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Capital controls and spillover effects: Evidence from Latin-American countries*

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Résumé

L'augmentation des flux de capitaux vers les pays émergents après 2009 a ravivé le débat sur les contrôles de capitaux. Ce papier analyse les conséquences internationales de ces contrôles. Nous utilisons les statistiques de balance des paiements ainsi que des données à plus haute fréquence sur les flux de portefeuille en actions et obligations pour un échantillon de pays latino-américains, afin d'étudier les effets de spillover des flux de capitaux sur les pays voisins lors de l'introduction de restrictions à la mobilité des capitaux dans un pays donné. Notre étude économétrique montre que la hausse de la taxe imposée par le Brésil sur les achats de titres obligataires par les non-résidents a entraîné une augmentation significative des flux de portefeuille en obligations et en actions vers les autres pays d'Amérique latine. Cette augmentation est généralement de courte durée et suivie d'une baisse rapide des flux. Elle est toutefois importante. Nos estimations suggèrent que la hausse de la taxe brésilienne sur les flux de portefeuille en obligations pourrait expliquer la totalité de l'augmentation des flux obligataires entrant au Mexique entre septembre et octobre 2010.

Codes JEL : F32, F33, F42.

Mots clés: flux de capitaux, contrôles des capitaux, effets de spillover, Amérique latine, VAR.

Abstract

The surge in capital inflows towards emerging countries after 2009 has revived the debate about capital controls. This paper analyzes some of the international implications of restrictions on capital inflows. Focusing on a sample of Latin-American countries, we use detailed balance of payments data and higher frequency data on portfolio bond and equity flows to investigate the potential spillover effects that capital controls imposed in one country may have on neighboring economies. Using various econometric approaches, we find that a rise in the Brazilian tax on portfolio bond inflows has been affecting other Latin-American economies through significant surges in portfolio funds invested either in fixed income or equity securities. The effect is usually short lasting and followed by rapid reductions in those inflows. Yet it can be large. According to our estimates, the increase in the Brazilian tax on portfolio bond inflows may account for the entire surge in bond inflows to Mexico between September and October 2010.

JEL Classification Codes: F32, F33, F42.

Keywords: capital flows, capital controls, spillovers, Latin America, VAR.

1 Introduction

The surge in capital inflows towards emerging countries after 2009 has revived the debate about capital controls.[1] There are two key issues: (i) are capital controls effective at reducing the volatility of international capital flows, e.g. by decreasing the volume of inflows or preventing sudden capital outflows? and (ii) what are the effects of such controls on other countries? The second question lied at the core of the discussions that led to the adoption of "coherent conclusions for the management of capital flows" by G20 Leaders in November 2011.[2] As a matter of fact, if the introduction of capital controls in one country has positive or negative spillover effects on other countries, there is a strong motivation for multilateral cooperation to maximize global welfare.

Capital controls can be defined as measures aimed at restricting international capital mobility that discriminate between residents and non-residents. This definition does not distinguish between controls on outflows and controls on inflows, nor does it reflect the variety of possible measures from market-based restrictions or price controls to quantitative controls. The focus of this paper will be on controls on capital inflows.

There are two main reasons for which countries may impose controls on capital inflows. First, capital controls may be motivated by prudential considerations (Korinek, 2011). Countries may want to limit capital inflows to prevent the build-up of asset price bubbles and excessive external indebtedness. As shown by Korinek (2010), the risks posed by capital inflows stem from the existence of a pecuniary externality that results in distortions of the financing and investment decisions of private market participants. Small private agents take prices, especially exchange rates, as given, and neglect the price effects of their actions and the resulting balance sheets effects. In bad times, those effects may constrain the access of economic agents to external finance, which in turn forces them to cut back on their spending and contract aggregate demand following a financial amplification mechanism. Prudential capital controls can thus help to reduce the incentive for excess risk-taking on the part of private agents and the level of financial fragility in the economy. Second, capital controls may be used for mercantilist reasons to prevent an appreciation of the exchange rate while keeping the autonomy of monetary policy.

The existing literature provides mixed evidence of the effectiveness of controls to af-

[1] For a recent study including a large survey of the literature, see e.g. Magud, Reinhart and Rogoff (2011).

[2] G20 Leaders Summit, Final Communiqué, Cannes, November 2011.

fect inflows. Empirical studies suggest that capital controls have been more successful at altering the composition of flows entering a given country than at reducing their volume (De Gregorio, Edwards and Valdés, 2000). Very few studies however look at the international spillover effects of such controls.

Controls on capital inflows may have three types of spillover effects. First, the adoption of controls in one country may produce higher capital flow volatility in other countries with similar characteristics. If international capital flows are mainly driven by exogenous "push" factors, they will go where they are allowed to. Thus capital controls could in turn act as another "push" factor driving inflows in other countries. Second, capital controls that lead to persistently undervalued exchange rates, do produce externalities insofar as they affect the relative price-competitiveness of countries in international trade. Third, capital controls and restrictions to capital mobility may prevent an optimal international allocation of capital resulting in lower global economic growth.

This paper provides a first attempt to assess the magnitude of the first effect. Using detailed balance of payments data and higher frequency data on portfolio flows for a large sample of emerging countries, we construct correlation matrices of inflows in emerging economies to identify groups of countries among which spillover effects from capital controls might be the largest. We show that cross-country correlations of inflows are stronger within the same regional area and increase in crisis times.

Focusing on Latin-American countries, we look for significant divergences in the comovements of inflows following the introduction of capital controls in some countries. For our econometric analysis, we rely on monthly data on portfolio investments in bonds and equities compiled by EPFR (Emerging Portfolio Fund Research) calibrated and fitted on balance of payments data. Using single equation regressions, we provide evidence of the extent to which the Brazilian tax on portfolio inflows or IOF may have contributed to divert capital flows to other Latin American economies. Those spillover effects are significant. Using impulse response functions from VARs, we estimate that the increase in the Brazilian tax on portfolio bond inflows from 2 to 6% in October 2010 led to additional bond inflows to Mexico of about USD 1.8bn in the same month. This figure is consistent with monthly data on bond inflows to Mexico, which increased from USD 3.7bn in September 2010 to USD 5.1bn in October (EPFR data calibrated and fitted on balance of payments flows), while at the same time bond inflows to Brazil dropped from USD 4.2bn to 2.2bn. Thus, according to our estimation, in the absence of any change in the Brazilian tax on inflows, the inflows to Mexico would have slightly decreased.

The rest of the paper is organized as follows. Section 2 presents some stylized facts on international capital flows and provides anecdotal evidence of international spillover effects from Brazil's tightening of capital controls. The econometric analysis is the focus of Section 3. Section 4 concludes.

2 Stylized facts on international capital flows

2.1 EPFR and balance of payments data

In most emerging economies (Brazil is an exception), balance of payments data on portfolio liabilities, with a breakdown between debt and equity flows, are only available on a quarterly frequency. In addition, we therefore use data on portfolio investments in bonds and equities, compiled by EPFR (Emerging Portfolio Fund Research). Our dataset is comprised of bond and equity country flows data at monthly frequency, over the periods April 2004-June 2011 for bond flows and February 1996-June 2011 for equity flows. To our knowledge, those data have been used by few papers so far (Bernanke, 2010; Jotikasthira, Lundblad and Tarun, 2010; Fratzscher, 2011).

EPFR collects data from investment funds, mostly based in the OECD, on their international transactions in bonds and equities. Accordingly, the reported flows are part of the portfolio investments carried out by non-residents in emerging countries recorded in those countries' balances of payments. Yet the comparison between EPFR and balance of payments data shows a significant discrepancy in coverage, as EPFR flows only account for some 15% of balance of payments flows. To match monthly EPFR data with quarterly balance of payments flows and limit biases related to measurement errors and limited coverage, we implement a calibrating and fitting procedure, similar to the one used to compute French quarterly national accounts (see Appendix A for details).

2.2 Co-movements within and across regions

Figure 3 represents bilateral correlation coefficients of bond flows in first difference, computed from raw EPFR data, for all countries in the sample with colored squares depending on the sign and strength of the correlation. Only coefficients that are significant at the 5% threshold are represented with a colored square.

Overall EPFR bond flows look highly correlated across countries, suggesting a large role for common international drivers of portfolio flows. This result holds for all sample periods (before and after the 2007-2009 crisis) but is stronger for the crisis period defined as July 2007-March 2009 where bilateral correlations are the highest. It is striking that correlations look stronger among emerging countries than between emerging and advanced economies, as well as within certain geographical areas (diagonal blocks). The last observation is valid for both bond and equity flows (Figure 4).

Focusing on flows to emerging economies, we note that such flows were very dynamic, especially towards emerging Europe, until 2007. The global financial crisis triggered sharp and simultaneous reversals in capital flows in late 2008 and early 2009 (Figure 1). From March 2009 onwards, capital flows rebounded, mostly towards emerging Asia and Latin America. They reached on average more than 3 percent of recipient countries' GDP, with portfolio flows accounting for nearly half of total flows (IMF, 2011).

While we do not provide an assessment of the drivers of this dynamics, we guess that "pull" factors, such as better growth outlook, higher interest rates, lower public and private debt in emerging countries, did play a role in this rebound. Using a factor model, Fratzscher (2011) emphasizes the role of idiosyncratic, country-specific shocks as a dominant determinant of capital flows, particularly for countries in Emerging Asia and Latin America. Yet "push" factors such as abundant global liquidity, resulting from extremely accommodative monetary policies in advanced economies, and increased uncertainty about growth prospects in those countries may also be playing a role. In that respect, it is illustrative that correlations of flows have increased between the pre- and post-crisis periods (Figure 5).

The potential importance of "push" factors is a strong motivation to investigate the existence of possible spillover effects of capital controls. If capital flows are driven by global factors unrelated to domestic circumstances, one can imagine that restrictions to entry in a given country may lead to an increase in inflows to neighboring countries, as we have seen that flows seem very strongly correlated within the same region.

2.3 Anecdotal evidence of spillover effects of controls

Faced with volatile and short-term inflows, some emerging countries have been reacting by imposing controls on inflows. In Latin America, it has been especially the case of Brazil, which reinstated the IOF (*Imposto sobre Operações Financeiras*) in October 2009, at a 2% rate on portfolio inflows from non-residents (either equities or bonds). The tax rate was subsequently raised to 6% for bond inflows in October 2010 (see Appendix B).

Figure 6 looks at changes in bilateral correlations of flows between Brazil and seven other Latin-American countries over four periods, depending on whether the IOF was in place or not. The comparison is far from conclusive. With the exception of Argentina, Mexico and Venezuela, bond flows tend to be more correlated in the last period (2009-2011) when the IOF was in place than immediately after the IOF was removed in October 2008 (The IOF was first introduced on bond purchases by non-residents in March 2008 to limit the surge in inflows).

However there is some anecdotal evidence of spillover effect. In an article from *Valor Econômico*,[3] the authors relate that Japanese investors, traditionally important buyers of Brazilian bonds, were increasingly focusing on Mexico at the expense of Brazil, especially because of the higher uncertainty created by the IOF. According to the same article, Tandem Partners, an investment fund dedicated to emerging economies, cancelled its position on Brazilian securities while increasing the share of Mexican securities in its portfolio. The contrasted evolution of CDS spreads (rising in Brazil, decreasing in Mexico) may also be interpreted as another clue of such a shift in investors' portfolios.

Brazilian balance-of-payment data show some decrease in inflows since the reinforcement of the IOF in October 2010: while inflows from non-residents into fixed-income securities reached USD 10.6 billion in the third quarter of 2010, they fell to USD 3.6 billion in the second quarter of 2011. At the same time, inflows from non-residents into Mexican fixed-income securities increased dramatically, from USD 4.9 billion in the second quarter of 2010 to USD 14.4 billion in the first quarter of 2011. Figure 2 illustrates those diverging trends.

Other factors than a diverting effect of capital controls may however explain the diverging trends of non-resident capital inflows into bonds, in Mexico and in Brazil. On the one hand, Mexico was the first emerging country to be included in Citigroup Inc.'s World Government Bond Index, on October 1st, 2010, which may have boosted fixed-income investment into the country. On the other hand, the decrease in bond inflows to Brazil may be partly explained by circumventing strategies: shortly after the IOF reinforcement in October 2010, foreign direct investment into Brazil increased significantly, especially through intercompany loans.[4] As noted by Carvalho and Garcia (2006), such a strategy had already been used in the past to avoid the IOF tax. Moreover, subsequent measures adopted by the Brazilian authorities, taxing foreign exchange derivatives transactions,

[3] "México rivaliza com Brasil por capitais", May 3rd 2011.

[4] "Empréstimo intercompanhia cresce para driblar IOF", *Valor Econômico*, June 28th 2011.

may indicate that investors had been pursuing carry-trade strategies through the forward exchange rate market rather than through the Brazilian bond market. Hence the need for a more thorough and systematic analysis.

3 Econometric evidence

We investigate if a tightening of controls in Brazil has been diverting short term flows to third economies in Latin America. We look at the largest recipient countries of capital flows in the region other than Brazil: Argentina, Chile, Colombia, Mexico and Peru. We use first a static econometric approach (time series and panel estimations, in which the lagged dependent variable is not included as a regressor) then we introduce some dynamics by estimating a vector autoregressive model (VAR) and we simulate impulse-response functions country by country.

3.1 Data and variables

We use monthly EPFR data on portfolio flows calibrated and fitted on quarterly balance of payments data. The sample period (2004m4-2011m6) was determined by the availability of EPFR series for both bond and equity funds. To get comparable estimators, we relate bond and equity inflows to domestic GDP (*Bd_r*, *Eq_r*). Quarterly GDP data have been seasonally adjusted using the census X11 (Historical) method, and then converted into monthly series by linear interpolation. Data on the remaining variables come from the IMF International Financial Statistics and from national sources.

We focus on the effect of recent Brazilian capital controls on portfolio inflows to Brazil and to third Latin American countries. The static specifications are estimated for three dependent variables in the case of Brazil: the ratios of gross inflows of portfolio bonds (*Bra_Bd_r*), of portfolio equity (*Bra_Eq_r*) and of intercompany loans (*Bra_ICL_r*). For third countries, estimations are carried out on two dependent variables: the ratios of gross inflows of portfolio bonds (*_Bd_r*) and of portfolio equity (*_Eq_r*).

As for the regressors, our main explanatory variable is the prevailing value at the end of the month of the IOF on bonds (*IOF_Bd*). The IOF on bonds is an ad valorem tax on purchases of Brazilian fixed income securities by non-residents, which has been ranging from 0% to 6% since March 2008 (see Appendix B). Hereafter we refer to this tax simply as “IOF”.

The other explanatory (control) variables correspond to "push" and "pull" factors commonly highlighted in the literature on capital flows. Domestic growth is proxied by the (seasonally adjusted) monthly growth rate of the industrial production index (*InProd_v*). We compute a tax equivalent measure of capital controls (*_TaxEquiv*) for third countries that have implemented required reserves on some categories of external financing (see Appendix C). This proxy for capital controls varies over time for Colombia and Peru and is equal to zero for the other countries. As a proxy for the world interest rate (*WIR*), we calculated an average of the money market rates in the main reserve currencies areas (U.S.A., the Eurozone, Japan, the United Kingdom and Switzerland), weighted by their respective GDP in 2010. As a measure for domestic interest rates (*IR*), we use nominal interbank money market rates.[5] Besides, we construct a measure of "pure" expected depreciation (*EER_v*), aimed at avoiding colinearity with other explanatory variables. The variable *EER_v* captures future expected exchange rate variations in a given country, once the effect of domestic and foreign interest rates is removed from the observed three-month forward exchange rates.[6] Combined, the latter three variables capture the excess return that a foreign investor can get by investing in domestic riskless assets, corrected by the expected exchange rate depreciation (*IR* - *EER_v* - *WIR*). Finally, using the cyclical component of the volatility index VIX[7], we construct a dummy variable (*VIX_extreme*) aimed at capturing periods of extreme uncertainty, during which a widespread retrenchment of financial flows may occur. The variable *VIX_extreme* is equal to one when the absolute value of the HP-detrended VIX is larger than two times its standard deviation, and zero otherwise.

Other variables are tested in alternative specifications (see below). As long as they reflect fears of excessive currency appreciation, they may be used as instruments for the Brazilian *IOF_Bd*. Purchases of foreign currency by Brazilian authorities are proxied by the changes in the ratio of official reserve assets to GDP (*IRes_r*).[8] The realized appre-

[5]For the sake of homogeneity across countries we chose interbank rates of very short maturities rather than three-month rates.

[6]Using the usual notations, the covered interest rate parity condition yields $1 + i = f/s(1 + i^*)$. In practice, a measure of expected depreciation such as $(f - s)/s$ is strongly correlated with both i and i^*. To avoid colinearity problems, we use as a proxy for the "pure" expected depreciation *EER_v* the residuals $\hat{\varepsilon}$ of the following OLS regression (with no constant): $(f - s)/s = \alpha i + \beta i^* + \varepsilon$

[7]The trend was obtained using a Hodrick-Prescott filter and subtracted from the series to obtain the cyclical component of the VIX.

[8]However the quality of such a proxy is biased by the keenness of the Brazilian central bank to intervene through foreign exchange swap contracts.

ciation rate is calculated as the percentage variation of the spot exchange rate (SER_v). The domestic inflation rate (INF) is computed as the percentage monthly variation of the Consumer Price Index, on a year-on-year basis. We also checked the suitability of the IOF on equity (IOF_Eq) either as an additional control or else as an instrument. IOF_Eq may be assimilated to a "dummy" exogenous variable since it shows almost no variability: it was raised from 0% to 2% around October 2009, and has remained at this level since then. Some descriptive statistics of the variables are provided in Table 1.

3.2 Static analysis

We estimate four types of equation, each one explaining a different variable: portfolio bond or equity inflows, to Brazil or to third countries. Our baseline regression is carried out using ordinary least squares (OLS), and subsequent estimations include fixed effects (FE) or instrumental variables. For our dependent variables, gross inflows/GDP, all the series are stationary according to augmented Dickey-Fuller (ADF) tests at least at 95% confidence levels. As for our explanatory variables, with very rare exceptions, series are found to have a unit root. We therefore use the ratio of flows to GDP as such and all the regressors in first differences (denoted by $d(.)$) or in percentage variations (denoted by $\_v$).[9] As an additional test, each regression is also estimated by OLS with the dependent variable in first differences. Finally, along with portfolio bond and equity inflows towards Brazil we estimate the effect of the IOF on intercompany loans from foreign corporations to Brazilian affiliates (measured as a ratio to GDP). A positive reaction of the latter to the IOF might reflect some by-passing of Brazilian capital controls by foreign investors.

In our baseline specification, we use contemporaneous and lagged values (up to four lags) of $d(IOF\_Bd)$, which is the focus of our analysis. This variable is not autocorrelated, so we can include in the same regression different lags of $d(IOF\_Bd)$. The rest of the explanatory variables are included in a contemporaneous way.

Taking for instance bond inflows as the dependent variable, the time series specification for Brazil takes the form:

$$Bra\_Bd\_r_t = c + \sum_{l=0}^{4} \alpha_l d(IOF\_Bd)_{t-l} + \sum_{k=1}^{K} \beta_k X_{k,t} + \varepsilon_t \tag{1}$$

where $X_1, ..., X_K$ is a set of K control variables and ε_t are supposed to be zero mean and

[9] This specification is quite common in previous work on the drivers of capital flows (see e.g. Cardoso and Goldfajn, 1998; De Gregorio et al., 2000; De Vita and Khine, 2008).

constant variance errors. The specification for bonds includes an AR(1) term, whereas AR(1), AR(2) terms are added for equity and intercompany loans.

For third countries, the estimations are carried out in panel to check for the homogeneity of the responses of inflows to a change in the IOF. The pooled specification can be written as:

$$Bd\_r_{i,t} = c + \sum_{l=0}^{4} \delta_l d(IOF\_Bd)_{t-l} + \sum_{k=1}^{J} \eta_k X_{k,t} + \sum_{k=J+1}^{K} \gamma_k X_{k,i,t} + \varepsilon_{i,t} \qquad (2)$$

where $X_1, ..., X_J$ are control variables common to all countries, $X_{J+1,i}, ..., X_{K,i}$ are country specific control variables, and $\varepsilon_{i,t}$ are assumed to be zero mean and constant variance errors. All the specifications in panel include AR(1) and AR(2) terms.

By contrast with the pooled equation, in the FE specification, the constant c_i is allowed to vary across individuals. This aims at capturing structural country-specific effects that could have remained undetected in the pooled regression, i.e. embedded in the error term.

Both equations (1) and (2) are also estimated using instrumental variables. Indeed, OLS estimations in which incoming short term flows appear positively related to capital controls in the current period $d(IOF\_Bd_t)$ might suffer from an endogeneity bias: Brazilian authorities may react to an observed surge in inflows, either to Brazil or to neighboring countries, by raising the IOF. Related work documents this phenomenon as capital controls were set up in Chile and Brazil in the 1990s (De Gregorio et al., 2000; Cardoso and Goldfajn, 1998). To address this issue we estimate the same equation using two stage least squares (TSLS), then we check the suitability of the instruments and the exogeneity of the IOF. Only $d(IOF\_Bd)_t$ is suspected of being endogenous, as its lagged values $d(IOF\_Bd)_{t-p}$ are predetermined and thus exogenous.

Consider equation (2) for third countries. If $d(IOF\_Bd)_t$ is simultaneously determined along with $Bd\_r_{i,t}$, then $E(d(IOF\_Bd)_t \varepsilon_{i,t}) \neq 0$, so that the OLS estimators are biased. To overcome this problem we chose the IOF on equity $d(IOF\_Eq)_t$ and the previous month observed appreciation SER_v_{t-1} as exogenous instruments (denoted by Z) to explain $d(IOF\_Bd)_t$.[10] At the first stage of TSLS, $d(IOF\_Bd)_t$ is regressed on the exogenous

[10] We tried current and lagged values of other variables as instruments for $d(IOF\_Bd)_t$. Neither the ratio of international reserves to GDP ($IRes\_r$), nor the inflation rate (INF) (both in first differences), nor lags of SER_v higher than one appeared to explain $d(IOF\_Bd)_t$. Along with SER_v_{t-1}, the IOF on equity $d(IOF\_Eq)_t$ was chosen as an instrument rather than as a regressor, since it is collinear with $d(IOF\_Bd)_t$. Moreover, according to the Hansen-Sargan test, $d(IOF\_Eq)_t$ is not correlated with the error $\varepsilon_{i,t}$ in the baseline equations, except in the case of bond inflows to third countries.

instruments Z and on all exogenous regressors from equation (2). This auxiliary OLS regression yields an instrumented variable $d(\widehat{IOF\_Bd})_t$ and first stage residuals $\nu_{i,t}$. As long as the variables Z are uncorrelated with the error $\varepsilon_{i,t}$ of the main regression, they constitute suitable instruments, so that $d(\widehat{IOF\_Bd})_t$ is no more simultaneously determined along with $Bd\_r_{i,t}$. According to the Hansen-Sargan test, one cannot reject that $E(Z\varepsilon_{i,t}) = 0$ as the the p-value of the J-statistic (reported in Tables 2 and 3) is larger than 0.1. In turn, to check the endogeneity of the suspected variable $d(IOF\_Bd)_t$, we included the first stage residuals $\nu_{i,t}$ as an additional regressor in the original regression. Conditional to the suitability of the instruments Z, one can accept that $d(IOF\_Bd)_t$ was simultaneously determined along with $Bd\_r_{i,t}$ if $\nu_{i,t}$ appears to significantly explain $Bd\_r_{i,t}$ in equation (2).

The main results are summarized in Tables 2 and 3. Measured by the adjusted R^2, the fit of the regressions varies noticeably across specifications. Durbin-Watson statistics are all close to 2, which suggests that the autocorrelation of errors is corrected by the inclusion of AR terms.[11] Therefore, there is no risk of spurious regressions.

In the case of Brazil (Table 2), the OLS regression with the dependent variable in level yields no clear cut conclusions about the effectiveness of controls. A tightening of the IOF significantly reduces the inward flow of portfolio bonds only with one month lag. Moreover, between two and three months after a given tightening of *IOF_Bd*, Brazil seems to experience a significant increase in bond inflows. The latter effect is robust to TSLS estimation. On the one hand, this counterintuitive result would be in line with the rationale advanced by Cordella (2003): foreign investors are likely to prefer countries imposing capital controls, as long as the recipient economy is expected to reduce its external vulnerability. On the other hand, this increase appears to be contemporaneous to a surge in bond inflows to other Latin America countries (see Table 3). Hence, the significant surge in portfolio bond inflows between two and three months after a tightening of the IOF might be due to some type of rebalancing strategy of portfolio funds, missed by our control variables.[12]

[11]Since AR terms are added to account for the serial autocorrelation of errors, the DW statistic is computed from the estimated one-period ahead forecast errors, rather than from the unconditional residuals. The DW statistic remains a valid indicator as long as the specification does not include lagged values of the dependent variable on the right hand side.

[12]For the whole sample, the fit of the regression on the level of bond inflows (in GDP %) to Brazil is quite poor. This is mainly due to the fact that portfolio bond inflows to Brazil were quite volatile until 2007, whereas the IOF began to vary only from mid 2008. Including the dependent variable with one lag (instead of the AR term) on the right hand of the equation yields a non-significant coefficient and does

Turning to short-term effects, the negative impact of IOF on bond inflows to Brazil found in $t+1$ is not robust to the TSLS estimation. Yet we cannot reject that $d(IOF\_Bd)_t$ is exogenous in this specification. As endogeneity bias is ruled out and OLS estimators are more efficient than TSLS, we trust the initial result: the IOF is negatively related to the level of bond inflows to Brazil (relative to GDP) in $t+1$. The results with the dependent variable in first difference confirm a sharp slowdown in the growth of foreign investments in portfolio bonds shortly after a rise in the IOF. This slowdown is statistically significant at times t, $t+1$ and $t+4$. Thus if a tightening of capital controls does not imply an outright drop in portfolio investments in bonds, it has an effect on the pace of entry of those flows.

Quite surprisingly, equity inflows to Brazil are better explained by the bond tax and the rest of regressors than bond inflows themselves. In period t, as soon as the IOF is raised, Brazil experiences significant and large inflows of equity. In this case, there is some evidence that $d(IOF\_Bd)_t$ is simultaneously determined with equity inflows to Brazil. As the instruments appear to be valid, the estimation by TSLS is preferred to OLS. Still, OLS results are robust to TSLS. This result is consistent with previous empirical work showing that capital controls tend to be effective at altering the composition of inflows if not their magnitude. However this might also reveal circumventing strategies. There is some evidence that the existence of loopholes enabled investors to by-pass capital controls in the past in Brazil (Garcia and Barcinski, 1998) and Chile (De Gregorio et al., 2000). This may have been the case again in October 2009 when conversion of ADRs into Brazilian equities was used to avoid the IOF tax; those transactions were subsequently taxed at 1.5% from November 2009 onwards to close this loophole (see Appendix B).[13] Similarly, we also find some support for the view that foreign investors may have been using intercompany loans (from foreign corporations to their Brazilian affiliates), which are recorded as FDI, as another way to by-pass capital controls: the coefficient on $d(IOF\_Bd)_{t-4}$ in the last column of Table 2 is positive and significant. The four-month lag suggests that these strategies take some time to be implemented. The Brazilian authorities eventually adressed this loophole in August 2011.

not change the results. However, as the regression is run for the period 2007m7-2011m6, the adjusted R^2 increases significantly. We chose to work with the largest possible data sample rather than by subperiods, to keep as many degrees of freedom as possible.

[13]Part of the strong effect we find may also be related to Petrobras' large equity issue in September 2010, since some of the corresponding equity flows may have been reported in October, at a time when the IOF was being reinforced.

As for the control variables, the coefficients on the return differential corrected by depreciation are not statistically different from zero. Indeed expected returns may be more relevant for investors to discriminate among countries at any given point in time, than to determine the pattern of capital flows to a given country over time.[14] In turn, the proxies for growth and for extreme uncertainty do significantly affect portfolio inflows towards Brazil, especially in the case of bond inflows. Still, the exclusion of those controls, while worsening the fit of the regressions, does not change our results.

Next, we investigate the existence of potential spillovers of the Brazilian tax on portfolio bond inflows towards the five Latin American economies mentioned above (Table 3). The specification for third countries presents almost no difference with respect to that used for Brazil. We simply add one more regressor: the variable $d(\_TaxEquiv)$ controls for the relative cost for a foreign investor of holding external liabilities issued by Colombia and Peru.

By contrast to Brazil, equity inflows towards third countries seem to be better explained by return differences[15] than by uncertainty periods. In turn, only bond inflows are positively related to growth. The OLS regression points to significant coincident common responses of inflows to third countries following a tightening of capital controls in Brazil. Following an increase in the IOF for bonds in Brazil, third countries experience significant surges in portfolio bond inflows and, to a lesser extent, equity inflows. Yet, this diverting effect from the IOF on bond inflows to third countries seems to vanish after one month.[16] Regressing the dependent variable in first difference tends to confirm that the surge in the level of inflows in period t is short-lived: indeed we find a significant slowdown in the growth of incoming portfolio flows, of both bonds and equity, in period $t+1$. The spillover effects found in the pooled estimations are also robust when controlling for country fixed effects (FE).

[14] This is especially the case for Brazil, where interest rates have been on a decreasing trend over the past ten years (due to stabilization policies), without this implying a fall in capital inflows.

[15] The negative sign suggests that equity flows increase as the expected yield of alternative portfolio (bond) investments decrease.

[16] As noted above, the significant surge in flows to third countries three months after a given tightening in Brazilian capital controls (see table 3) also characterizes the evolution of bond inflows to Brazil (table 2). Thus, rather than the direct impact of the IOF on bonds, an omitted determinant of the investment funds' behavior could drive such a lagged effect. The VAR analysis sheds some light on the duration of the IOF's effects over time.

The above OLS and FE results might again be driven by simultaneity bias: Brazilian authorities could have tightened controls as they observe increases in inflows towards other Latin America economies. We therefore apply for third countries the same instrumental variables strategy (TSLS) as for inflows to Brazil. The spillover effects on portfolio bonds flowing to third countries appear to be more important and long-lasting than those estimated by OLS. Yet, while $d(IOF\_Bd)_t$ appears to be endogenous, we could not find a set of suitable instruments. Even as we instrumented solely by SER_v_{t-1}, the p-value of the J-statistic did not allow to reject that $E(Z\varepsilon_{i,t}) \neq 0$. For bond inflows to third countries, OLS results are thus preferred to TSLS. Namely, we accept that spillovers are significant but tend to last on average no more than one period. As regards equity inflows, the spillover effect is not robust to the TSLS estimation. In this case the test confirming the exogeneity of $d(IOF\_Bd)_t$ is backed by the validity of the instruments. The OLS and FE estimators are thus preferred to TSLS results. We can therefore accept the existence of some diverted equity flows to third countries, which are again very short-lived.

3.3 Dynamic specification

To investigate potential spillovers on a country basis, we estimate a vector autoregression (VAR) model for each country. This type of specification accounts for contemporaneous and lagged feedback effects among variables, all of which can be treated as endogenous. Not only potential endogeneity bias are ruled out but the VAR seems also a good approach as capital flows respond to a given shock at different lags, depending on their determinants and on the characteristics of the recipient country. We can therefore study how potential spillovers evolve over time taking into account the dynamic path of all the variables. The reduced autoregressive (AR) form of an open VAR can be written as follows:

$$Y_t = \mathbf{A}_p^*(L)Y_t + \mathbf{B}^*X_t + e_t \tag{3}$$

where[17] Y_t is a vector of m endogenous variables, $\mathbf{A}_p^*(L)$ is an invertible matrix containing the $m \times m$ coefficients $a_{kp}^j(L)$ of lagged endogenous variables. (L) denotes a lag polynomial: for p lags, the coefficient of a given lagged j variable in the equation for variable k is $a_{kp}^j(L) = (a_{k1}^j L + ... + a_{kp}^j L^p)$. X_t is a vector of n exogenous variables, with an associated $m \times n$ matrix of coefficients $\mathbf{B}^*$. e_t is a vector of m reduced form errors, with an associated

[17]We omit the constant in (3). Variables are thus written in deviations from their respective average.

$m \times m$ variance-covariance matrix $\mathbf{\Sigma}_e$. Rewriting the endogenous variables in (3) as a moving average (MA) of errors gives:

$$Y_t = \mathbf{C}_p^*(L)e_t + \mathbf{D}^* X_t \tag{4}$$

where $\mathbf{C}_p^*(L) = [\mathbf{I} - \mathbf{A}_p^*(L)]^{-1}$ and $\mathbf{D}^* = [\mathbf{I} - \mathbf{A}_p^*(L)]^{-1}\mathbf{B}^*$.

As long as the perturbations e_t are stationary, the systems (3) and (4) can be estimated by OLS since all the right hand variables are predetermined. The number of lags of each VAR(p) was chosen as the one recommended by a majority of the following information criteria: sequential LR test statistic, final prediction error, Hannan-Quinn, Akaike, and Schwarz. We retained one lag for Brazil, Chile and Mexico, and two for Argentina, Colombia and Peru. The main continuous variables from the static specification were modeled as endogenous variables, each representing one equation in the VAR: d(*IR - EER_v - WIR*), *InProd_v*, d(*IOF_Bd*), *_Eq_r*, *_Bd_r*. In addition, we included d(*_TaxEquiv*) in the VAR for Colombia and Peru. Finally, the two "dummies", *VIX_extreme* and d(*IOF_Eq*), were used as exogenous variables for every country.

We exploit the dynamic properties of the estimated VAR through the simulation of impulse-response functions. This cannot be done directly from (3) though. A structural VAR model (SVAR) is needed to get economically interpretable impulse-responses.[18] The SVAR is assumed to summarize the underlying "true" structure of the modeled economic relationships. The following structural representation takes into account the potential contemporaneous relationships between endogenous variables through a matrix $\mathbf{M}$.

$$\mathbf{M}Y_t = \mathbf{A}_p(L)Y_t + \mathbf{B}X_t + u_t \tag{5}$$

In (5), u_t is a vector of m structural perturbations and $\mathbf{\Sigma}_u$ is its associated variance-covariance matrix. $\mathbf{M}$ is assumed to be an invertible $m \times m$ matrix and the dimensions of the other matrices and vectors are the same as in (3). The relationship between a structural model (5) and its reduced form (3) is given by $\mathbf{M}$ as follows:

$$\mathbf{A}_p^* = \mathbf{M}^{-1}\mathbf{A}_p \tag{6}$$

$$\mathbf{B}^* = \mathbf{M}^{-1}\mathbf{B} \tag{7}$$

$$e_t = \mathbf{M}^{-1}u_t \tag{8}$$

$$\mathbf{\Sigma}_e = \mathbf{M}^{-1}\mathbf{\Sigma}_u\mathbf{M}^{-1\prime} \tag{9}$$

[18]See Amisano and Giannini (1997) and Gottschalk (2001) for a more complete discussion on the SVAR models.

The impulse-response functions are computed from the MA structural form of the SVAR. From (4) and (6-9) we can write:

$$Y_t = \mathbf{C}_p(L)u_t + \mathbf{D}^* X_t \tag{10}$$

where $\mathbf{C}_p(L) = \mathbf{C}_p^*(L)\mathbf{M}^{-1}$ and $u_t = \mathbf{M}e_t$.

A SVAR implies that the structural perturbations of the model u_t are not mutually correlated, unlike the reduced form errors e_t obtained above. The elements of u_t represent unexpected 'primitive' innovations, with no common causes. The impulse-response functions (i.e. the coefficients of $\mathbf{C}_p(L)$) have then an economic interpretation since the effect of a shock to u_t is computed anything else in Y_t being equal (i.e. as a lagged serie of partial derivatives). As long as innovations u_t are orthogonal, the associated variances and covariances $\mathbf{\Sigma}_u$ form a diagonal matrix.

The SVAR appears therefore well suited to the issue of capital controls in Brazil: the response function to a given variable (say, *d(IOF_Bd)*) is simulated from a shock on its structural perturbation, ($u_{d(IOF\_Bd)}$), i.e. from an unexpected variation of the policy instrument. This represents reasonably what has happened in practice, since Brazilian authorities have not so far preannounced the changes in the tax on portfolio inflows.

Yet, the unrestricted SVAR (5) is not identified. Several combinations of the coefficient matrices $\mathbf{M}$ and $\mathbf{A}_p(L)$ in (5) can yield the reduced form (3) estimated above. Thus, in the relationship implied by (10) additional identifying restrictions are usually imposed on $\mathbf{M}^{-1}$.

In our baseline identification scheme, we impose simple exclusion restrictions: we assume that $\mathbf{M}^{-1}$ is a $m \times m$ lower triangular matrix.[19] A coefficient $a_k^j = 0$ on the k^{th} row, j^{th} column of $\mathbf{M}^{-1}$ implies a recursive order in the contemporaneous causality: at time t an unexpected shock on the variable y_j (i.e. u_j) leaves unchanged the observed variation of y_k (or equivalently the forecast error e_k).

For third countries in our sample (for example, Mexico), $\mathbf{M}^{-1}$ multiplies the vector of structural shocks to endogenous variables in the following order: $u_t = (u_{d(IOF\_Bd)}, u_{d(IR\text{-}EER\_v\text{-}WIR)}, u_{InProd\_v}, u_{Mex\_Eq\_r}, u_{Mex\_Bd\_r})$. By assuming a lower triangular matrix $\mathbf{M}^{-1}$, we actually specify which variables are considered "external" or "weakly endogenous", so that they cannot be contemporaneously affected by shocks on other variables. We ordered the variables of a given country VAR model following economic reasoning, with

[19]Note that, with the lower triangular matrix used to get our baseline results, the impulse-response functions are equivalent to those yielded by the Cholesky method of decomposition of errors.

our dependent variables (portfolio inflows) always in the last position (i.e. as the "most endogenous" variables). In general, "push" variables appear before "pull" ones. We consider that the interest rate spread may have a contemporaneous effect on industrial growth in a given country, but the latter cannot affect the interest rate differential in the same period. As for the ordering between *d(IOF_Bd)*, *Mex_Eq_r* and *Mex_Bd_r*, the above TSLS estimations showed that it is hard to disentangle the sense of causality between the IOF and portfolio inflows. Yet, we consider that global investment funds can react to changes in capital account regulations in a given country faster than the country's authorities can respond to changes in incoming capital flows. First, local authorities face lags when collecting and processing information on capital flows; second, changes in the rules take time to be implemented. Thus, we assume that Brazilian capital controls on bonds may explain (but not be explained by) flows to third countries (*Mex_Bd_r*) within the current period (a month length). Similarly, for third countries where capital controls have been set (Colombia and Peru) the tax equivalents for required reserves *d*(*Col_TaxEquiv*) and *d*(*Per_TaxEquiv*) are ordered just before the inflows. As for the order of flows, we assume that equity inflows to a given country (*Mex_Eq_r*) are more stable than portfolio bond inflows (*Mex_Bd_r*), which are more volatile (see Table 1) and thus supposed to be affected by all the precedent variables. The VAR for Brazil was specified following the same order, except that *d*(*IOF_Bd*) is a less "exogenous" variable than in third countries: it follows *InProd_v* and precedes *Bra_Eq_r*. As in third countries, we suppose that capital inflows are quite reactive to the IOF, so that portfolio decisions by fund managers can be modified within a month. By contrast, Brazilian authorities react at least with one month lag after having observed a surge in inflows.

Our impulse responses are computed supposing that the magnitude of a shock on u_{jt} corresponds to the root of the $\mathbf{\Sigma}_u$ diagonal elements (i.e. to one standard deviation σ_{uj}). Figure 8 focuses on the responses of portfolio bond and equity flows to a one standard deviation shock to the IOF (equivalent to a 53 basis points increase in the tax).[20] As noted by Hamilton (1994), the fact that the confidence intervals tend to be quite wide is a

[20] Although we are mainly interested in potential diversion of capital inflows, we also simulated the response of nominal exchange rates to capital controls. We found significant evidence of currency appreciation only in Argentina, following the implementation of the IOF in Brazil. Still, spillover effects on exchange rates are difficult to show since many variables other than portfolio inflows influence exchange rates. A study focused on exchange rates should also control for the whole set of net operations denominated in foreign currency, including sterilization by central banks. We leave that for future research.

common feature of the simulation of VAR impulse responses. Related works often display 68% confidence intervals. In our work, the confidence threshold is more stringent. Figure 8 shows ± 2 standard deviation intervals, i.e. a 95% confidence level. Still, one way to increase the statistical significance of the responses would be to impose more restrictions on the SVAR. In our baseline model, we chose to impose only the exclusion restrictions summarized in the triangular matrix $\mathbf{M}^{-1}$, since they already imply strong assumptions on the contemporaneous causality of the variables.

Most results yielded by the VAR are consistent with those found in the static approach. The VAR impulse-response functions shed some light on the timing of effects, while determining which countries of the sample are more likely to receive diverted flows after a rise in the IOF. Since variables are stationary, the response to an innovation reverts back to the equilibrium level in the subsequent periods.

For Chile, Mexico and Peru we find significant evidence of a boost in bond inflows immediately after a tightening of the IOF. In particular, we estimate that the increase in the Brazilian tax on portfolio bond inflows from 2 to 6% in October 2010 may have triggered additional bond inflows to Mexico of about USD 1.8bn in the same month. Other countries also experience a surge in inflows, although the statistical significance is lower, especially after one month. Surges tend therefore to be short-lived and are often followed by inflows temporarily below the stationary level.

The response of equity inflows to a positive innovation in *IOF_Bd* is strong and significantly positive in Brazil and Colombia and, to a lesser extent, in Mexico. The effect on portfolio equity flows is also short-lasting and generally vanishes two months after the tightening in capital controls.

4 Conclusion

This paper analyzed how portfolio inflows, both towards Brazil and towards other large Latin American countries, have responded to capital controls recently set by the Brazilian authorities. We focused on the impact of the IOF for bonds, which has been the instrument most actively used to tax portfolio financial inflows to Brazil. We found some evidence that bond inflows to Brazil tend to slow down as capital controls on fixed income securities are tightened. We found much stronger econometric evidence that a tightening of the tax on bonds has encouraged equity inflows to Brazil and some type of inward FDI such as inter-company loans, probably aimed at circumventing capital controls.

The main contribution of this paper concerns the potential international spillovers of such measures. Unlike previous studies, usually restrained to the effects on the country that tightens the access of foreign financial flows itself, we enlarged the analysis to the effect of capital controls on third countries in the region. Besides, the high frequency of our data on portfolio inflows (monthly) enables the identification of effects that could have gone unnoticed otherwise. We found significant evidence of spillovers arising from Brazilian controls on bond inflows. Our panel and VAR estimations showed that bond inflows and, to a lesser extent, equity inflows to most of the economies of our sample (Argentina, Chile, Colombia, Mexico and Peru) are positively related to a rise in the IOF in Brazil. The surge in inflows tends to be short lived, but the evidence of such an externality deserves to be highlighted, because of the potential effects on domestic macroeconomic and prudential policies pursued by neighboring countries.

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A Calibrating and fitting method

While EPFR data are available at a high frequency, they suffer from various biases, as well as insufficient coverage. On the contrary, balance of payment data are exhaustive and computed according to a standardized methodology. For many countries they are however available only at a quarterly frequency, which may undermine the robustness of certain econometric estimations (Habermeier, Kokenyne and Baba, 2011). We therefore use a method to adjust the high frequency EPFR data and ensure their consistency with quarterly balance-of-payments data. This method, called calibrating and fitting, is being used extensively to produce French quarterly national accounts.

Let BOP_q be quarterly balance-of-payment flows (either bond or equity flows), and $EPFR_m$ denote the corresponding monthly EPFR data. Then $EPFR_q$ is defined as the sum of monthly flows over a quarter: For each quarter q,

$$EPFR_q = \sum_m EPFR_m \tag{A-1}$$

We assume that although EPFR data suffer from various biases, they have a quarterly profile similar to that of BOP data (a reasonable assumption, confirmed by the visual observation of the time series' plots or by regression results, especially R^2 statistics). Then we estimate the following econometric relationship (on a quarterly basis):

$$BOP_q = \alpha + \beta EPFR_q + \varepsilon \tag{A-2}$$

This is the calibration part of the exercise. From (A-2) we derive $\hat{\alpha}$ and $\hat{\beta}$ through an OLS regression. Quarterly residuals are denoted by $\hat{\varepsilon}_q$. We then calculate monthly residuals (fitting step) through optimization techniques, so that these monthly residuals be as smooth as possible, and not distort the overall series profile. The optimization program to calculate $(\hat{\varepsilon}_m)$ is defined as follows:

$$\min \sum_m (\hat{\varepsilon}_{m+1} - \hat{\varepsilon}_m)^2 \tag{A-3}$$

subject to the constraints: $\sum_m \hat{\varepsilon}_m = \hat{\varepsilon}_q$

Finally, we assume that (A-2), which was estimated at a quarterly frequency, holds on a monthly basis as well. We can then compute monthly series, denoted by (BOP_m), as:

$$BOP_m = (\hat{\alpha}/3) + \hat{\beta} EPFR_m + \hat{\varepsilon}_m \tag{A-4}$$

From (A-4), it is easy to check that (BOP_m) has the two required properties:

1. (BOP_m) has a monthly profile similar to that of ($EPFR_m$);

2. (BOP_m) is consistent with the initial quarterly balance-of-payments series: for each quarter q,

$$\sum_m BOP_m = BOP_q \tag{A-5}$$

B Recent capital flows management measures in some Latin American countries

B.1 Brazil

Mar 08	Introduction of 1.5% IOF (ad valorem tax on financial operations) on portfolio bonds purchases by non residents.
Sep 08	The IOF is removed.
Oct 09	Reinstatement of 2% IOF on bond and equity flows from non residents.
Nov 09	1.5% ad valorem tax on the conversion of ADR (certificates of deposit issued by Brazilian corporations in the U.S.) into Brazilian stocks.
Oct 10	Increase to 6% of IOF on portfolio bonds and on the conversion of ADR. Increase from 0.38% to 6% on deposits guaranteeing non-residents' investments on exchange rate futures.
30 Dec 10	IOF on the conversion of ADR into Brazilian stocks is reduced from 6% to 2% (same level as tax on stock purchases by non-residents).
7 Jan 11	Introduction of 60% unremunerated required reserves (URR) on bank short positions in foreign currency beyond a maximum threshold (either banks' own funds or USD 3 bn).
Apr 11	The tax base of borrowings in foreign currency from the private sector is enlarged. Borrowings up to two years maturity become taxable (before, only maturities up to 3 months were taxable).
July 11	URR on the excess of bank short positions in foreign currency are strengthened: the maximum position is lowered to either 1 bn USD or the bank's own funds.
Aug 11	Introduction of 1% IOF tax on the excess of bank short positions on the foreign exchange derivatives market (over a minimum USD 10 million).
Aug 11	6% IOF tax is also applied to the inter-company loans with maturities less than two years, recorded as FDI but suspected to be used as covert inward portfolio investments.

B.2 Peru

May 06	Introduction of a 30% URR marginal rate (over the 6% minimum URR) on foreign currency bank deposits and on external bank liabilities (URR to be held in the corresponding currency).
Sept 07	The marginal URR rate on long term external bank liabilities (essentially credit lines) is removed.
Apr 08	The marginal URR rate on domestic currency bank deposits held by nonresidents is raised (from 15% to 40%) further than that on residents (from 15% to 20%).
Jul 08	The marginal URR rate on domestic currency bank deposits held by nonresidents reaches 120% (25% for residents).
Oct 08	The marginal URR rate on short term external bank liabilities is removed.
Dec 08	The marginal URR rate on domestic currency bank deposits held by nonresidents is lowered to 35% (0% for residents).
Jul 10	Increase in the commission on the sale of Central Bank securities to non-residents, from 0,01% in December 2009 to 4% in July 2010.
Jul-Sep 10	Limits on the investments of pension funds abroad are raised from 22% to 28% in July, then up to 30% in September (removing restrictions on capital outflows).
Feb-Oct 10	The marginal URR rate on short term external bank liabilities is reinstated at 35% then progressively raised to 75%.
Jul-Sep 10	The marginal URR rate on domestic currency bank deposits held by nonresidents is raised from 40% to 120% (from 0% to 15% for residents).
Feb 11	A Government bill intends to raise from 30% to 50% the limit on the investments of pension funds abroad (removing restrictions on capital outflows).

B.3 Colombia

May 07	Introduction of a 40% URR ratio (for a minimum period of 6 month deposit in domestic currency) mainly on portfolio debt inflows from non-residents. A 2 years minimum stay was required for an inflow to qualify as FDI.
May 08	The URR ratio on portfolio inflows is raised to 50%.
Oct 08	The URR ratio is removed.

C Computation of a tax equivalent to the reserve requirements

A standard tax equivalent for required reserves (RR) on foreign liabilities is used e.g. by Valdés-Prieto and Soto (1998), De Gregorio et al. (2000) and Edwards and Rigobon (2009). We compute this tax for the more general case in which required reserves can be remunerated. We follow the notation of De Gregorio et al. (2000).

Suppose that, for each dollar of incoming funds, a percentage u is to be held as RR for a minimum period h and that $(1-u)$ is invested for k periods. The world interest rate is denoted by i^*, the domestic yield is i_k and the interest rate paid on RR (if remunerated) is r. Assuming that $k \geq h$ and that the RR, once reimbursed, are reinvested outside the country that has set capital controls, the non-arbitrage condition states:

$$(1-u)(1+i_k)^k + u(1+r)^h(1+i^*)^{(k-h)} = (1+i^*)^k \tag{C-1}$$

The tax equivalent for the RR, denoted by μ_k, has to be such that $i_k = i^* + \mu_k$. Applying the latter condition and the approximation $(1+x)^a \approx (1+ax)$ one gets:

$$\mu_k = \frac{u}{(1-u)}\frac{h}{k}\left[i^* - r\left(1 + i^*(k-h)\right)\right] \tag{C-2}$$

In calculating our proxy for the tax equivalent of the RR, we assume for simplicity that the investment period equals the minimum holding period. The tax equivalent for the RR used in our empirical analysis is then:

$$\mu_k = \frac{u}{(1-u)}(i^* - r) \tag{C-3}$$

In our empirical analysis, r is nil for Colombia and 0.6*LIBOR for Peru. For i^* we use the world interest rate (WIR). The series of u are constructed from the tables in Terrier, Valdés, Tovar, Chan-Lau, Fernández-Valdovinos, García-Escribano, Medeiros, Tang, Vera Martin and Walker (2011) and Rincon and Toro (2010) for Colombia, and in Rossini, Quispez and Rodriguez (2011) for Peru.

Table 1: Descriptive statistics

	Observations	Mean	Median	Std. Dev.	Minimum	Maximum
Main variables for Brazil						
Portfolio bond inflows (% GDP)	87	0.45	0.87	2.13	-8.99	3.89
Portfolio equity inflows (% GDP)	87	0.99	0.59	1.95	-4.29	9.63
Intercompany loans to Br. Affiliates (% GDP)	87	0.41	0.34	0.68	-1.79	3.23
Domestic interest rate (%)	86	13.33	12.62	3.33	8.65	19.75
Expected FX rate depreciation (%)	86	-0.02	-0.12	0.39	-0.95	1.08
Industrial Production (index)	87	119.40	118.22	7.91	103.81	132.35
FX rate change (%)	85	-0.85	-1.33	3.40	-6.64	17.54
Inflation rate (%)	87	5.30	5.26	1.29	2.96	8.07
IOF on portfolio bond inflows	87	0.98	0.00	1.87	0.00	6.00
IOF on portfolio equity inflows	87	0.48	0.00	0.86	0.00	2.00
Variables for other Latin American countries (Argentina, Chile, Colombia, Mexico and Peru)						
Portfolio bond inflows (% GDP)	424	1.04	0.80	3.02	-8.00	11.43
Portfolio equity inflows (% GDP)	424	0.23	0.06	0.73	-1.76	6.52
Domestic interest rate (%)	424	5.81	5.74	2.76	0.41	15.21
Expected FX rate depreciation (%)	424	-0.07	-0.04	1.59	-2.92	23.83
Industrial Production (index)	424	139.51	133.70	27.04	95.99	229.25
FX rate change (%)	424	-0.17	-0.28	2.43	-9.04	15.62
Inflation rate (%)	424	4.81	4.30	2.89	-3.39	12.33
Tax equivalent of capital controls in Colombia (URR on portfolio debt inflows)	87	0.48	0.00	0.98	0.00	2.78
Tax equivalent of capital controls in Peru (marginal RR on short term bank liabilities)	86	0.05	0.00	0.21	-1.07	0.50
Other variables						
World interest rate (%)	86	2.00	2.12	1.40	0.22	4.12
Global risk aversion: VIX	86	20.81	17.62	10.62	10.31	68.51

Table 2: Effect of Brazilian capital controls on portfolio inflows to Brazil

Dependent variable Y:		Portfolio inflows to Brazil						Inter-company loans from foreign corporations to Brazilian affiliates
Sample : **2004m4-2011m6** Frequency: **M** Estimation: **OLS, TSLS**		**Bonds**		**Δ Bonds**	**Equity**		**Δ Equity**	
		OLS	**TSLS**	**OLS**	**OLS**	**TSLS**	**OLS**	**OLS**
Brazilian capital controls on portfolio bonds (current and lagged)	$d(IOF\_Bd)_t$	0.106 (0.076)	0.403 (0.427)	-0.247** (0.122)	2.038*** (0.408)	3.725*** (0.698)	1.698*** (0.247)	0.130 (0.108)
	$d(IOF\_Bd)_{t-1}$	-0.188* (0.109)	-0.122 (0.141)	-0.365*** (0.090)	-0.166 (0.185)	-0.470 (0.386)	-2.202*** (0.316)	0.053 (0.054)
	$d(IOF\_Bd)_{t-2}$	0.442** (0.205)	0.462** (0.204)	0.376 (0.252)	0.050 (0.115)	0.327 (0.296)	-0.015 (0.310)	0.024 (0.097)
	$d(IOF\_Bd)_{t-3}$	0.337** (0.155)	0.356** (0.157)	-0.223 (0.203)	-0.231** (0.110)	-0.241 (0.196)	-0.244 (0.301)	-0.047 (0.073)
	$d(IOF\_Bd)_{t-4}$	-0.048 (0.196)	-0.049 (0.187)	-0.509*** (0.156)	-0.081 (0.179)	0.004 (0.200)	-0.119 (0.186)	0.197** (0.080)
Industrial production growth	$InProd\_v_t$	0.279*** (0.084)	0.271*** (0.088)	0.192** (0.088)	0.114* (0.068)	0.090 (0.064)	0.023 (0.047)	-0.026 (0.039)
Money market rate differential adjusted by expected depreciation	d(IR– EER_v– WIR)	0.362 (0.683)	0.345 (0.679)	0.249 (0.569)	-0.302 (0.301)	-0.439 (0.307)	-0.107 (0.179)	0.070 (0.214)
Extreme uncertainty	VIX_extreme	-1.321*** (0.465)	-1.165** (0.526)	-0.618 (0.633)	-1.963*** (0.345)	-1.162* (0.658)	-0.123 (0.324)	0.557 (0.451)
Number of observations		80	80	80	79	79	79	79
Adjusted R^2		0.09	0.08	0.17	0.40	0.18	0.62	0.04
Durbin-Watson		2.02	2.02	2.36	2.02	2.00	2.17	2.00
Instruments Z			$d(IOF\_Eq)_t$, SER_v_{t-1}			$d(IOF\_Eq)_t$, SER_v_{t-1}		
H0: all instruments Z are uncorrelated with Y Prob(J-Statistic)			0.71			0.58		
H0: d(IOF_Bd)t is exogenous p-value (t statistic)			0.72			0.00		

Standard errors are in parenthesis; *, ** and *** denote significance at the usual confidence levels (90%, 95% and 99%). All values are heteroskedasticity-consistent (White estimators). The regressions for bonds include an AR(1) term, whereas those for equity and intercompany loans include AR(1) and AR(2) terms. Neither the constant, nor the AR terms are reported on this table. Along with exogenous instruments Z, the instrumental equation for TSLS estimation includes lagged values of the dependent variable, and current and lagged values of all the regressors.

Table 3: Effect of Brazilian capital controls on portfolio inflows to other Latin American countries

Sample : **2004m4-2011m6** Frequency: **M** Estimation: **OLS, TSLS**	Dependent variable *Y* :	**Portfolio inflows to other Latin America countries (Argentina, Chile, Colombia, Mexico, Peru)**							
		Bonds			**Δ Bonds**	**Equity**			**Δ Equity**
		OLS	**FE**	**TSLS**	**OLS**	**OLS**	**FE**	**TSLS**	**OLS**
Brazilian capital controls on portfolio bonds (current and lagged)	$d(IOF\_Bd)_t$	0.549**	0.554**	2.846***	0.311	0.063**	0.064**	0.014	0.053
		(0.220)	(0.220)	(0.915)	(0.235)	(0.032)	(0.031)	(0.034)	(0.038)
	$d(IOF\_Bd)_{t-1}$	0.113	0.129	2.850***	-0.719***	-0.006	-0.003	-0.045	-0.085**
		(0.294)	(0.293)	(1.084)	(0.157)	(0.039)	(0.038)	(0.039)	(0.036)
	$d(IOF\_Bd)_{t-2}$	0.282	0.306	2.388***	-0.108	-0.029	-0.025	-0.057	-0.034*
		(0.311)	(0.309)	(0.855)	(0.131)	(0.034)	(0.032)	(0.036)	(0.019)
	$d(IOF\_Bd)_{t-3}$	0.732**	0.756**	1.860***	0.267	-0.058*	-0.054*	-0.073**	-0.028
		(0.300)	(0.299)	(0.515)	(0.207)	(0.030)	(0.029)	(0.032)	(0.039)
	$d(IOF\_Bd)_{t-4}$	0.083	0.097	0.366	-0.583***	-0.078***	-0.076***	-0.082***	-0.030
		(0.257)	(0.257)	(0.227)	(0.220)	(0.020)	(0.020)	(0.021)	(0.023)
Industrial production growth	$InProd\_v_t$	0.023*	0.023*	0.031	0.027	-0.003	-0.003	-0.003	-0.004
		(0.013)	(0.013)	(0.020)	(0.022)	(0.004)	(0.004)	(0.004)	(0.014)
Tax equivalent of RR on foreign liabilities (Peru, Colombia)	d(TaxEquiv)	-0.056	-0.058	-0.006	0.115	-0.099	-0.098	-0.268**	0.079
		(0.192)	(0.198)	(0.148)	(0.295)	(0.105)	(0.105)	(0.130)	(0.085)
Money market rate differential adjusted by expected depreciation	d(IR– EER_v– WIR)	0.055	0.055	0.072	0.127**	-0.009*	-0.009*	-0.009*	-0.002
		(0.041)	(0.042)	(0.044)	(0.057)	(0.005)	(0.005)	(0.005)	(0.009)
Extreme uncertainty	VIX_extreme	0.630	0.630	0.624	0.778	-0.058	-0.062	-0.085	0.066
		(0.412)	(0.409)	(0.402)	(0.920)	(0.071)	(0.070)	(0.073)	(0.119)
Number of observations		394	394	394	394	394	394	394	394
Adjusted R^2		0.72	0.72	0.54	0.18	0.53	0.54	0.53	-0.01
Durbin-Watson		2.01	2.02	2.13	2.14	2.10	2.12	2.10	1.99
Instruments *Z*				SER_v_{t-1}				$d(IOF\_Eq)_t$, SER_v_{t-1}	
H0: all instruments *Z* are uncorrelated with *Y* Prob(J-Statistic)				0.04				0.99	
H0: d(IOF_Bd)t is exogenous p-value (t statistic)				0.00				0.29	

Standard errors are in parenthesis; *, ** and *** denote significance at the usual confidence levels (90%, 95% and 99%). All values are heteroskedasticity-consistent (White estimators). The regressions include AR(1) and AR(2) terms. Neither the constant, nor the AR terms are reported on this table. Aong with exogenous instruments Z, the instrumental equation of TSLS includes lagged values of the dependent variable and current and lagged values of all the regressors.

Figure 1: Financial flows to emerging economies

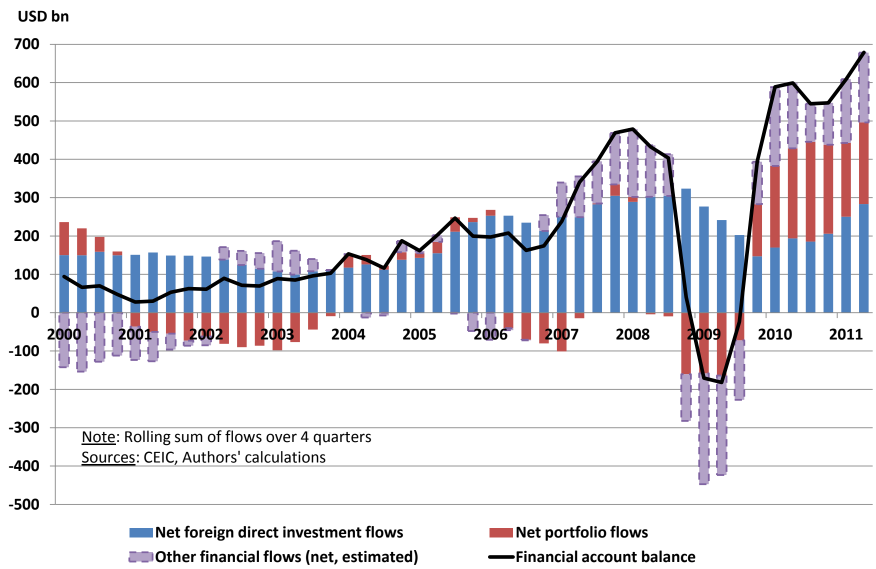


Figure 2: Portfolio bond inflows to Brazil and Mexico and Brazilian IOF

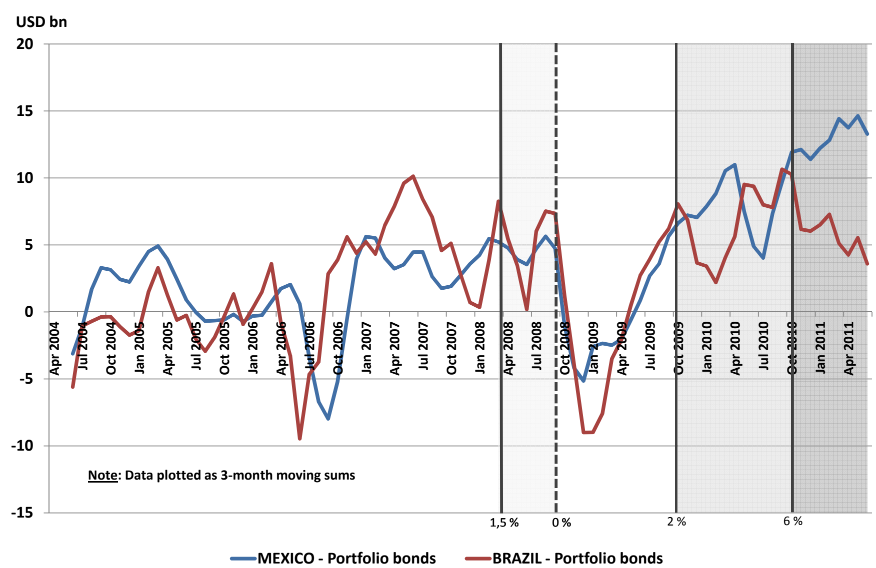


Figure 3: EPFR Bond flows - Correlation matrix

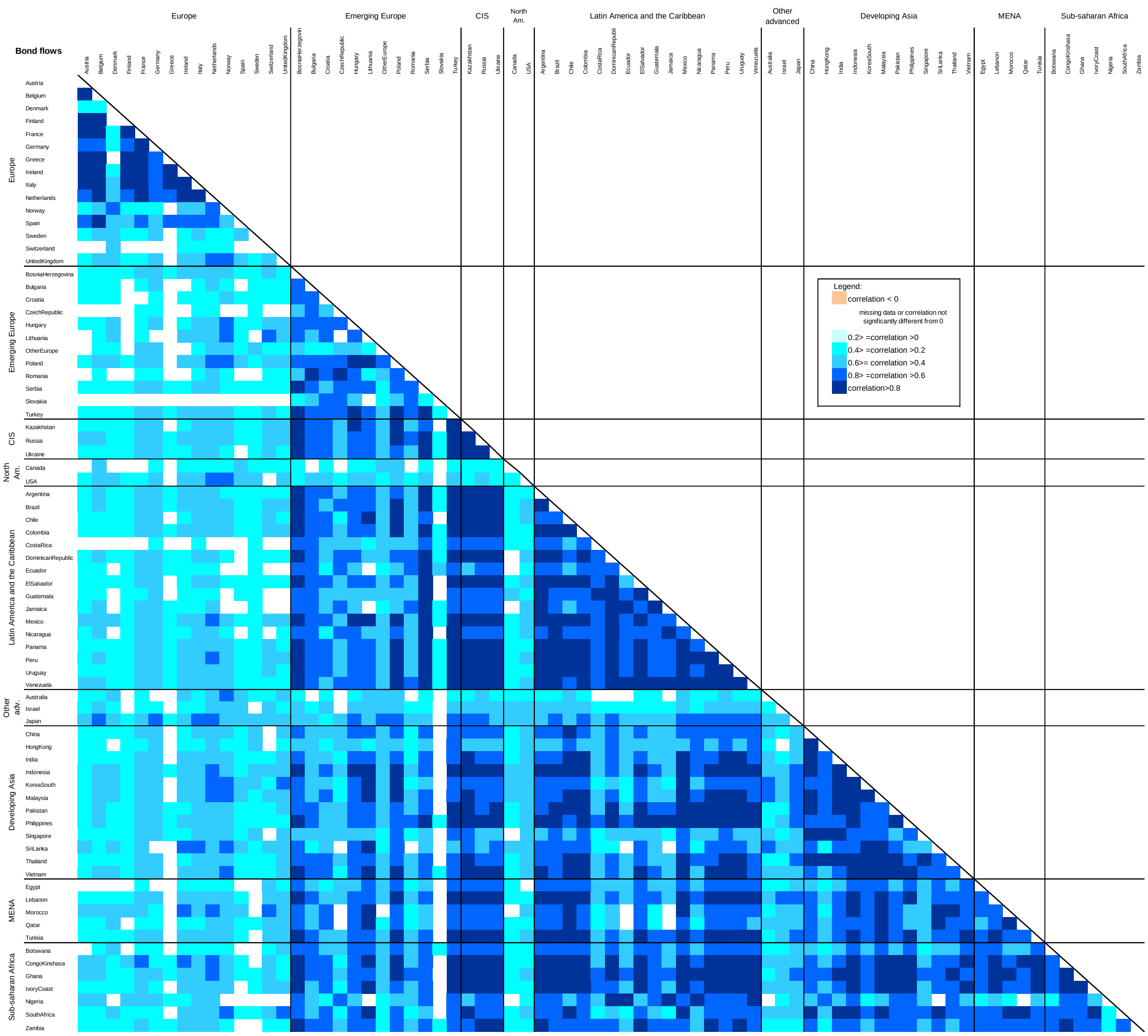


Figure 4: EPFR Equity flows - Correlation matrix

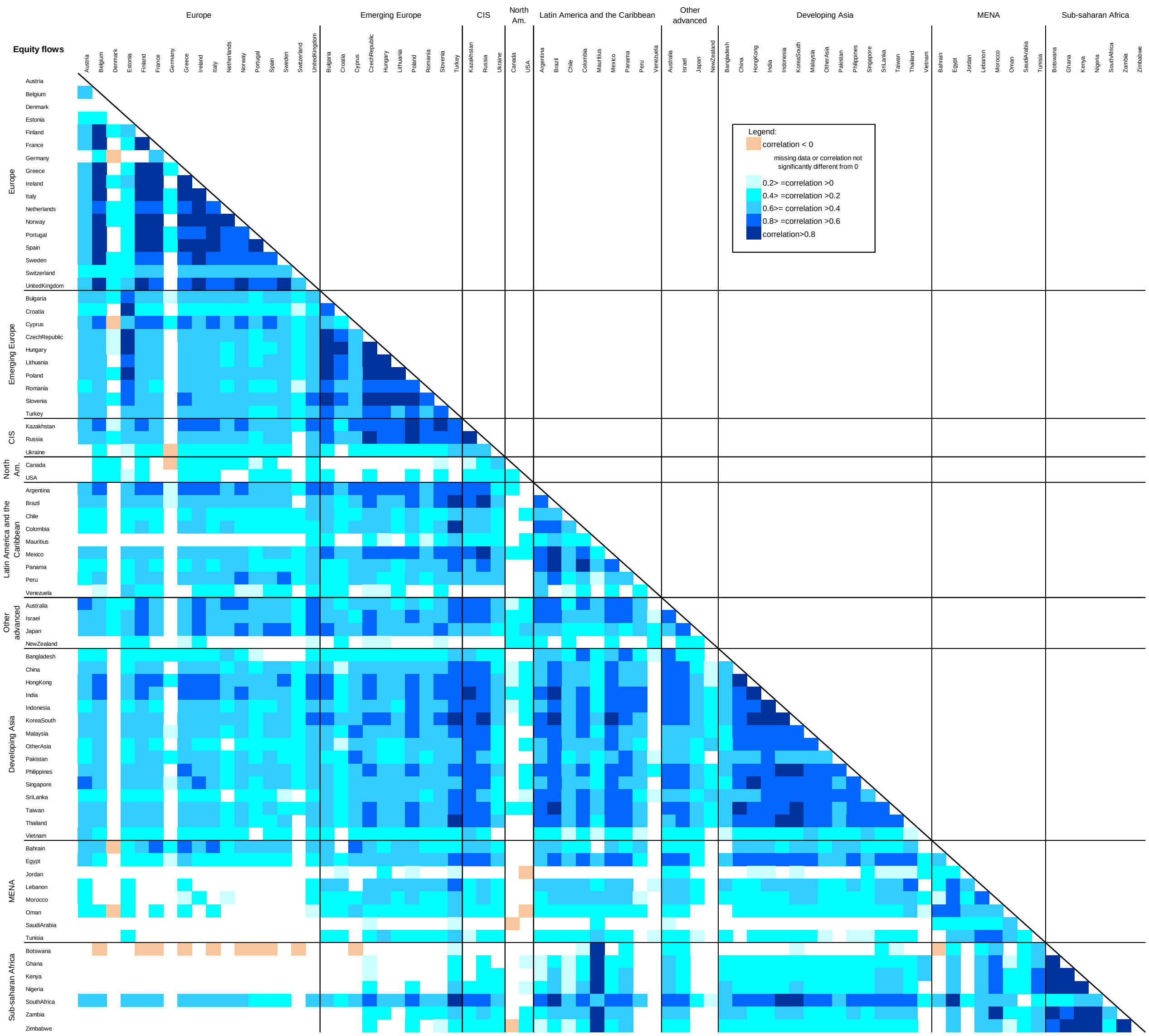
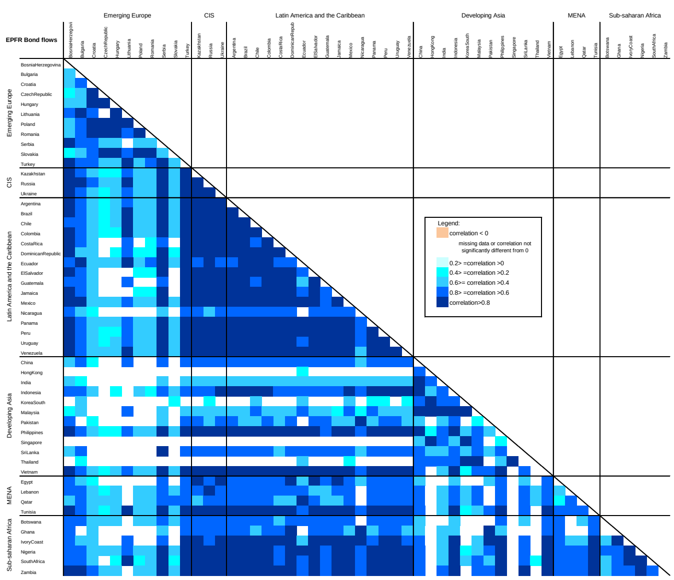
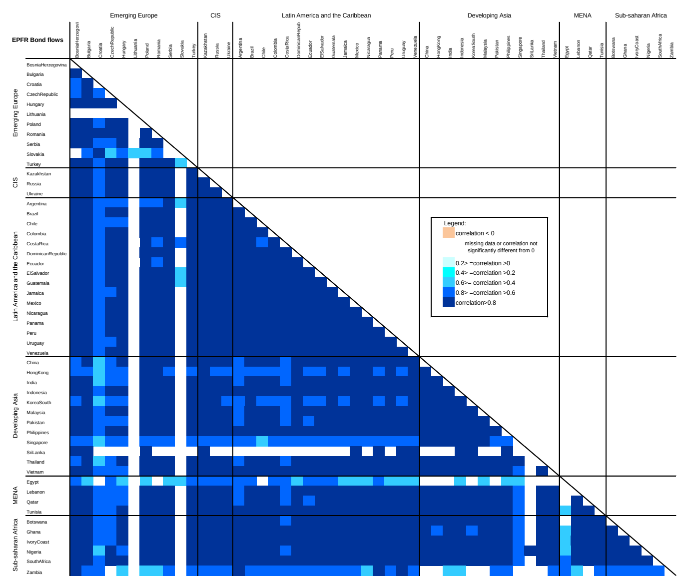


Figure 5: EPFR Bond flows to emerging countries - Correlation matrix before, during and after the crisis

(a) Pre-crisis (April 2004 - July 2007)



(b) Crisis period (August 2007 - March 2009)



(c) Post-crisis (April 2009 - June 2011)

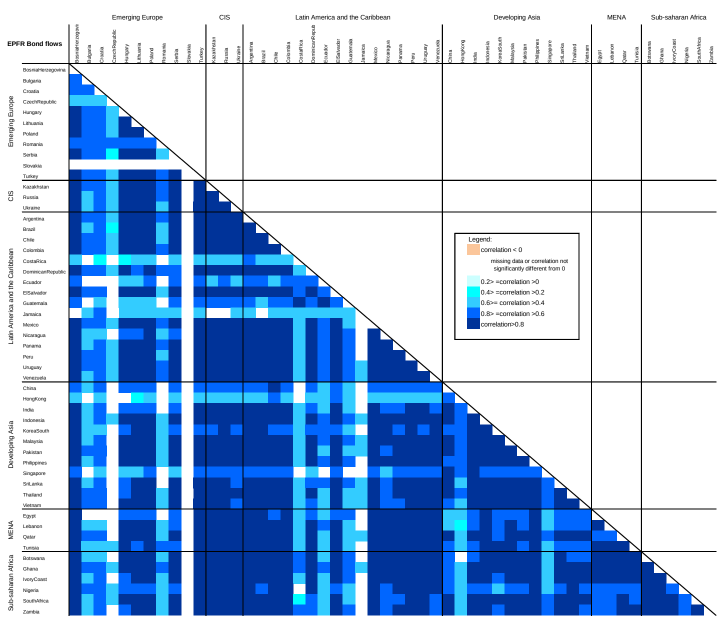


Figure 6: Bilateral correlations of portfolio bond inflows vis-á-vis Brazil (BoP adjusted EPFR monthly data, first differences)

In red, periods where the IOF was in place	Argentina	Chile	Colombia	Mexico	Peru	Uruguay	Venezuela
May 2004 - April2008							
May - October 2008							
November 2008 - September 2009							
October 2009 - June 2011							

Figure 7: Capital controls in Brazil, Colombia and Peru

(a) Brazil (b) Colombia and Peru

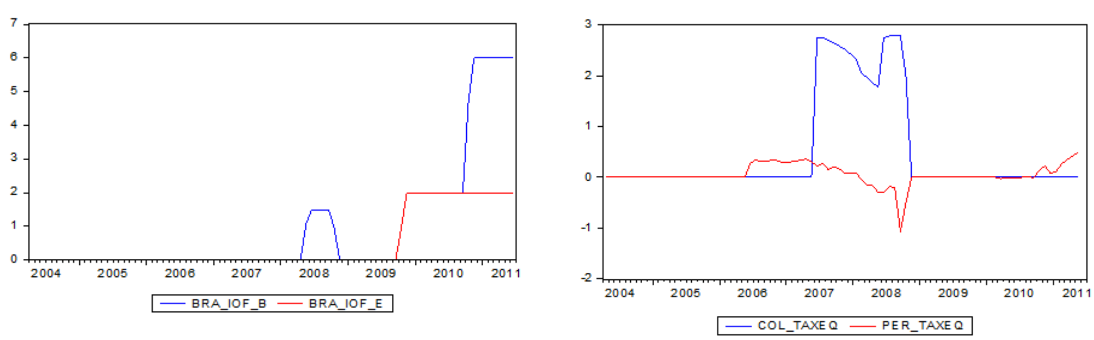
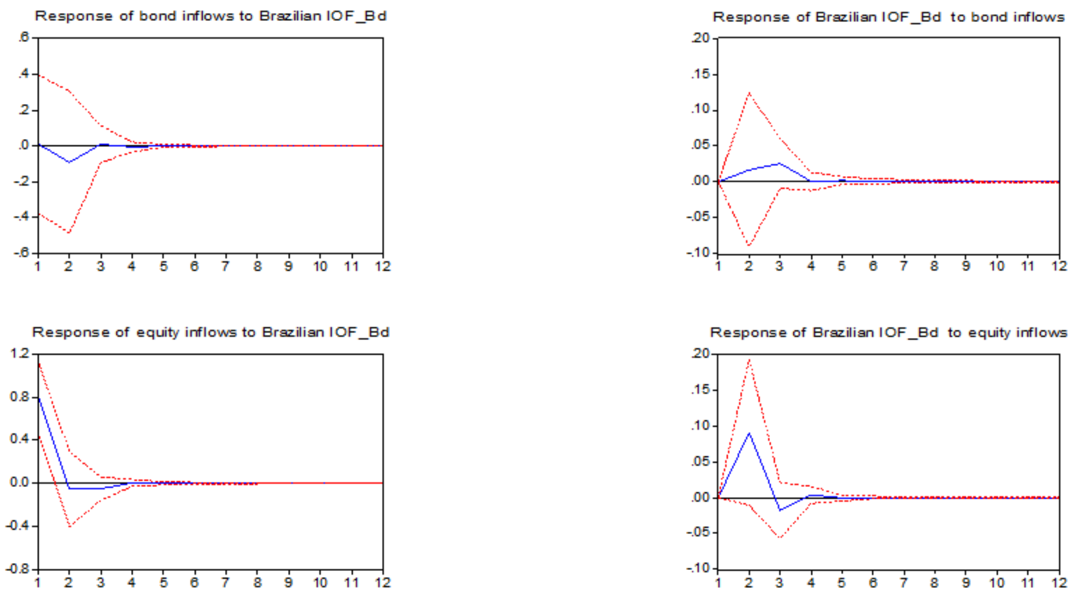
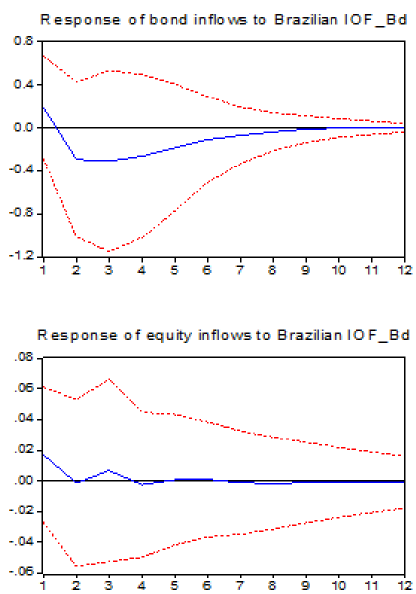


Figure 8: Dynamic responses to a one-standard-deviation shock on Brazilian capital controls

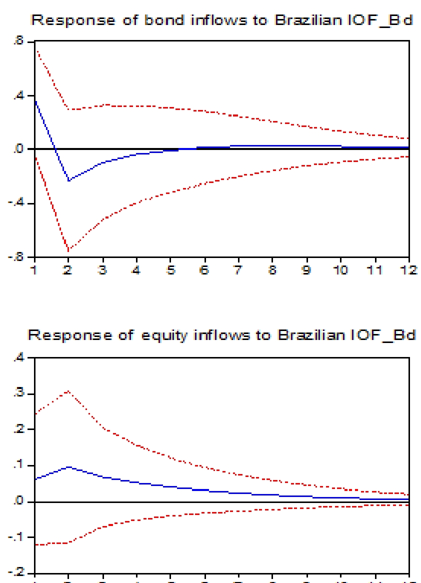
(a) Brazil



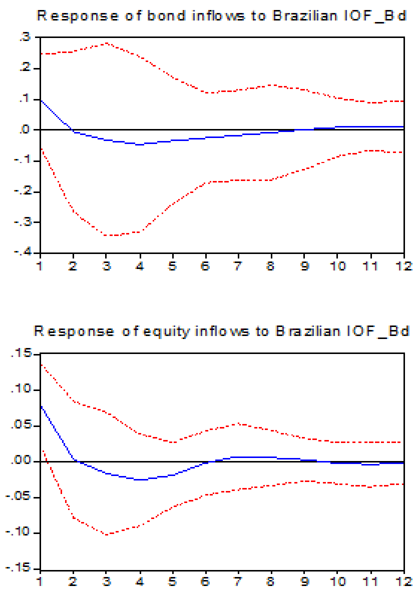
(b) Argentina



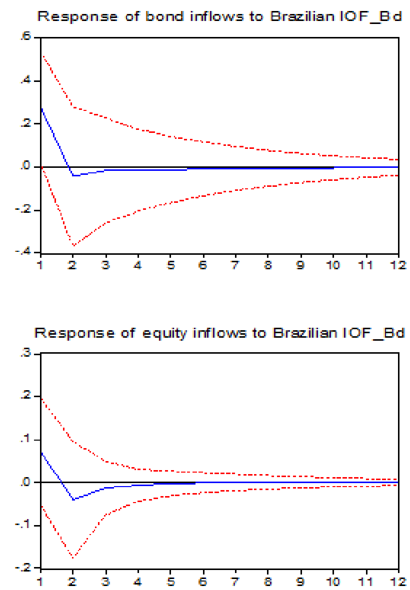
(c) Chile



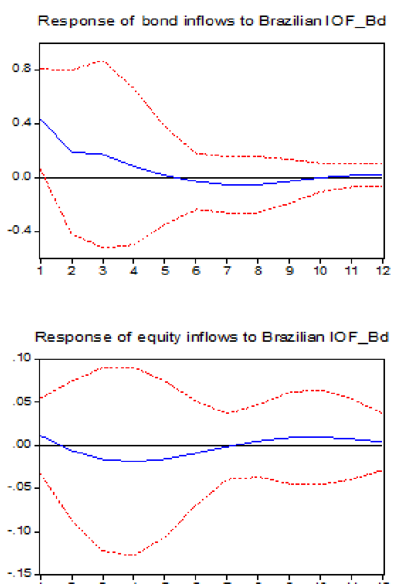
(d) Colombia



(e) Mexico



(f) Peru



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