50	1.00	0.05	0.92	0.87
Peru	0.68	0.42	0.05	1.00	0.56	0.05
Chile	0.82	0.62	0.92	0.56	1.00	0.86
Colombia	0.65	0.32	0.87	0.05	0.86	1.00

TABLE 2

**Unit Root, Cointegration and Tests for Long Run Relationship between Returns
and (log) Dividend Yields.**

	RALS: Unit Root Test		RALS: Cointegration Test	Johansen Cointegrating Vectors*		ARDL: Test on Long Run Relationship
	Total Returns	Log of Dividend Yield		No Data Trend	Linear Data Trend	
ARGENTINA	I(0) ^a	I(1)	Not cointegrated	0	2	Long-run relationship ^b
BRAZIL	I(1)	I(1)	Not cointegrated	0	0	No long-run relationship
CHILE	I(1)	I(1)	Not cointegrated	0	0	No long-run relationship
COLOMBIA	I(1)	I(1)	Not cointegrated	0	0	No long-run relationship
MÉXICO	I(1)	I(0)	Not cointegrated	0	2	No long-run relationship
PERU	I(1)	I(1)	Not cointegrated	0	0	No long-run relationship

*Results based on the Trace. Two cointegrating vectors reflect the fact of stationarity in the series and therefore there is no cointegrating relation.

^a Results are sensitive to the inclusion (or not) of constant and trend. The series are I(1) when constant and trend are not included in the regressions.

^b Results are sensitive to the inclusion (or not) of constant and trend. There is no long-run relationship when trend is not included in the regressions.

Table 3
Correlation Matrix of Residuals of the Intrinsic Bubble Model 1980-2001

	RINTRINBUB_ARG	RINTRINBUB_BRA	RINTRINBUB_CHL	RINTRINBUB_COL	RINTRINBUB_MEX	RINTRINBUB_US
RINTRINBUB_ARG	1.00	0.03	0.29	-0.21	0.09	-0.16
RINTRINBUB_BRA	0.03	1.00	0.37	0.47	0.23	-0.24
RINTRINBUB_CHL	0.29	0.37	1.00	0.33	0.37	-0.29
RINTRINBUB_COL	-0.21	0.47	0.33	1.00	0.26	-0.09
RINTRINBUB_MEX	0.09	0.23	0.37	0.26	1.00	0.08
RINTRINBUB_US	-0.16	-0.24	-0.29	-0.09	0.08	1.00

Table 4
Correlation Matrix of Residuals of the Intrinsic Bubble Model 1995-2001

	RINTRINBUB_ARG	RINTRINBUB_BRA	RINTRINBUB_CHL	RINTRINBUB_COL	RINTRINBUB_MEX	RINTRINBUB_US
RINTRINBUB_ARG	1.00	-0.25	0.61	-0.31	0.52	-0.01
RINTRINBUB_BRA	-0.25	1.00	-0.16	-0.10	-0.25	-0.27
RINTRINBUB_CHL	0.61	-0.16	1.00	-0.02	0.51	-0.16
RINTRINBUB_COL	-0.31	-0.10	-0.02	1.00	-0.04	0.33
RINTRINBUB_MEX	0.52	-0.25	0.51	-0.04	1.00	0.40
RINTRINBUB_US	-0.01	-0.27	-0.16	0.33	0.40	1.00

Table 5
Dummy variable for bubbles in the stock market 1984:12-2001:03*

Dates	Argentina	Brazil	Chile	Colombia	Mexico	USA
1985-1989	<ul style="list-style-type: none"> • 1985:01-1985:09 • 1987:04 • 1989:05-1989:09 	<ul style="list-style-type: none"> • 1985:08-1986:10 	<ul style="list-style-type: none"> • 1985:03-1985:07 • 1986:03 		<ul style="list-style-type: none"> • 1987:02-1987:09 	<ul style="list-style-type: none"> • 1986:03-1987:10 • 1989:07-1989:12
1990-1994	<ul style="list-style-type: none"> • 1991:03-1992:07 	<ul style="list-style-type: none"> • 1990:02 • 1992:01-1992:02 • 1993:11-1994:11 	<ul style="list-style-type: none"> • 1991:09 • 1992:03-1992:07 • 1994:01-1995:10 	<ul style="list-style-type: none"> • 1992:01-1993:01 • 1993:10-1995:02 	<ul style="list-style-type: none"> • 1991:12-1992:05 • 1992:12-1994:03 	
1995-2001				<ul style="list-style-type: none"> • 1997:08-1998:01 	<ul style="list-style-type: none"> • 1999:12-2000:06 	<ul style="list-style-type: none"> • 1998:02-1998:05

*Except for Colombia 1985:12-2001:03

Number of Bubbles and Crashes

LAC	BUBBLE	CRASH 1-CMAX	CRASH 2-DLPRICE
Total	22	24	41
Total bubbles ending with a crash	14	11	24
Independent bubbles and crashes	6	9	13
Crashes during bubbles or vice versa	2	4	4

Table 6
Tests of Unit Roots for Series with Structural Breaks

	Log of stock market prices in real terms (LGPRLCR)	6-month moving average of capital flows proxy (KFLO6)	Domestic credit in real terms (DCR)	Domestic credit growth in real terms (DCRG)
Argentina	1) 89:02	1) 91:09	1) 89:06	91:05
	2)	2)	2) 84:05, 90:04	
	3) 89:02	3) 91:10	3) 82:08, 89:06	
	SETAR* 89:02 & 91:06			
Brazil	1) 86:11	1)	1) 89:11	92:02
	2) 90:10	2)	2) 91:09	
	3)	3) 92:01	3)	
	SETAR* 93:04			
Chile	1)	1) 91:09, 96:07	1) 89:05	82:06?
	2) 91:07	2) 97:09	2) 83:07, 91:02	
	3)	3) 91:09, 97:12	3)	
Colombia	1) 91:11	1)	1)	91:10
	2) 94:05	2) 97:02	2) 97:01	
	3) 91:06, 98:02	3)	3)	
Mexico	1)	1) 94:03	1)	88:08
	2) 91:11	2) 91:10	2) 92:07	
	3) 86:05, 94:10	3)	3)	
Peru	1) 97:10	1) 94:03	1) 88:02	90:08
	2)	2) 96:09	2) 89:12	
	3)	3)	3)	

Note: Numerals on the table denote the equation used below. The methodology followed for unit root tests is that as described in Zivot and Andrews (1992) and Vogelsang and Perron (1998). Initially we use the "innovational outlier" (IO) model where a dummy for a break in the level is allowed along with a dummy for a break in the trend at an unknown time (equation 1). We used two restricted models (equation 2 and 3), when only a break in the trend or a break in the level were more appropriate. Also to complement the break in the level model, we made use of the SETAR method as described by S.M Potter (1995).

$$y_t = \mu + \beta_t + \gamma DT_t + \theta DU_t + \alpha y_{t-1} + \sum_{i=1}^k c_i \Delta y_{t-i} + u_t \quad (1)$$

$$y_t = \mu + \beta_t + \gamma DT_t + \alpha y_{t-1} + \sum_{i=1}^k c_i \Delta y_{t-i} + u_t \quad (2)$$

$$y_t = \mu + \beta_t + \theta DU_t + \alpha y_{t-1} + \sum_{i=1}^k c_i \Delta y_{t-i} + u_t \quad (3)$$

where $DU = 1$ if $t > T_b$, and 0 otherwise, $DT = t - T_b$ if $t > T_b$, and 0 otherwise, k = number of lags according to the ADF general to specific methodology for lag selection.

Table 7

PANEL DATA: LOGIT RESULTS (ARGENTINA, BRAZIL, CHILE, COLOMBIA AND MEXICO)

Dependent Variable: DUSUM

Method: ML - Binary Logit (Quadratic hill climbing)

Sample(adjusted): 60 1272

Included observations: 815

Excluded observations: 398 after adjusting endpoints

Convergence achieved after 6 iterations

QML (Huber/White) standard errors & covariance

Variable	Coefficient	Std. Error	z-Statistic	Prob.
NSPR103	1.481	0.176	8.407	0.000
NRINTRINBUBUS	-0.282	0.135	-2.098	0.036
NEQREIDEV	0.637	0.151	4.217	0.000
NDCRGHAT(-12)	-0.426	0.195	-2.184	0.029
NDCRGHAT(-18)	0.530	0.143	3.696	0.000
NKFLO6HAT	0.408	0.210	1.938	0.053
NTOTDEV	0.659	0.107	6.150	0.000
NVOL24DLDCRHAT	0.367	0.227	1.614	0.107
NVOL24DLGTRLCRHAT	0.600	0.301	1.996	0.046
DARG	-3.281	0.328	-10.006	0.000
DBRA	-3.465	0.429	-8.085	0.000
DCHL	-1.395	0.275	-5.070	0.000
DCOL	-1.079	0.226	-4.775	0.000
DMEX	-2.117	0.270	-7.843	0.000
Mean dependent var	0.189	S.D. dependent var		0.392
S.E. of regression	0.324	Akaike info criterion		0.706
Sum squared resid	84.322	Schwarz criterion		0.787
Log likelihood	-273.880	Hannan-Quinn criter.		0.737
Avg. log likelihood	-0.336			
Obs with Dep=0		Total obs		815
Obs with Dep=1	661.0			

DUSUM = Bubbles dummy; it is equal to 1 when durintrinbub = 1 (when rintrinbub > 1 standard deviation), or when dudiffnorm dummy = 1 [when the difference of the normalized price (gprlcr) - normalized Present value (K*Dividends)>1.3], and 0 otherwise.

NSPR103 = Normalized Spread between the 10-year and 3 month US t bills

NRINTRINBUBUS = Normalized residual of US intrinsic bubble

NEQREIDEV = Normalized deviations of the equity real estate variable

NDCRGHAT(-12) = Normalized domestic credit growth in real terms at t-12 months

NDCRGHAT(-18) = Normalized domestic credit growth in real terms at t-18 months

NKFLO6HAT = Normalized 6-month moving average capital flows proxy

NTOTDEV = Normalized Terms of trade deviation

NVOL24DLDCRHAT = Normalized 24 month moving window of a standard deviation of the differenced log of domestic credit in real terms.

NVOL24DLGTRLCRHAT = Normalized 24 month moving window of a standard deviation of the first difference of the log of the real global total return series

DARG = Argentina dummy

DBRA = Brazil dummy

DCHL = Chile dummy

DCOL = Colombia dummy

DMEX = Mexico dummy.

Note: Variables with the suffix HATP are fitted values of the variables used as instrumental variables

TABLE 8
Marginal Probabilities

	Domestic credit growth	EQR EIDEV	Capital flows	U.S. residual intrinsic bubble	Spread between 10-year and 3- month t-bill	Terms of trade deviation	Volatility in real domestic credit	Volatility in real total returns
Argentina	0.0798	0.0504	0.0365	-0.0196	0.1199	0.0566	0.0306	0.0523
Brazil	0.0698	0.0441	0.0319	-0.0171	0.1048	0.0495	0.0267	0.0457
Chile	0.0906	0.0573	0.0414	-0.0223	0.1362	0.0642	0.0348	0.0594
Colombia	0.0977	0.0618	0.0447	-0.0240	0.1469	0.0693	0.0375	0.0640
México	0.0922	0.0582	0.0421	-0.0227	0.1385	0.0653	0.0353	0.0604

Table 9
Causality Results: Stock Market Returns, Consumption, and Investment

To	From	D(Investment)	D(Return)
D(Investment)			ARG CHL COL* MEX PER
D(Return)		CHL MEX* PER	

To	From	D(Consumption)	D(Return)
D(Consumption)			ARG COL* MEX* PER
D(Return)		BRA(annual series) PER	

Note: Countries with * the series are both I(1) and we tested cointegration. For those variables no cointegrated we did granger causality in the differences of the variables. In variables cointegrated we performed a wald test for the relevance of the difference of the lagged explanatory variable and the lagged error correction term.

Table 10
Panel results: SUR estimation

Dependent Variable: GDITOIY?
Method: Seemingly Unrelated Regression
Sample: 1980 1999
Included observations: 20
Number of cross-sections used: 5
Total panel (unbalanced) observations: 91
One-step weighting matrix

Variable	Coefficient	Std. Error	t-Statistic	Prob.
C	20.735	1.758	11.795	0.000
FDITOHATP	0.500	0.275	1.817	0.073
GOVYCHATP	-0.099	0.034	-2.908	0.005
GOVYKHATP	0.454	0.110	4.124	0.000
LGPRLCRHATP	0.923	0.361	2.559	0.012
VOL12DLGTRLCRHATP	-10.133	4.585	-2.210	0.030
GROHATP(-1)	0.364	0.155	2.348	0.021
STATEFAILAVG	-12.071	1.861	-6.487	0.000
Weighted Statistics				
Unweighted Statistics				
R-squared	0.393	Mean dependent var	20.69	
Adjusted R-squared	0.341	S.D. dependent var	3.57	
S.E. of regression	2.893	Sum squared resid	694.90	
Durbin-Watson stat	0.811			

GDITOIY = Gross domestic investment (% to GDP)

FDITOIYHATP = Foreign direct investment (% to GDP)

GOVYCHATP = Government consumption (% to GDP)

GOVYKHATP = Public capital expenditures (% to GDP)

STATEFAILAVG = Average of "Episodes of State Failure" as defined by the State Polity 98 from the University of Maryland. It includes Ethnic Wars, Revolutionary Wars, Abrupt or Disruptive Regime Changes, and Genocides/Politocides.

VOL12DLGTRLCRHATP= 12 month moving window of a standard deviation of the first difference of the log of the real global total return series.

GROHATP = GDP growth.

Note: Variables with the suffix HATP are instrumental variables; they are derived by running a panel regression for each (except statefailavg) independent variable (X_{it}) on the other independent variables; the residuals from each panel regression are saved and the fitted values are calculated by subtracting the residuals from the dependent variable. These fitted values are the ones used in the above estimation.

Table 11

	Argentina				Colombia				Mexico			
	Credit to housing as share of total credit	Credit to construction as share of total credit	Housing Price Index/CPI		Credit to construction as share of total credit	Housing Price Index/CPI		Credit to construction and housing as share of total credit	Housing Price Index/CPI			
			(Average)	(Growth)		(Average)	(Growth)		(Average)	(Growth)		
1984												
1985												
1986	10.80		0.69						10.80	0.77		
1987	11.10		0.68	-1.00					9.70	0.71	-7.79	
1988	13.10		0.71	4.40					7.40	0.69	-2.82	
1989*	16.20		0.64	-9.90					6.80	0.59	-14.49	
1990**	13.05		0.71	10.60	27.99	0.91			11.10	0.61	3.39	
1991	14.20	6.20	0.71	10.60	27.99	0.91			11.50	0.78	27.87	
1992	12.80	6.50	0.82	15.60	28.49	0.90	-1.10		13.50	0.89	14.10	
1993	9.80	5.40	0.88	6.80	21.04	0.90	0.00		13.60	0.95	6.74	
1994	10.30	4.20	0.94	7.00	20.55	0.89	-1.11		12.10	0.97	2.11	
1995	12.20	4.40	0.99	1.60	24.46	0.92	3.37		13.10	0.98	1.03	
1996	12.70	4.10	0.99	0.10	30.11	0.94	2.17		13.40	1.00	2.04	
1997	13.00	3.60	0.98	-1.10	30.35	0.96	2.13		14.20	0.96	-4.00	
1998					28.21	0.98	2.08		13.40	0.92	-4.17	
1999					22.69	0.99	1.02		13.70	0.92	0.00	
					16.70	0.98	-1.01					
					18.82	0.97	-1.02					

Sources: Argentina: Boletín Estadístico del Banco Central de la República Argentina. Colombia: Banco de la República. Mexico: INEGI.

Argentina: Housing credit from banks and savings institutions. Construction credit from financial institutions.

Colombia: Construction credit from savings and housing corporations.

* Argentina 1989 credit to housing is datum for Aug. '89 and it excludes foreign currency lending.

** Argentina 1990 credit to housing is datum for January '91 and it includes foreign currency mortgage lending.

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