Do Controls on Capital Inflows Insulate Domestic Variables Against External Shocks?

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Abstract

In this paper we attempt to analyse whether price-based controls on capital inflows are successful in insulating economies against external shocks. We present results from VAR models that indicate that Chile and Colombia, countries that adopted controls on capital inflows, seem to have been relatively well insulated against external disturbances. Subsequently, we use the ARDL approach to co-integration in order to isolate the effects of the capital controls on the pass-through of external disturbances to domestic interest rates in those economies. We conclude that there is strong evidence that the capital controls allowed for greater policy autonomy.

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Introduction

It is usually recognised by analysts that Chile and Colombia were spared from the financial turmoil originated in Mexico by the end of 1994; nevertheless, it is also evident that those two countries suffered difficulties following the Asian and Russian debacle at the end of that decade. In this paper, we will attempt to analyse whether the capital control measures[1] adopted by the two countries were successful in insulating those economies against external shocks. The obvious limitation of those polices lies in the fact that they cannot possibly tackle all the existing transmission channels of external disturbances, therefore they are bound to be ineffective against certain types of shocks, especially, strong ones that are transmitted in multiple ways.

One issue that will inevitably hinder the task we proposed to accomplish is the difficulty in isolating the impact of the unremunerated reserve requirements (URR) from the effects of other policies, such as the exchange rate regime, for instance. Therefore, it will often be the case that our analysis will consider the whole external economic policy "package" and not merely the controls on inflows. Nevertheless, we will attempt to disentangle the specific role played by the capital controls in one the sections below.

The ultimate aims of the URR were the reduction of the costs of sterilised intervention, the improvement of the term structure of external

[1] Those countries adopted unremunerated reserve requirements on capital inflows throughout the 1990s. For a description of these policies see, for instance, David 2005b.

debt and the mitigation of the appreciation of the real exchange rate, whilst giving room for manoeuvre for monetary policy. We believe that the study of its impact on the transmission of external shocks is an important issue that still needs to be addressed. Epstein, Grabel & Jomo (2003) for instance, state that “Chilean-style” controls are effective tools in reducing the impact of external shocks, but do not present systematic empirical evidence to that effect.

This paper is divided into four sections. Firstly, we will review the channels through which the reserve requirements on capital inflows could possibly affect the transmission of international shocks to the domestic economy. Subsequently, we will present some preliminary evidence of the vulnerability of Chile and Colombia to external disturbances and of their response to some crisis episodes when compared to other Latin American countries. In the third section, we will present results from Vector Autoregressive models in order to assess the transmission of global financial shocks to the Chilean and Colombian economies. Finally, in section 4, we will use the ARDL approach to co-integration in order to attempt to isolate the effects of the capital controls on the pass-through of external disturbances to domestic interest rates.

1 The URR and International Shocks

This section will sketch some of the channels through which controls on capital inflows, especially asymmetric[2] taxes, such as the unremunerated reserve requirements, could affect the transmission of international shocks to the domestic economy and, possibly, aid in the prevention of financial crises. Those issues have been thoroughly analysed elsewhere (see for instance David, 2005a and Dooley & Walsh, 2000).

Arguably the most clear-cut case for the introduction of restrictions on capital inflows with the objective of averting financial crises relates to periods of excessive liquidity (euphoria), which are a characteristic of the upswing of financial cycles, as emphasised by Kindleberger (2000). Historically, according to this author, the origin of several financial market bubbles has been associated with some form of "good news" that would yield profitable opportunities to investors in certain assets. In response to these events arbitrageurs begin to bid-up the prices of these assets, trying to sustain the "euphoria" of noise traders. Eventually, when the bubble bursts, the collapse of asset prices entails a collapse in the value of collaterals and may lead to bankruptcies, inefficient liquidations and falls in investment.

During the credit expansion phase, there is an increase in credit risk because the quality of financial intermediaries portfolio's decreases, leading to an increase in non-performing loans. This is aggravated by the fact that financial market agents frequently have inappropriate responses to changes in risk over the economic cycle, such that risks are underestimated during booms and overestimated during recessions. The so-called overborrowing syndrome (Mckinnon & Pill, 1999) sometimes linked to moral hazard due to government guarantees could be considered a special case of the previously outlined general “anatomy” of crises.

In the presence of these dynamics, the adoption of restrictions on capital inflows in a countercyclical manner in other to mitigate the expansionary effects associated with surges in capital inflows seems to be an appropriate response by policy makers in capital importing countries with the objective of reducing economic volatility.

The literature on currency crises highlights other circumstances during which controls on inflows might also be a useful device. The currency mismatch strand of the literature argues that if a substantial amount of debt is denominated in foreign currency due to “Original Sin”[3], a country could become financially fragile (vulnerable to self-fulfilling crises), as emphasised by Aghion et al. (2000). By taxing capital inflows it is possible to restrict the level of external indebtedness and limit the negative real effects of devaluations due to the exposure of domestic firm’s balance sheets. In fact, benefits arise because the decision by an individual firm to borrow in foreign currency imposes costs on the rest of the economy that are internalised by the tax on inflows.

Furthermore, one of the stylised facts in the empirical literature on currency crises is that the maturity structure of external debt seems to

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[2] Those taxes are asymmetric in the sense that they fall more heavily on short-term capital inflows and did not apply to all capital flows (trade credits were usually exempt).

[3] Original Sin is a term coined by Eichengreen and Hausman (1999) that essentially refers to the inability of developing countries to issue debt abroad in their own currency (when they do, those are usually bonds indexed to the exchange rate, which stills fit in the Original Sin concept).

matter for the occurrence and severity of currency collapses (see Carlson and Hernandez, 2002). This fact has been addressed theoretically by the maturity mismatch literature (see especially Rodrik & Velasco, 1999). This type of model emphasises that in the presence of market failures, borrowers would not consider the effects of the stock of short-term debt on contractual interest rates[4] and would choose the privately less costly option of taking short-term rather than long-term debt. Hence, reserve requirements on short-term capital inflows such as the deposit requirements adopted by Chile and Colombia could play a role in reducing the likelihood of liquidity problems by mitigating the externality associated with short-term debt, forcing borrowers to evaluate the social costs of external borrowing translated into financial crises. In fact, an explicit or implicit tax on short-term debt would reduce the difference between the relative costs of short-term and long-term debt, thus reducing the incentives for potentially dangerous excessive short-term capital inflows.

Another important channel through which the URR might be able to reduce the vulnerability of a country to external shocks relates to the prevention of real exchange rate overvaluation. Dornbusch, Goldfajn and Valdes (1995) demonstrate empirically that an overvalued real exchange rate is a fairly good predictor of future currency crises; therefore the impact of the reserve requirements on the real exchange rate seems to be of considerable importance.

There is a vast literature that focuses on specific transmission channels of financial crises from one country/region to another and

[4] In the Rodrik & Velasco model the interest rates paid depend on the level of short-term debt, since it affects the likelihood of long-term debt contracts to be honoured.

emphasises mostly trade and financial links as the main carriers of spill-over effects associated with financial stress (see Dungey et al., 2003 for a survey). The spread of crises through trade links can take place via competitive devaluations or simply by the reduction in demand for exports due to a weaker economic situation in the "ground zero" and other affected countries.

The role of financial links is slightly more complex. Whilst some authors concentrate on common lender effects (a lender is affected by crisis in one country, faces liquidity problems such as margin calls and has to liquidate its assets elsewhere); others focus on "rebalancing effects" in the context of standard portfolio models, where investors pull away from risky assets (emerging markets) when faced with a large negative shock, i.e. an increase in volatility (Schinasi and Smith, 1999). The latter effect is akin to a simple increase in risk aversion by international players. Calvo and Mendoza (2000) emphasise the importance of information asymmetries and costs in acquiring information about specific markets, which encourage herd behaviour by international investors as a major source of "contagion". We could also consider a more general story, where in the presence of multiple equilibria, a crisis in one country can act as a sunspot and coordinate investor's expectations. In this case, the transmission of crises occurs purely because of changes in investor's beliefs.

What role could the URR play in addressing those transmission channels of crises? This type of controls on capital inflows is capable of mitigating problems related to liquidity risk by increasing the maturity structure of external debt, hence it may reduce the exposure to sudden capital outflows associated with the "common lender", "risk aversion", "Calvo and Mendoza-style herd behaviour" and other multiple equilibria stories. Nonetheless, the reserve requirements cannot affect trade links in a direct clear-cut way and therefore cannot prevent distress when trade is a major transmission channel of shocks.

We should conclude this section by noting that one cannot expect controls on capital inflows to insulate an economy against all sources of difficulties in its external sector. These polices may be ineffective against runs (i.e. massive capital outflows) by domestic residents, which were particularly important in the Mexican and Brazilian crisis episodes see Frenkel & Schmukler (1996) and Goldfajn (2000)[5].

2 Preliminary Evidence

We will start this section by analysing selected indicators of external vulnerability for Chile and Colombia presented in Tables 1 and 2. One should note the gradual build-up of vulnerability during the second-half of the 1990s, evidenced, among other factors, by persistent current account deficits and real exchange rate appreciation in both economies, despite a visible amelioration in the maturity structure of external debt. In fact, this deterioration of the vulnerability position of the two countries may be confirmed when we consider other solvency and liquidity indicators.

In the Chilean case, it is apparent that during the 1990s, indicators such as the ratio of M2 to reserves, the ratio of debt service to exports and the ratio of reserves to total debt remained stable (despite decreases in the latter since 1995). Nevertheless, the ratio of total foreign debt to exports presented large increases in the second half of the decade (see Table 1). One should note that this variable is particularly important for currency mismatch stories of financial fragility and currency crises. Models in the vein of Krugman (1999) place the ratio of debt denominated in foreign currency to exports as one of the main determinants of financial vulnerability in developing countries.

As far as Colombia is concerned, the ratio of reserves to total external debt decreased significantly since 1996, the ratio of external debt to exports also deteriorated since 1995 and the ratio of M2 over reserves also increased, reaching a level of 2.7 in 1999. These figures indicate that the solvency and liquidity positions of the country were worse by that period (see Table 2). On the other hand the ratio of debt service to exports remained fairly stable in the second half of the decade.

Hence, despite what could be seen as an overall relatively secure financial position over the decade, which might explain the resilience of those economies against certain shocks, by 1998 the presence of significant disequilibria in what concerns the current account, the real exchange rate and several vulnerability indicators has to be acknowledged.

In the Chilean case, authors such as Ffrench-Davis & Tapia (2001) hold the complacency in capital account management polices verified after 1996 directly responsible for the inability of this country to limit its

[5] Nevertheless, one has to note that by changing the maturity structure of external debt, the controls make runs more expensive and hence could play an indirect role in preventing them.

increasing external vulnerability. They argue that the authorities did not respond to the increasingly alarming signs of fragility, such as the large successive current account deficits, and that the URR should have been increased to face the surge in capital inflows verified after 1996.

This lack of action could be explained by an excessive optimism by Chilean authorities regarding the sustainability of the macroeconomic disequilibria that were arising. Nevertheless, it is also important to note that, during that period, multilateral institutions were strongly encouraging developing countries to adopt full capital account convertibility, i.e. eliminate all forms of capital controls, including the URR. In addition, especially after 1997, it became clear that the central bank gave priority to its inflation stabilisation goal over other objectives, such as maintaining a stable and competitive real exchange rate.

The extensive liberalisation of capital outflows is another factor that might have contributed to the difficulties faced by Chile in its external sector following the Asian and Russian crises. For instance, the percentage of foreign investment by Chilean pension funds permitted by law was raised from 6% of their portfolio in 1995 to 12% in April 1997.

Chile also had strong trade links with the East Asian countries, according to data from the IMF's Direction of Trade Statistics, 37.5% of Chilean exports were directed to Emerging Asia, China, Hong-Kong, Japan and Singapore in 1997. Albeit this may suggest a trade component in the transmission of that crisis, we have to bear in mind that the strongest external shock affecting the Chilean economy was actually related to the Russian default of August 1998. Indeed, by mid-1998 Chile faced large reversals in capital inflows.

Thus, it seems that a number of important transmission channels of external disturbances to Chile are precisely the ones that the unremunerated reserve requirements were not designed to affect, i.e. trade related channels. In addition, it seems that the capital control policies might have been effective in tackling increased vulnerabilities associated with large imbalances, had they been used more actively in the second half of the 1990s.

The deterioration of the external position of Colombia during the 1990s was caused by the unsustainable increase in government expenditures (see Table 2) and the boom-bust financial cycle verified from 1991 to 1998 following the external financial liberalisation process. The credit boom period of the early 90s was accompanied by a large increase in private indebtedness, which was financed, in part, by large capital inflows. The expansionary pressure fuelled a bubble in asset prices and a deterioration of the quality of loans. A pro-cyclical fiscal policy increased aggregate demand even further in the second half of the decade.

In addition, competition by foreign banks that entered the country following the financial liberalisation process may also have contributed to the reduction in the quality of the loan portfolio of domestic banks. In fact, this deterioration in the quality of loans is evidenced by the increase in non-performing loans from around 2% in 1995 to 6% at the end of 1997, 10% in 1999 and 13% in 2000 according to Central Bank figures. Evidently, at least a certain part of this increase in non-performing loans is endogenous to the financial distress verified in 1998/1999.

Later in the decade the public sector resorted to large external borrowings to finance its deficit (Uribe and Vargas, 2002), as we can see from Table 2, the government budget deficit amounted to 3.1% of GDP in 1997, 3.4% in 1998 and 5.1% in 1999. In this context, a succession of large current account deficits took place, rendering the country vulnerable to reversals in capital flows and negative shocks in its terms of trade, both of which occurred in the 1998-1999 period, culminating in the currency crisis of September 1999[6].

Contrary to Chile, Colombia did not have extensive trade links with East Asian countries (in fact, merely 3.6% of Colombian exports were directed to Emerging Asia, China, Hong-Kong, Japan and Singapore in 1997). The bulk of Colombia's trade was concentrated in Latin America and the United States (66.6% of Colombian exports in 1997, once again according to the IMF's direction of trade statistics).

How serious was the 1998/1999 banking and currency crisis in Colombia when compared to the experience of other developing countries around that period? This is an extremely difficult question to answer. The loss in terms of output growth seems to have been significant. As shown in Table 2, output grew by only 0.5% in 1998 and fell by 4.1% in 1999, just to mention the immediate aftermath of the financial difficulties. In addition, according to Herrera and Perry (2003), the rescue costs of distressed financial institutions might have amounted to 8% of GDP.

With the objective of obtaining a first assessment of the response of the economies that adopted controls on inflows to external disturbances, we will examine the behaviour of domestic interest rates during periods of major international financial disturbances. Figure 1 presents weekly data

[6] Note that the devaluation by itself does not mean that controls were ineffective, but simply that they were not sufficient to prevent the bust in the financial cycle.

for 90-day deposit interest rates for Chile and Argentina (a country that did not adopt capital controls) from March 1994 to March 2000.

We can clearly see that Argentina responded more to the Mexican currency crisis of December 1994 than Chile has[7]. Nevertheless, Chilean rates increased significantly more than Argentinean ones following the development of financial difficulties in East Asia and especially after the Russian default in August 1998 despite the presence of capital controls from June 1991 until September 1998. In fact, Chile had a very strong reaction to this episode. Ffrench-Davis and Tapia (2001) argue that the authorities overreacted in order the preserve the credibility of the inflation stabilisation achieved during the decade. As we mentioned before, one should remember that Chile had strong trade links with East Asia and also suffered a severe terms of trade shock at that time. Evidently, the comparison between the reactions of the two countries is only indicative as the policy regimes, inflation rates and the economies themselves are very distinct, i.e. Argentina could hardly be regarded as a rigorous "control group", since the adoption or not of reserve requirements is certainly not the only factor that differentiates the two economies.

Similar conclusions arise when we compare the movements in Colombian and Argentinean deposit rates for the same period as illustrated in Figure 2 (note that Argentinean rates are presented on the left-hand-side axis and Colombian rates on the right-hand-side axis). Colombian deposit rates did not react as much as the Argentinean ones to the Mexican currency collapse, i.e. there was no sudden increase in

[7] One has to bear in mind, since we are using nominal interest rates, the differentials in inflation. In any case, our focus in these figures is the volatility of the series not their level.

Colombian rates as in the Argentinean case following the Peso devaluation. Nevertheless, Colombian rates reacted substantially in the period of “contagion” from East Asia and Russia, although this could also be linked to internal problems, as we emphasised previously.

Figures 3 and 4 present weekly data from June 1997 to June 2002 for deposit rates in Chile, Colombia and Mexico (In Figure 4 Mexican rates are represented on the left hand side axis and Chilean rates on the right hand side axis). It is clear that sharp increases in domestic rates were verified in the three countries following the August 1998 financial crisis in Russia, which indicates that those economies were affected by spill-overs from that period of stress in international financial markets.

This initial cursory look at the movements in interest rates for these countries in periods of high volatility does not allow us to distinguish whether countries that adopted controls on capital flows were systematically affected by external disturbances in a different manner than countries that did not resort to those policies. We will attempt to address this issue in Section 3 below, using VAR models.

Overall, we may conclude that when we simply consider the descriptive evidence to compare the reactions of different countries to crisis episodes, it seems that that Chile and Colombia, two countries that adopted policies to manage actively their capital accounts, have been relatively well insulated against external disturbances, at least until the mid-1990s. Those countries were also relatively successful in keeping macroeconomic imbalances in check during that early period. In the next section, we will attempt to produce more systematic evidence on the transmission of international shocks to Latin American economies using Vector Autoregressive models.

3 Evidence from Vector Autoregressive Models

Our objective in this section is to obtain some stylised facts about the response of domestic macroeconomic variables to external shocks, measured by the spread paid by emerging markets over U.S. instruments of the same maturity. For this purpose, we will use the global Emerging Markets Bond Index (EMBI) spreads constructed by the American investment bank J.P. Morgan. This index tracks the traded market for U.S. Dollar denominated Brady and other similar bonds and is widely regarded as an appropriate measure of foreign investor's risk perceptions regarding emerging markets. We chose to use the EMBI rather than the commonly used EMBI+ spreads (which is a less restrictive index, including Eurobonds and other sovereign debt instruments), because of data availability as the latter series only starts in 1998[8].

We will implement reduced form models of the transmission of international shocks to macroeconomic variables in some Latin American countries using the Vector Autoregressive (VAR) methodology. Other authors such as Edwards (2000) and Edison and Reinhart (2001) have previously used this framework for similar purposes. Applications of VAR models generally consist of estimating an appropriate statistical model of the data and computing impulse response functions and variance

[8] Chilean securities are not included in the EMBI index and Colombian ones were only included after 1998. In any case, the VAR methodology would be able to tackle the possible endogeneity problems that might arise.

decompositions to analyse the dynamic effect of a standard shock in one variable on the rest of the system.

When specifying a reduced-form VAR model it is crucial to assess whether the variables included in the system are stationary or not for inference to be valid, since it is well known that testing hypothesis on coefficients of integrated variables requires non-standard asymptotic theory[9] (Canova, 1995). Harris and Sollis (2003) emphasise that standard univariate Unit Root testing procedures such as the Dickey-Fuller, Augmented Dickey-Fuller and Phillips-Perron tests suffer from a lack of power, i.e. in finite samples they tend not to reject the null hypothesis of non-stationarity, when it should be rejected.

When performing the Unit root tests for the variables included in our models we chose to implement tests using the GLS de-trending procedure proposed by Elliot, Rothenberg and Stock (1996), which optimises the power of the test. We report results obtained for the ADF test with GLS de-trending and also for the Ng-Perron (2001) testing procedure that incorporates both GLS de-trending and a new information criterion for selecting the lag length of the model[10].

In our first model, we will use a foreign exchange market pressure index for Argentina, Chile, Colombia and Mexico as endogenous variables in the system. We construct the pressure index following the empirical literature on currency crises in order to account for the fact that in periods of stress due to external shocks the burden of adjustment does not fall exclusively on interest rates, but it is also reflected in changes in

[9] Although Sims, Stock & Watson (1990) prove that coefficients are consistently estimated independently of the order of integration.
[10] See once again Harris and Sollis (2003) for details.

international reserves and changes in the nominal exchange rate. Our index is a weighted average of changes in domestic interest rates, nominal exchange rates and the log of foreign reserves, where the weights are given by the inverse of the standard deviation of each series such that none of the individual components would dominate the index (see Eichengreen, Rose and Wyploz, 1996 for a similar index). Hence, the pressure index on foreign exchange markets for country j at time t is given by:

$$emp_{jt} = \left(\frac{1}{\sigma_i}\right)\Delta i_t + \left(\frac{1}{\sigma_s}\right)\Delta s_t - \left(\frac{1}{\sigma_R}\right)\Delta R_t$$

We estimated the model from May 1991 to June 2001 and included the global EMBI spread and the pressure indexes, as constructed above for the different countries, as endogenous variables in our system. The data on reserves and nominal exchange rates used to build the pressure indexes comes from the IMF's IFS database. In addition, the Junk bond spread and the log of world oil price index (included to capture real shocks and obtained from the IMF's IFS database) were used as exogenous variables in the system. The series on the risk premium over U.S. government securities paid by American non-investment grade ("Junk") bonds was obtained from the Bloomberg database. This variable has been used in the literature as a measure of international investor's risk appetite (see Mody & Taylor, 2002).

We chose a lag structure of 4 for this model, as there was a conflict between the different information criteria, the LM test does not detect serial correlation of the residuals for this specification. The Unit Root tests performed and reported in Table 3 showed that all the pressure indexes are stationary. Note also that both tests reject the null of a Unit Root for the EMBI spread, the Junk bond spread and the world oil price series at the 5% level.

It is important to note that we decided not to include Brazil in our models due to the fact that the 1994 Real stabilisation plan represents a significant structural break that would have reduced the sample size of the specifications (prior to 1994 this country was experiencing a turbulent period of high inflation). In addition, Brazil adopted controls on inflows of a quite different form (see David, 2005b) and, therefore, neither constitutes an adequate “observation” of a country that imposed reserve requirements nor an appropriate “counterfactual” for our analysis.

One should also bear in mind that we have included in our sample a number of years during which the controls on capital inflows were not in place as well. Nevertheless, we believe that this does not pose a problem as far as the interpretation of the results is concerned, as the controls were designed to deal with surges in capital inflows lasting several months, which were not verified in the post 1998 period. Hence, tax-equivalent rates would probably have been small (non-binding) if the controls were in place. On the other hand, considering a larger sample may allow us improve the reliability of the estimation and infer the lasting effects (in the medium term) of those polices in terms of a reduction of financial fragility.

The generalised impulse responses presented in Figure 5 demonstrate that the pressure indexes for Chile and Colombia do not respond significantly[11] to shocks to the EMBI spread nor do they respond significantly to shocks to the other countries' (Argentina and Mexico) pressure indexes. Nevertheless, the Argentinean and Mexican indexes do respond to EMBI shocks and Mexico is the country that presents the strongest response. It is also of interest to note that the Argentinean pressure index presents a statistically significant response to shocks in the Mexican pressure index indicating that financial "contagion" occurred between these two countries.

The advantage of using generalised impulse responses when compared to standard Cholesky ones is that they do not depend on the ordering of the variables in the system. Generalised impulse responses compare the conditional expectation of a variable in the model, given a shock and the history of the model, to the conditional expectation of that variable given the historically observed information of the model (see Koop, Pesaran & Potter, 1996 for details). The results obtained from the impulse response functions are robust to the impulse definition, for instance when a Cholesky decomposition was used we got qualitatively similar responses.

Table 4 shows the variance decomposition for the VAR model estimated above. One should note that only a small percentage of the forecast errors in the Chilean and Colombian pressure indexes can be attributed to EMBI shocks. In fact, the figures are 0.07% and 1.17% for the first month for each country respectively. Nevertheless, when we look

[11] Throughout this paper a "significant" impulse response for each month following a shock means that the interval defined by the error bands does not contain the value zero. Note that error bands were calculated using Monte Carlo simulations (1000 repetitions).

at the Argentinean and Mexican indexes for the same horizon, those figures become much larger (15.69% and 40.26% respectively).

These results indicate that the unremunerated reserve requirements (combined with other capital account polices) might have helped to insulate the Chilean and Colombian economies from certain types of global external financial shocks, namely the ones captured by the EMBI spread. Evidently, at this stage, we cannot distinguish whether this difference is due to capital account policies, other macroeconomic policies, or simply the type of exchange rate regime adopted by the different countries. One also has to note that the precise role played by the capital controls in insulating those economies was not clarified in our empirical analysis so far.

In order to confirm the validity of our results we decided to estimate models for Chile and Colombia individually, including a wider selection of macroeconomic variables. Firstly, we estimated a model of the Chilean economy using monthly data from January 1991 to December 2000[12], including as endogenous variables the EMBI spread, the output gap, the difference between inflation in the past 12 months (change in the consumer price index) and the inflation target, the exchange rate indexed deposit rate, the Chilean sovereign risk premium and the cyclical component of the real exchange rate. We also included as exogenous variables the logarithm of the Chilean terms of trade and the Junk bond spread. The Akaike, Schwartz and H-Q information criteria suggested a model with two lags and there is no evidence of serial correlation for this specification according to the LM test.

The construction of the Chilean (and Colombian) country risk series is described in David, 2005b. The output gap variable used for Chile was defined as the difference between the log of the economic activity index (IMACEC) and its trend and irregular components obtained using a structural time series model (estimated by Kalman filtering techniques). The monthly inflation target series was kindly provided by Gallego at al. (2002) and the cyclical component of the real exchange rate was obtained by fitting a structural time series model to the series and removing the trend and irregular components. The source for all the variables included is the Central Bank of Chile, except for the terms of trade series that was constructed by Benett and Valdes (2001). The Unit Root tests performed (see table 5) show that the null of non-stationarity is rejected at the 5% level for the inflation, interest rate and country risk series. Non-stationarity is also rejected for the terms of trade series at the 10% level and for the output gap and real exchange rate series at the 1% level.

The generalised impulse responses are presented in Figure 6. One should note that the domestic exchange-rate-indexed deposit rate does not present a statistically significant response to shocks to the EMBI spread, whereas the real exchange only presents a marginally significant response between the second and fourth months. Therefore, the impact of EMBI shocks on the real exchange rate seems to be relatively short-lived and small. In addition, the domestic deposit rate also seems to be resilient to shocks to the Chilean sovereign risk premium and so does the real exchange rate.

[12] We chose this period because of data availability issues and to ensure that a single policy regime is considered.

These conclusions are confirmed by the variance decomposition analysis (see Table 6). Shocks to the EMBI spread are only responsible for a small part of the forecast errors in the real exchange rate and the domestic deposit rate. In the first period, the EMBI accounts for 1.13% of forecast errors in the interest rate and 1.17% in the real exchange rate, whereas in period 10 the figures are 4.79% and 10.43% respectively. Hence, it seems that interest rates were insulated against external shocks in Chile, whereas the real exchange rate is slightly more vulnerable.

To sum up, the URR and other capital account policies seem to have been capable of reducing the pass-through of external shocks to the Chilean economy, especially as far as deposit interest rates are concerned. Nonetheless, the capital account policies did not completely insulate Chile, as the real exchange rate was more vulnerable to shocks.

In addition, we estimated a similar model for the Colombian economy for the period from January 1993 to June 2002[13]. We included as endogenous variables the EMBI spread, the output gap, the difference between the inflation rate and the annual inflation target set by the Central Bank, the 90-day real deposit rate[14], the Colombian sovereign risk premium and the cyclical component of real exchange rate. We also included as exogenous variables the log of the world oil price index, the Junk Bond spread and a dummy variable that takes the value of zero before September 1999 and one thereafter to account for the change in the exchange rate band.

The output gap variable was constructed as the difference between the log of the industrial production index and its trend, seasonal and

[13] The period was chosen because of data availability.

irregular components using a structural time series model. The cyclical component of the real exchange rate was also obtained by de-trending the series using a structural time series model. The primary source for all these variables is the Central Bank of Colombia.

The Akaike information criterion suggests a model with two lags and the LM test does not detect serial correlation for that specification. Unit Root tests reported in table 7 show that all the variables included are stationary, except for the country risk premium. We decided to estimate a model including this variable in levels to make the interpretation of the results more clear. McCallum (1993) argues that if the residuals of each equation in the system are stationary and there is no evidence of serial correlation, the model in levels can be correctly estimated. The Unit Root tests presented in the appendix indeed show that the residuals of each equation in the system are stationary.

The generalised impulse response functions obtained are reported in Figure 7. One can observe that the domestic interest rate presents a marginally significant response to EMBI shocks between periods 2 and 5. The real exchange rate also presents a marginally significant response to EMBI shocks from period two to four. Hence, it seems that both domestic interest rates and the real exchange rate were relatively insulated against external financial disturbances, as the effects of EMBI shocks were small (barely different from zero) and short-lived.

When shocks to the country risk premium are considered, it is possible to observe that according to this model, the domestic deposit rate only presented a marginally significant response to country risk shocks.

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[14] The nominal interest rate minus inflation in the past 12 months.

Nonetheless, the country risk variable seems to have larger effects on the real exchange rate (the response of the real exchange rate to country risk shocks is significant until the forth month and the shock dies out after 6 months).

The variance decomposition analysis (see Table 8) shows that shocks to the EMBI spread explain 23.24% of the forecast errors in the Colombian sovereign risk after 2 months, but only 3.32% of errors in interest rates and 2.66% of errors in the real exchange rate. These results clarify the conclusions obtained from the impulse response functions. The Colombian country risk seems to co-move with global risk premia, whereas domestic interest rates and the real exchange rate are relatively insulated from those shocks. We may conclude that the capital account management policies were relatively successful in insulating the Colombian economy against global shocks (as measured by the EMBI spreads).

4 Disentangling the Effects of the URR

In this section, we will attempt to isolate the possible role of the reserve requirements on capital inflows as far as the pass-through of external shocks to domestic interest rates is concerned and confirm whether those capital account restrictions contributed towards a greater autonomy for domestic monetary policy. Indeed, following the results from previous sections, our main conjecture is that the URR did reduce the effect of foreign shocks on domestic rates even in the context of a crawling band exchange rate regime.

It is well known in the econometrics literature that interest rates frequently behave in ways close to Unit Roots in finite samples. Nonetheless, it seems counterintuitive that they would present Unit Roots as this would mean that some series could go to infinity and others reach negative values. Series that are quite persistent in this way can present the same estimation problems in standard models as regressions involving non-stationary variables. In addition, the application of standard co-integration analysis tools such as the Johansen procedure requires the classification of the variables included in the model into I(1) or I(0), which is difficult in this case.

We will estimate error correction models that circumvent that issue for the Chilean indexed policy interest rate and include the 90-day Libor interest rate (in real terms) and the tax equivalent of the reserve requirement as explanatory variables. Hence, we implemented a model describing the dynamics of the Chilean interest rate, which is given by:

$$\Delta i_t = C + \sum_{i=1}^{n} \beta_i \Delta i_{t-i} + \sum_{j=1}^{m} \alpha_j \Delta i^*_{t-j} + \sum_{w=1}^{k} \delta_w \Delta URR_{t-w} + \psi(i_{t-1} - \lambda_1 i^*_{t-1} - \lambda_2 URR_{t-1})$$

This expression determines the first differences in the domestic interest rate as a function of past changes of the rate itself, past changes in the LIBOR rate (i^*), changes in the URR and an error-correction term.

As mentioned before, it is not possible to determine unambiguously whether the variables that would form the long run relationship are stationary or not. Therefore, we decided to adopt the Autoregressive-Distributed Lag (ARDL) approach to Co-integration proposed by Pesaran et al. (2001) that does not require the classification of variables into stationary and non-stationary. This approach consists of two stages.

Firstly, a test for the existence of a long-run relation between the variables is implemented. It amounts to a simple F-test for the joint significance of the lagged levels of the variables in the error-correction model outlined above. This F-statistic has a non-standard distribution irrespective of whether the variables are I(0) or I(1) and the appropriate critical value bounds have been tabulated by Pesaran et al. in order to test for the null of no co-integration (i.e. coefficients are not significantly different for zero).

In fact, there are two sets of critical values: one assuming that all the variables in the model are I(0) and the other assuming that all the variables are I(1). This provides a band covering all possible classification of variables. If the computed F-statistic falls outside the band, it is possible to reach a conclusion regarding the existence of a co-integration relationship. On the other hand, if the statistic falls inside the band the result of the inference is inconclusive and depends whether the variables are I(1) or I(0). The second stage of the analysis refers to estimating the coefficients for the long-run relationships and making inferences about their values.

We estimated models for the Chilean rate on lagged values of itself, lagged values of the LIBOR interest rate (obtained from the IMF's IFS database) and lagged values of the tax equivalent of the reserve requirements[15] for the period from January 1991 to December 2000 (six lags were used in all specifications)[16]. We experimented two basic

[15]Details of the calculation of the tax-equivalent of the URR were given in David, 2005b. In our specifications, we used an average of the tax equivalent for different maturities of capital inflows.
[16] One of the reasons why we did not extend our sample beyond December 2000 is that the inclusion of the 1999 dummy variable would change the asymptotic

different models: one with a dummy for the turbulent period following the Russian default in 1998, which represents a transition away from the use of controls on inflows and another model including a dummy for the change in the exchange rate regime in September 1999.

As argued by Pesaran et al. (2001), ARDL estimation is applicable even when the explanatory variables are endogenous, provided that the order of the model is large enough to account for contemporaneous correlations between errors in the data generating process. In our case, the foreign interest is evidently long-run forcing and our tests for a model including the URR as the endogenous variable fail to reject the null that the level variables do not enter significantly in the equation, which indicates that the URR can be treated as an exogenous variable[17].

The F-statistic for those models of Chilean interest rates are 11.834 and 4.947 respectively, whereas the critical bounds at the 5% level, calculated by Pesaran et al. are 3.10 and 3.87 (the critical bounds at the 1% level are 4.13 and 5.00). Hence, we can conclude that there exists a long run relationship between those variables for the period analysed, irrespectively of their order of integration, as the test statistic exceeds the critical bounds.

We now proceed to the estimation of the ARDL model. The Akaike information criterion suggests a ARDL(2,0,4) model for the specification including the 1998 dummy. If we consider the long-run relationship

results of the tests, if the fraction of the sample during which the dummy is different from zero does not go to zero as the sample size increases (see Pesaran et al, 2001, p.307).

[17] The test statistic is 1.83, which is below the critical bounds both at the 5% level and at the 10% level.

between the variables, the estimated coefficients are (with p-values in brackets):

$$i_t = \underset{[0.00]}{4.51} + \underset{[0.00]}{0.27}\, i_t^* + \underset{[0.00]}{0.20}\, URR_t + \underset{[0.00]}{2.96}\, Dummy1998$$

One should note that all the coefficients are highly significant (even at the 1% level). The coefficient of the URR is positive indicating that the URR allowed for higher interest rates in the long-run than would have been the case in the absence of capital controls. An increase in the URR by one standard deviation would lead to an increase in the interest rate of 0.2 percentage points.

Hence, the adoption of the URR mitigated part of the expansionary pressures due to high capital inflows and allowed for greater monetary policy autonomy. One also has to note that the coefficient for the foreign interest rate is only 0.27, therefore indicating that the adjustment of local rates to changes in foreign rates is less than proportional, thus confirming the evidence of relative monetary independence.

The estimation results for the error-correction model outlined above are presented in Table 9. All regressors are statistically significant and the diagnostic statistics do not show any evidence of serial correlation of the residuals and the high R-squared (0.56) indicates a good fit.

One interesting feature of the results presented above concerns the coefficient for the error correction term, which is relatively small (-0.306). This implies that the half-life, calculated as ln(0.5)/ln(1+error-correction coefficient) is of approximately 2 months. This suggests that the speed of adjustment of domestic interest rates to foreign ones is only moderate, hence indicating greater monetary policy independence in the Chilean case. Therefore, we can conclude that the evidence so far corroborates our previous inferences concerning the insulating properties of controls on capital inflows.

Subsequently, we also applied the ARDL approach to analyse the effects of the reserve requirements on Colombian 90-day deposit interest rates[18]. We estimated a very similar model to the one outlined for Chile. The sample period goes from January 1991 to December 2000 and a dummy variable taking the value of 1 after September 1999, when the exchange rate band was abandoned was included in the regressions[19]. The F-statistic for this model is 4.242, which lies outside the bounds at the 5% level, but lies inside the bounds at the 1% level, therefore yielding inconclusive results. Nonetheless, given the small sample size, significance at the 5% level is more than adequate.

Hence, we can be reasonably confident that a long run relationship between those variables exists, irrespective of the order of integration of the variables. In addition, the tests indicate that the URR can be treated, once again, as a long run forcing variable (the F-statistic for the model with the URR as the dependent variable is 1.84, which is below the bounds at the 5% level).

The Akaike information criterion suggests a ARDL(4,1,6) model. The estimated coefficients for the long run relationship are (with p-values in brackets):

$$i_t = \underset{[0.07]}{3.96} + \underset{[0.02]}{1.74} i_t^* + \underset{[0.08]}{0.11} URR_t - \underset{[0.06]}{5.08} Dummy1999$$

[18] We used real (*ex post*) interest rates in this case to facilitate the comparison with the Chilean experience and eliminate the noise created by volatile inflation.
[19] As in the previous case, the measure of the tax equivalent used is an average for capital inflows of different maturities.

One should note that all the coefficients are significant at the 10% level and have the expected signs. The coefficient of the URR is positive permitting us to infer that the URR allowed for higher interest rates in the long-run. Nonetheless, the coefficient is smaller than in the Chilean case. An increase in the URR by one standard deviation would lead to an increase in the domestic interest rate of 0.11 percentage points.

The coefficient for the foreign interest rate exceeds 1, implying more than proportionate changes in domestic rate when the foreign rate changes. This is not very surprising for an emerging market and it's in line with the evidence towards over-adjustment of interest rates in developing countries to shocks in foreign interest rates provided by Frankel et al. (2004).

The estimation results for the error-correction model are presented in Table 10. Most regressors are statistically significant at conventional levels and the diagnostic statistics do not show any evidence of serial correlation of the residuals. Once again, the relatively high R-squared indicates a good fit of the model (although the normality of the residuals is rejected at the 5% level, but not at the 1% level).

The coefficient for the error correction term is small (-0.145), which indicates that adjustment towards equilibrium occurs slowly. In fact the half-life for Colombian rates is approximately 4 months (twice the one of Chile). This suggests, once again, a degree of monetary autonomy. Therefore, the evidence seems to indicate that the controls on capital inflows adopted were capable of mitigating the impact of external shocks on the domestic economy for both the Chilean and Colombian cases.

Conclusion

As emphasised by Ocampo & Palma (2005), unremunerated reserve requirements can be an effective flexible instrument to balance macroeconomic disequilibria arising from surges in capital inflows, such as large current account deficits linked to an excessive real exchange rate appreciation. Furthermore, this type of capital controls is instrumental for the application of countercyclical macroeconomic policies; since they allow for greater policy autonomy in the face of pro-cyclical capital flows, as confirmed by the evidence presented in this paper.

It is clear that Chile and Colombia were affected by spill-overs from financial stress related to the East Asian and Russian crises. Nonetheless, what was of direct interest in our analysis was how much these economies have been affected when compared to other countries. We concluded that, when a very specific type of financial shock is considered, namely shocks to the EMBI index, the Chilean and Colombian economies seem to have been relatively well insulated against external disturbances, when compared to other large Latin American countries.

In addition, when we look at the two countries individually, we found that, in the case of Chile, global financial shocks as measured by the EMBI index; did not affect domestic interest rates and only marginally affected the real exchange rate. The Colombian economy also seems to have been relatively well insulated against external disturbances, as EMBI shocks only marginally affected both interest rates and the real exchange rate. One has to bear in mind that the reserve requirements are not the only policy measures that could possibly have reduced the external vulnerability in those countries.

Furthermore, by disentangling the specific effects of the capital controls on domestic interest rates, using the ARDL approach to Co-integration, we were able to demonstrate that there is strong evidence that the reserve requirements allowed for greater policy autonomy and did reduce the pass-through of external disturbances to the economies considered. In fact, the effects of foreign interest rate shocks on domestic rates were reduced even in the context of a crawling band exchange rate regime.

One has to acknowledge that the insulating properties of the controls are not complete, in the sense that they are only capable of reducing the pass-through of external shocks, not eliminate it. Despite the fact that controls on capital inflows have the potential to be beneficial in terms of reducing financial vulnerability, this does not mean that these policies were applied “optimally” by the two countries at all moments in time.

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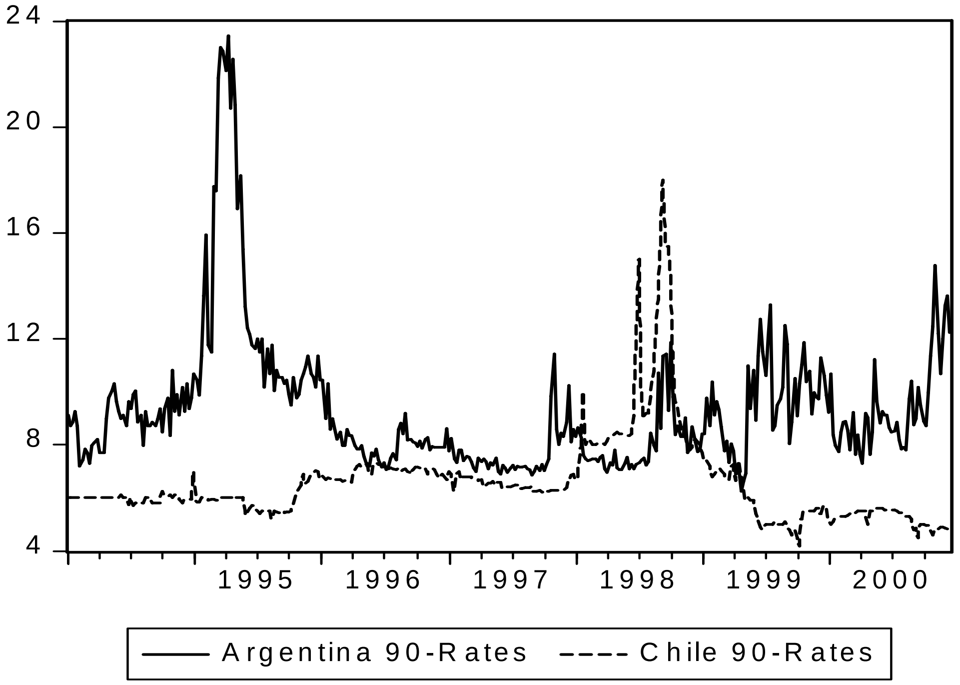
Figure 1

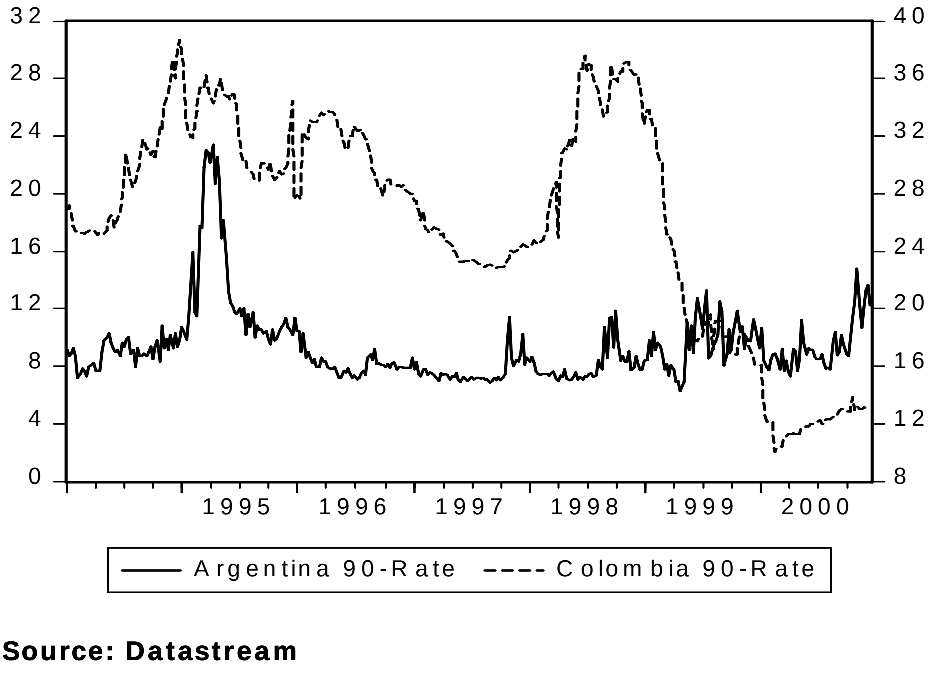
Chilean and Argentinean Interest Rates (1994-2000)



Source: Datastream.

Figure 2

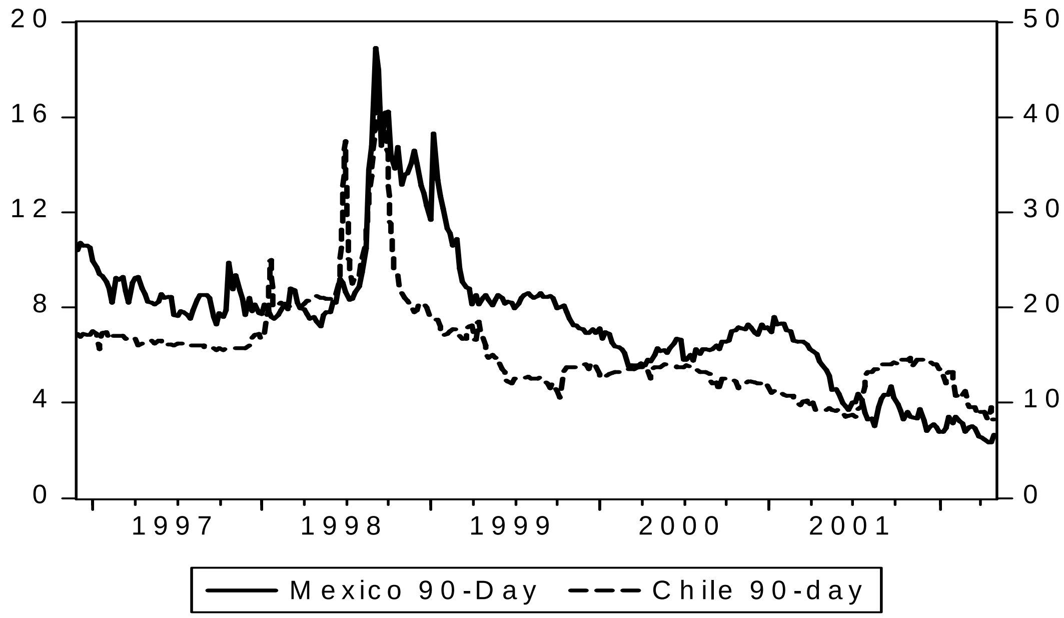
Colombian and Argentinean Deposit Rates (1994-2000)



Source: Datastream

Figure 3

Mexican and Chilean Deposit Rates (1997-2002)



Source: Datastream

Figure 4

Mexican and Colombian Deposit Rates 1997-2002

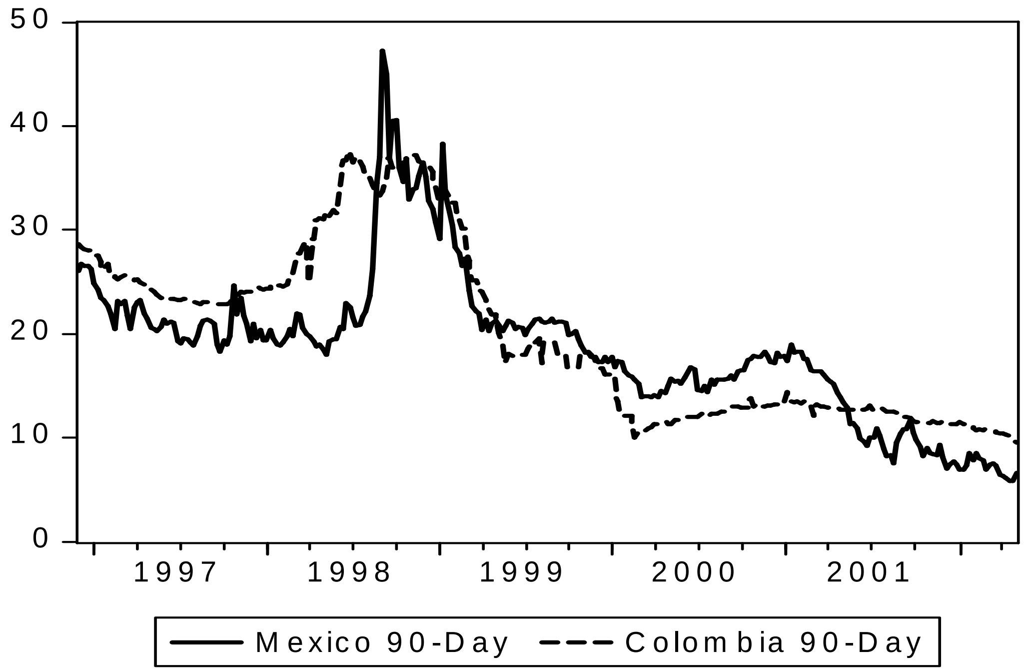


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Figure 5

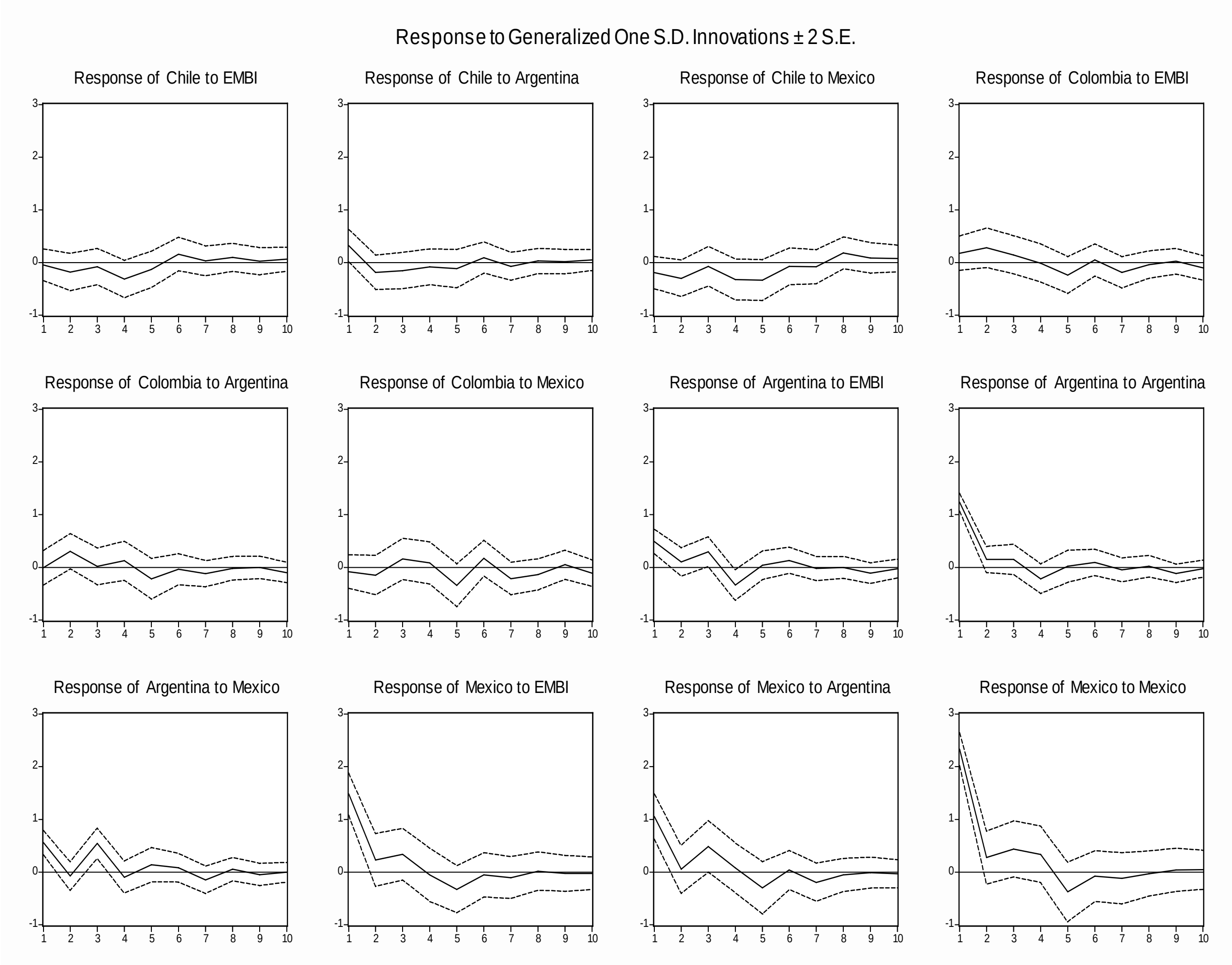

Response to Generalized One S.D. Innovations ± 2 S.E.
Response of Chile to EMBI
Response of Chile to Argentina
Response of Chile to Mexico
Response of Colombia to EMBI
Response of Colombia to Argentina
Response of Colombia to Mexico
Response of Argentina to EMBI
Response of Argentina to Argentina
Response of Argentina to Mexico
Response of Mexico to EMBI
Response of Mexico to Argentina
Response of Mexico to Mexico

Figure 6

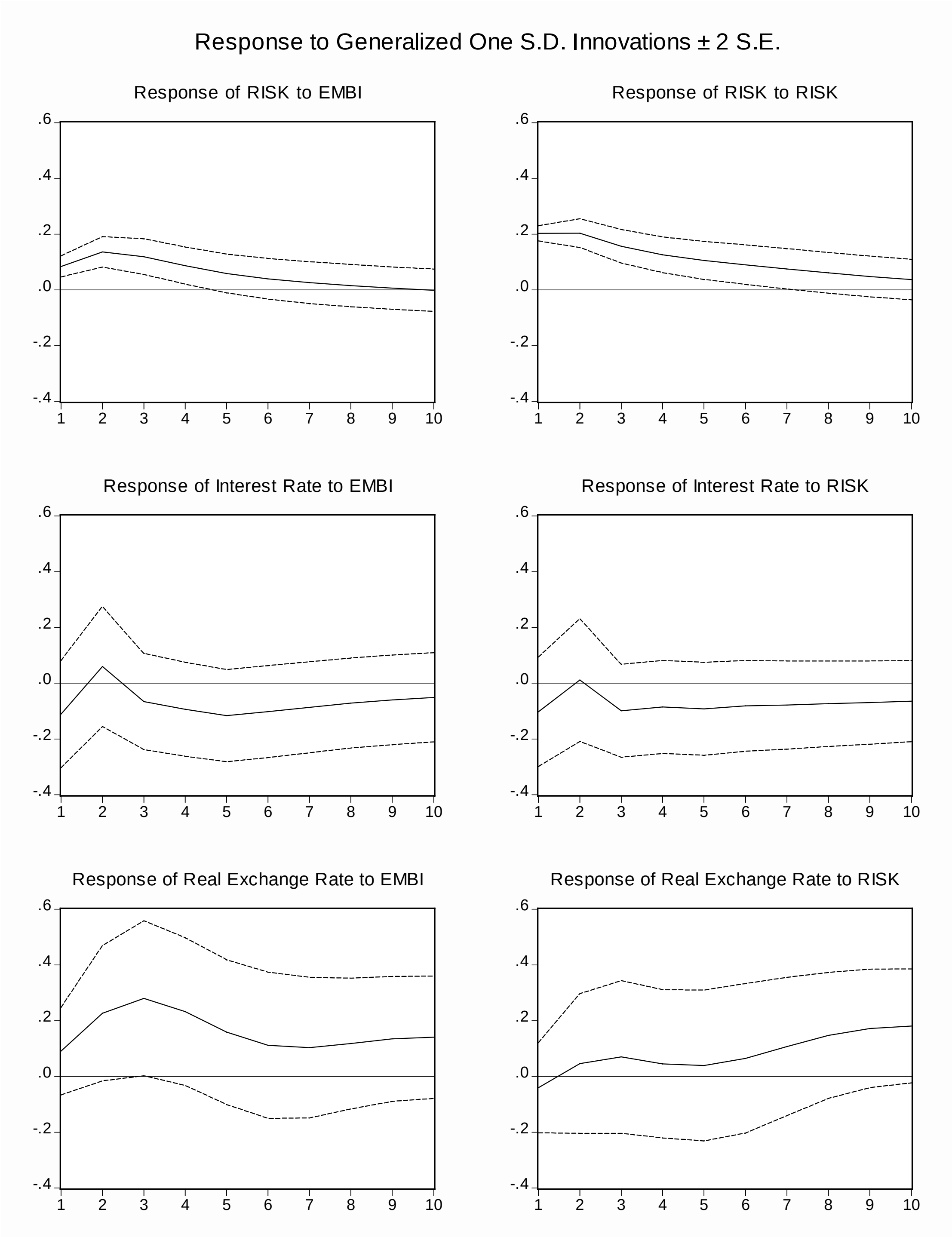
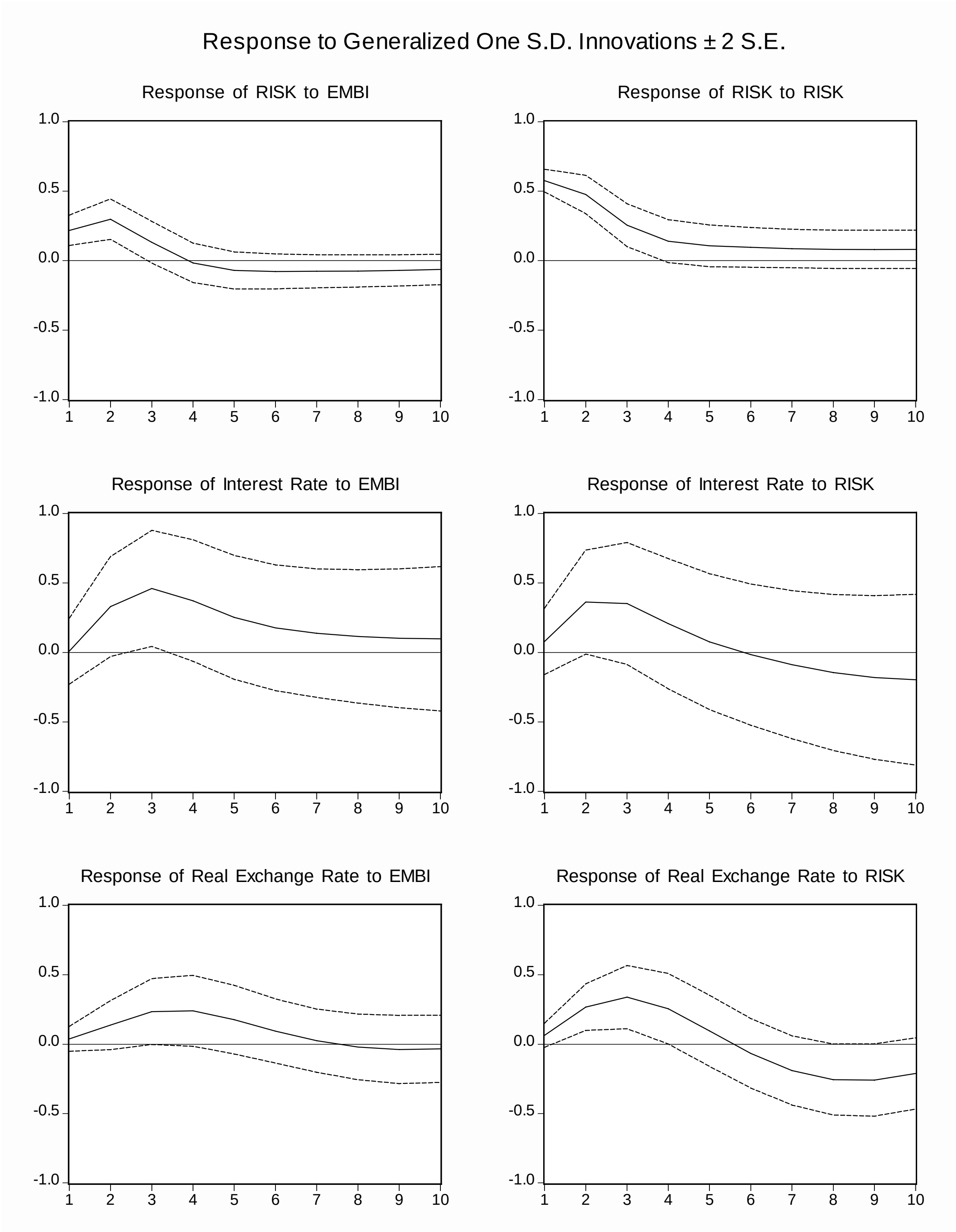

Response to Generalized One S.D. Innovations ± 2 S.E.
Response of RISK to EMBI
Response of RISK to RISK
Response of Interest Rate to EMBI
Response of Interest Rate to RISK
Response of Real Exchange Rate to EMBI
Response of Real Exchange Rate to RISK
.6
.4
.2
.0
-.2
-.4
1 2 3 4 5 6 7 8 9 10

Figure 7



Response to Generalized One S.D. Innovations ± 2 S.E.
Response of RISK to EMBI
Response of RISK to RISK
Response of Interest Rate to EMBI
Response of Interest Rate to RISK
Response of Real Exchange Rate to EMBI
Response of Real Exchange Rate to RISK
1.0
0.5
0.0
-0.5
-1.0
1 2 3 4 5 6 7 8 9 10

Table 1
Selected Indicators of External Vulnerability

Chile

	1990	1991	1992	1993	1994	1995	1996	1997	1998	1999
Short-term Debt/ External Debt (c)	17.6	12.3	16.9	20.0	17.4	15.6	11.4	5.6	5.3	3.4
Reserves/Total Debt (a)	35.3	42.9	51.2	50.2	55.8	58.1	56.6	56.7	44.1	38.2
External Debt/Exports (a)	179.6	154.1	148.0	168.9	164.8	126.3	137.9	143.8	180.8	184.2
Debt Service/Exports (a)	25.8	23.2	20.8	23.2	19.7	25.3	30.9	20.2	22.3	25.4
M2/Reserves Ratio (a)	1.8	1.8	1.6	1.7	1.4	1.7	1.8	1.8	2.0	2.3
Current Account/GDP (b)	-1.6	-0.3	-2.3	-5.7	-3.1	-2.1	-5.1	-5.0	-5.7	-0.1
Trade Balance[I] (b)	1080.3	1659.0	1206.7	125.1	321.5	-1501.6	-1686.6	-2486.9	-2522.8	1968.6
GDP Growth (%) (d)	3.7	7.9	12.2	6.9	5.7	10.6	7.2	7.5	3.9	-1.1
Government Budget Surplus (% of GDP) (c)	0.8	1.5	2.3	2.0	1.7	2.6	2.3	2.0	0.4	-1.5
Real Exchange Rate[II] (c)	112.7	106.4	97.6	96.9	94.3	88.9	84.7	78.2	78.0	82.3

Sources: Global Development Finance (a), World Development Indicators (b), Central Bank of Chile (c) and ECLAC Statistical Division (d).

I Exports of Good & Services minus Imports of Goods & Services (in Millions of U.S. Dollars of 2000).

II Real exchange rate index, 1986=100 a decrease in the index indicates an appreciation of the real exchange rate.

Table 2
Selected Indicators of External Vulnerability

Colombia

	1990	1991	1992	1993	1994	1995	1996	1997	1998	1999
Short-term Debt/ External Debt (c)	8.4	10.2	16.5	18.8	20.5	22.1	20.4	18.1	18.7	11.5
Reserves/Total Debt (a)	28.3	39.7	45.7	42.5	37.4	34.2	34.7	31.5	26.6	23.5
External Debt/Exports (a)	181.0	166.1	167.3	172.8	178.4	182.7	199.3	201.9	224.7	277.3
Debt Service/Exports (a)	40.9	36.3	38.8	33.8	45.3	31.7	37.3	28.7	30.8	43.5
M2/Reserves Ratio (a)	1.6	1.1	1.1	1.4	1.9	2.1	1.9	2.5	2.7	2.7
Current Account/GDP (b)	1.3	5.6	1.8	-3.7	-4.5	-5.0	-4.9	-5.5	-5.3	-0.2
Trade Balance[I] (b)	2841.9	3804.5	1311.6	-2222.7	-5446.1	-6313.1	-5492.6	-6272.8	-4342.9	1632.3
GDP Growth (%) (d)	4.0	2.4	3.8	5.4	5.8	5.2	2.0	3.4	0.5	-4.1
Government Budget Surplus (% of GDP) (c)	-0.7	0.2	-0.2	0.1	1.0	-0.6	-2.0	-3.1	-3.4	-5.1
Real Exchange Rate[II] (c)	132.5	128.7	118.3	112.1	100.0	98.6	92.3	82.8	86.9	96.8

Sources: Global Development Finance (a), World Development Indicators (b), Central Bank of Colombia (c) and ECLAC Statistical Division (d).

I Exports of Good & Services minus Imports of Goods & Services (in Millions of U.S. Dollars of 2000).

II Real exchange rate index, 1994=100 a decrease in the index indicates an appreciation of the real exchange rate

Table 3

Unit Root tests for Selected Variables

	ADF-GLS test statistic	Ng-Perron test statistic
EMBI Spread	-2.284 (**)	-2.223 (**)
Argentina Pressure Index	-4.628 (***)	-4.056(***)
Chile Pressure Index	-4.809(***)	-4.037(***)
Colombia Pressure Index	-7.817(***)	-5.185(***)
Mexico Pressure Index	-5.353(***)	-8.804(***)
Junk Bond Spread	-2.198 (**)	-2.205 (**)
World Oil Prices	-2.755(***)	-2.847(***)

(**) Denotes significance at the 5% level and (*) significance at the 10% level. Lag selection based on Schwartz information criterion. We chose to follow the sequential testing strategy proposed by Perron (1988) when modelling deterministic components.

Table 4

Variance Decompositions

Percentage of Variance Associated with EMBI Shocks

Horizon	Chilean Pressure Index	Colombian Index	Argentinean Index	Mexican Index
1	0.07	1.17	15.69	40.26
2	1.23	3.25	15.54	39.50
5	4.49	5.11	20.12	35.98
10	5.29	6.18	20.25	34.87

Table 5

Unit Root Tests for Selected Variables

	ADF-GLS test statistic	Ng-Perron test statistic
Output Gap	-10.477 (***)	-6.534 (***)
Inflation Gap	-2.204(**)	-2.136 (**)
Indexed Interest Rate	-2.711(***)	-2.598 (***)
Chilean Country Risk	-2.659 (***)	-2.496 (**)
Terms of Trade	-1.879 (*)	-1.720 (*)
Real Exchange Rate Gap	-5.487 (***)	-5.419 (***)

(***) Denotes significance at the 1% level, (**) significance at the5% level and (*) significance at the 10% level. Lag selection based on Schwartz information criterion. A constant was included in all tests. The Perron sequential procedure was used once again.

Table 6

Variance Decompositions

Percentage of Variance Associated with EMBI Shocks

Horizon	Chilean Interest Rate	Chilean Country Risk	Chilean Real Exchange Rate
1	1.13	17.28	1.17
2	1.37	29.54	3.67
5	3.01	33.69	8.87
10	4.79	28.83	10.43

Table 7
Unit Root Tests

	ADF-GLS test statistic	Ng-Perron test statistic
Output Gap	-4.170 (***)	-5.117 (***)
Inflation	-2.413 (**)	-2.546 (**)
Colombian Country Risk	-1.241	-1.304
Real Interest Rate	-1.620 (*)	-1.664 (*)
Real Exchange Rate Gap	-1.743 (*)	-9.615 (***)

(***) Denotes significance at the 1% level, (**) significance at the 5% level and (*) significance at the 10% level. Lag selection based on Schwartz information criterion. A constant was included in all tests. Once again, we used the Perron testing strategy.

Table 8

Variance Decompositions

Percentage of Variance Associated with EMBI Shocks

Horizon	Interest Rate	Country Risk	Real Exchange Rate
1	0.00	14.31	0.70
2	3.32	23.24	2.66
5	7.19	20.43	7.83
10	4.80	20.12	5.81

Table 9

Error-Correction Model for the Chilean Interest Rate (1991-2000)

Variable	Coefficient [P-values]	R-squared	Q-Statistic at lag 3	Normality
C	1.379 [0.00]	0.560	2.507 [0.43]	71.18 [0.00]
Δi_{t-1}	0.439 [0.00]			
Δi_t^*	0.083 [0.00]			
ΔURR_{t-1}	-0.075 [0.01]			
ΔURR_{t-2}	0.163 [0.00]			
ΔURR_{t-3}	-0.053 [0.10]			
Error-Correction-Term	-0.306 [0.00]			

Q-statistic refers to the test for serial correlation of the residuals of the model. Normality is the Jarque-Bera test for the normality of the residuals in the regression. P-values for all the test statistics are presented in brackets.

Table 10

Error-Correction Model for the Colombian Interest Rate

(1991-2000)

Variable	Coefficient [P-values]	R-squared	Q-Statistic at lag 3	Normality
C	0.576 [0.13]	0.331	0.877 [0.83]	8.67 [0.01]
Δi_{t-1}	0.352 [0.00]			
Δi_{t-2}	-0.010 [0.90]			
Δi_{t-3}	0.231 [0.01]			
Δi_t^*	-0.050 [0.85]			
ΔURR_{t-1}	-0.032 [0.31]			
ΔURR_{t-2}	-0.000 [0.98]			
ΔURR_{t-3}	0.053 [0.09]			
ΔURR_{t-4}	-0.001 [0.94]			
ΔURR_{t-5}	-0.073 [0.03]			
Error-Correction-Term	-0.145 [0.00]			

Q-statistic refers to the test for serial correlation of the residuals of the model. Normality is the Jarque-Bera test for the normality of the residuals in the regression. P-values for all the test statistics are presented in brackets.

Annex

Unit Root Tests for Residuals in Colombian VAR

	ADF-GLS	Ng-Perron
EMBI Equation	-8.134 (***)	-5.090 (***)
Country Risk Equation	-2.888 (***)	-2.174 (**)
Output Gap Equation	-2.549 (**)	-1.958 (*)
Inflation Gap	-5.325 (***)	-4.066 (***)
Real Interest Rate	-10.496 (***)	-5.267 (***)
Real Exchange Rate	-3.339 (***)	-2.699 (***)

(***) Denotes significance at the 1% level (**) Denotes significance at 5% level and (*) significance at 10% level. Lag selection based on Schwartz information criterion.