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Abstract

I study the determinants of capital flows to Argentina, Brazil, and Mexico, assessing the relative importance of domestic and global factors. I estimate six VECM models, one for each Latin American country plus the Euro Area, Japan, and USA, and then embed them in a multi-country Global VAR. The cointegrating space is identified in terms of theoretical long-run relations linking net foreign assets (NFA) to the other variables of the model. The results show that in the long-run external prevail on domestic factors as determinants of the equilibrium behaviour of NFA, with the relative importance of each factor varying from one country to another. Generalized Impulse Response Functions (GIRF) and Forecast Error Variance Decomposition (GFEVD) provide overwhelming evidence that domestic shocks are predominantly responsible for the short-run dynamics of Latin American NFA. Although all previous studies focus on North-American economic influence, one striking result of this paper is that the US variables are by no means the main external factors affecting Latin American NFA. Quite on the contrary, Japanese and, to a lesser extent, European cyclical conditions explain a large proportion of Latin American NFA short-run behaviour.

Keywords: Net foreign assets, capital flows, real exchange rate, Latin America, Emerging markets, VECM, Global VAR.

JEL Classification: F21, F32, C32, C50.

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1 Introduction

The wave of financial globalization characterizing the last two decades has been marked by a surge of private capital flows from industrial to developing countries. Latin American countries belong to the group that has garnered the lion's share of these flows, in particular from the beginning of the 1990s.

The international allocation of capital is determined by internal as well as external factors. The first category includes structural and cyclical conditions, such as economic and political reforms, affecting the prospective growth and the business cycle stance in the domestic economy. The second category refers to the state of global financial markets, including expected rates of return and business cycle phases.

This paper assesses the relative importance of external and internal factors in driving the pattern of capital flows to a selection of Latin American countries, namely Argentina, Brazil, and Mexico. To this end I study the long- and short-run dynamics of these economies' net foreign assets (NFA) in the context of a global economy including three major industrial countries/regions, namely US, Euro Area, and Japan.

The analysis involves two steps. Firstly, I estimate six vector error-correcting mechanism (VECM) models, one for each country/region. In each model the endogenous variables are related to corresponding foreign variables constructed according to the international trade pattern of the economy under consideration. Foreign variables are treated as weakly exogenous (or long-run forcing) based on a small open economy assumption. In order to identify the cointegration space, I impose and test two long-run over-identifying restrictions derived from economic theory. The first restriction is a long-run solvency requirement involving the NFA position, while the second restriction is derived from an intertemporal optimizing model of the real exchange rate. Secondly, the country/region specific models are combined in a consistent Global VAR of the world economy.

The econometric methodology sketched above nests two newly developed approaches to macro-econometric modelling. The first approach, proposed by Garratt, Lee, Pesaran, and Shin (2000, 2003, GLPS), provides a framework for embedding long-run theoretical conditions in an otherwise unrestricted cointegrating VAR. Essentially, this approach starts with a set of long-run relationships derived from macroeconomic theory and generally consisting of stock-flow and accounting identities, arbitrage conditions and long-run solvency requirements. The reduced-form disturbances (derived as functions of the structural shocks) are then incorporated in a VAR to obtain a VECM model having the long-run relations as its steady state solution. The second approach, proposed by Pesaran, Schuermann, and Weiner (2004a, PSW), consists of a procedure for estimating a GVAR composed by a number of country-specific VECM models.

This mixed approach combines the advantages of structural cointegrating VAR over traditional macro-econometric models such as large-scale simultaneous equations, unrestricted, Bayesian, and structural VAR, and DSGE models, with the advantages of GVAR over panel cointegration techniques.[1] As for the former,[2] structural cointegrating VARs provide a manageable mean for macroeconometric modelling with strong theoretical foundations explaining the long-run behaviour of the economy and flexible dynamics able to fit the time series well. As for the GVAR, its advantages over panel cointegration techniques are mentioned by Baltagi (2004) and PSW (2004b). These relate to the possible distortion of within-group cointegration test results caused by the existence of between-group cointegration, as shown by Banerjee *et al.* (2004). For example, Argentinean output is likely to be (and indeed it is, as shown later) cointegrated not only with domestic variables such as NFA or the real exchange rate, but also with foreign variables such as US or European output. This is overcome in the GVAR modelling strategy by allowing for cointegration between domestic and foreign variables at the country-specific model level.

Another strength of the GVAR is that it explicitly allows for interdependence in the global economy in a truly multi-country setting. Although the shocks hitting the economy are not identified according to their economic nature (supply, demand, policy), they are identified basing on their geographic origin. In a panel cointegration framework such an identification would require a formidable number of restrictions among country/region-specific shocks, which could be hardly justified in terms of economic theory. Using the GVAR it is possible to distinguish and identify the shocks originated in three industrial countries/areas, besides those originated in Latin America, rather than considering only one country (the US in the previous literature) or an ambiguous rest of the world.

This paper extends previous literature along four dimensions. First, from a methodological point of view, it is one of the first studies to estimate a GVAR with long-run theory explicitly embodied in the cointegrating relations.[3] In this respect it is a further development of the GVAR along the lines hoped by its authors (PSW (2004 a, b)). This improvement is particularly useful when studying capital flows. In fact, when long-run factors are important, a proper interpretation of short-term flows can only be achieved by taking into account explicitly the equilibrium conditions on the net foreign assets positions (Lane and Milesi Ferretti (2001b)). This implies, for example, that the consequences of current account imbalances are different depending on whether the NFA positions are moving away from or towards the long-run equilibrium.

Second, unlike previous literature, this paper does not only focus on USA or a comprehensive rest of the world as the main source of external shocks to developing countries' NFA, but rather considers the possible effects of disturbances originated also in the Euro Area and Japan. While the Euro Area is historically linked to the Latin American economy, Japan has stronger relationships with the Asian economies. However, surprisingly enough, my results show

[1] By incorporating the structural cointegrating VARs in a GVAR this paper responds to the criticism levied by Dennis and Lopez (2004) who emphasize the importance of stock-flow relations and intertemporal budget constraints for policy analysis.

[2] I refer to GLPS (2000) for a detailed discussion.

[3] To the best of my knowledge, Dees *et al.* (2007) provide the only other example.

that the response of Latin American NFA to Japanese shocks is considerable, thus supporting the world economy specification here presented.

Third, the paper builds on a recent strand of the literature adopting an international portfolio equilibrium approach to the analysis of the current account (Ventura (2003)). In this respect, it is the first paper, as far as I know, which studies push and pull factors of capital flows to developing countries over such a long sample period and with an explicit and coherent global perspective.

Finally, the NFA time series is constructed as an interpolation of the dataset on world foreign assets and liabilities recently developed by Lane and Milesi-Ferretti (2001a, 2006). The important feature of this dataset, based on direct estimates of stocks and on indirect estimates constructed from cumulated flows, is that asset prices adjustments are explicitly accounted for. This allows me to analyse the effects of external and internal factors on NFA considering both flows and stock valuation changes.

Many papers study the main factors driving private capital flows to developing countries. Some authors (e.g. Calvo *et al.* (1993, 1996), Fernàndez-Arias (1996)) attribute most of the flows to external factors, i.e. the cyclical conditions of creditor industrial countries. Accordingly, the flows to Latin America and East Asia at the beginning of the nineties were spurred by the burst of the speculative bubble in Japan, the US cyclical downturn in '91-'92, and the monetary policy stance of the UK following the ERM crisis. Chuhan *et al.* (1998), on the other hand, attribute the changing pattern of capital flows to the changed conditions and performance of emerging economies. Taylor and Sarno (1997) take a disaggregated view and show that in the long-run US bond and equity flows to developing countries are equally sensitive to global and country-specific factors. In the short-run, however, global factors seem to be more important than domestic ones in explaining bond flows, while both equally explain equity flows.

Prasad *et al.* (2003) take a broader view and argue that from a longer-term perspective "pushing" factors include the increasing importance of institutional investors and demographic trends. Institutional investors may have channelled increasing funds to emerging markets by reducing the riskiness of such allocations for individual investors. Moreover, the ageing population in industrial countries need higher returns from savings in order to be able to finance post-retirement consumption, while saving rates in industrial countries decrease as a consequence of the increasing old-age dependency ratios. Long-term "pulling" factors include capital account liberalizations and, more generally, opening up policies as well as large-scale privatization programs. Portes and Rey (2005) focus on the determinants of equity transaction flows for a sample of developed countries using the "gravity model". They find that informational frictions and "familiarity" effects are important for the geographical distribution of international equity flows, and that distance-related proxies affect asset trade in a similar fashion as goods trade.

Assessing the relative role of external and domestic factors in driving capital flows to developing countries has important policy implications. One could be tempted to argue that the adoption of policy measures should depend on whether capital flows are "pulled" or "pushed". In the former case inflows would not represent a problem because they are a consequence of the restoration of the country's creditworthiness, while in the latter case the external shock leading capital flows could be easily reversed and therefore call for policy intervention. However, Agénor and Montiel (1999) cast doubts on this view by emphasizing that what matter for policy design are the specific phenomena that are at work rather than the external or domestic origin of shocks. This is because the welfare implications of shocks depend on a wide variety of factors interacting in a complex way.

Further policy implications refer to the choice of the exchange rate regime. The successful adoption of such a regime might be influenced by the degree of economic and financial international integration and, in particular, by the correlation of cross-country financial shocks, as well as by the short- and long-run relationship between the real exchange rate and financial variables. These issues are addressed in more detail in Section 4.2.4.

The main findings of this paper can be summarized as follows: in the long-run, external prevail on domestic factors as determinants of the equilibrium behaviour of NFA, with the relative importance of each factor varying from one Latin American country to another. A common feature of all models is that short-term interest rates' influence on the long-run equilibrium is the largest and statistically most significant among all variables. The portfolio balance theory of NFA is supported in Argentina and Brazil, but not in Mexico. The evidence in favour of a relationship between the real exchange rate and the NFA in line with the "transfer effect" is robust in the models of Argentina and Mexico. Generalized Impulse Response Functions (GIRF) and Forecast Error Variance Decomposition (GFEVD) provide overwhelming evidence that domestic shocks are predominantly responsible for the short-run dynamics of Latin American NFA. Although all previous studies are concentrated on North-American economic influence, one striking result of this paper is that the US variables are by no means the main external factors affecting Latin American NFA. Quite on the contrary, Japanese and, to a lesser extent, European cyclical conditions play an important role.

The paper is organized as follows. Section 2 describes the theoretical conditions underpinning the long-run pattern of the NFA position and of the real exchange rate. Section 3 presents the empirical specification of the country/region-specific VECM models and of the GVAR resulting from their combination. Section 4 presents and analyse the results. Section 5 concludes.

2 Long-run theoretical framework

As a guide for empirical work, in this Section I present a small open economy model of long-run NFA positions. The model consists of two blocks, each corresponding to a long-run restriction, and it is specifically focused on NFA, the main object of this study. This allows me to neglect other possible cointegration relationships relating the variables in the model.

The first long-run relation derives from a long-run solvency condition expressed in terms of the NFA and results in a consistent relationship between capital flows (considered as NFA changes), *per capita* output, the real exchange rate, and real interest rates.

The second relation derives from an explicit intertemporal optimizing model of the "transfer effect" proposed by Lane and Milesi-Ferretti (2004), in which NFA are linked to the terms of trade, the real exchange rate, and relative *per capita* output.

2.1 Net foreign assets position

Define the stock of financial assets, $\tilde{L}_t$, held by the private sector as the sum of net government debt, $\tilde{D}_t$, and the NFA position, $\tilde{F}_t$. In turn, $\tilde{D}_t$ is the sum of high powered money and domestic government bonds minus the foreign exchange reserves; $\tilde{F}_t$ is the difference between the stock of foreign assets held by domestic residents and domestic assets held by foreign residents. All stocks are measured at the beginning of period t. The symbol '$\sim$' indicates nominal magnitudes.

As a condition for the private sector financial position to be sustainable in the long-term, the following must hold (see GLPS):

$$\frac{\tilde{L}_{t+1}}{\tilde{Y}_t} = \mu \exp(\eta_{l,t+1}) \tag{1}$$

where $\tilde{Y}_t$ is the gross domestic product, μ is a constant, and $\eta_{l,t+1}$ is a stationary process.[4]

The private sector can only hold high powered money and domestic and foreign assets. A "postulated" demand for NFA in the spirit of Branson and Henderson (1985) may be expressed in terms of the trade balance, the domestic and foreign net returns, and the expected change of bonds' market value. The trade balance, in turn, depends ultimately on domestic and foreign *per capita* output and the real exchange rate as determinants of imports and exports. Thus, a conceivable demand function for NFA can be characterized as follows:

$$\frac{\tilde{F}_{t+1}}{\tilde{L}_t} = F_{nfa}\left(\frac{Y_t}{P_t}, \frac{Y_t^*}{P_t^*}, \frac{E_t P_t^*}{P_t}, \rho_{t+1}^e, \rho_{t+1}^{*e}, t\right) \exp(\eta_{nfa,t+1}), \tag{2}$$

where $Y_t = \tilde{Y}_t / POP_{t-1}$, POP_t is the population at time t, P_t is the price level, E_t is the effective exchange rate defined as the domestic price of a unit of foreign currency, $\rho_{t+1}^e \equiv \left[(1+R_t)/\left(1+\frac{P_{t+1}^e - P_t}{P_t}\right)\right] - 1$ and

$\rho_{t+1}^{*e} \equiv \left[(1+R_t^*)\left(1+\frac{E_{t+1}^e - E_t}{E_t}\right)\right] / \left(1+\frac{P_{t+1}^{*e} - P_t^*}{P_t^*}\right) - 1$ are respectively the expected real rates of return on domestic and foreign assets, $*$ denotes foreign

[4] This assumption is consistent with the intertemporal national long-run budget constraint (LRBC) framework used by Obstfeld and Taylor (2004, p. 68) to study the current account dynamics.

variables, and the deterministic time trend, t, accounts for two long-run tendencies in international financial markets. Firstly, the demographic and institutional changes in developed world as well as the opening up of capital accounts in emerging markets (Prasad *et al.* (2003)) boosts financial integration over the very long-run; secondly, as noted by GLPS, the increasing share of wealth held in the form of interest-bearing assets reflects the diffusion of the use of credit cards and, more generally, the changing nature of financial intermediation. $\eta_{nfa,t+1}$ is a stationary process capturing the short-run deviations of the ratio $\tilde{F}_{t+1}/\tilde{L}_t$ from its long-run equilibrium position; it reflects costs and frictions to instantaneous portfolio reallocation which may arise from investors' imperfect information, congestion effects or investment adjustment costs.

There is a considerable ambiguity surrounding the effect of asset demand determinants. I consider each argument in more detail below.

Per capita output Overlapping generations models suggest that higher rates of growth in *per capita* income are associated with higher saving rates and, thus, larger NFA positions as young people's savings depend positively on contemporaneous income and negatively on old-age income; thus an increase in young people's income makes savings rise (Obstfeld and Rogoff (1996)). In models with investments, higher levels of output are associated to lower marginal products of capital which will drive abroad domestic savings. In models with habit formation, if income raises consumption lags behind and savings increases. Therefore, a positive relation between output and NFA should be expected.

On the other hand, infinite-horizon representative-consumer models predict that shocks to permanent output does not affect the current account, and so the NFA position, since consumption smoothing makes agents increase consumption by the same amount as output. Lane and Milesi-Ferretti (2001b) argue that for developing economies operating under credit constraints a negative relationship between NFA position and output may hold if higher output relaxes financial constraints and allows greater access to foreign financial markets. Similarly, Calderón *et al.* (2003) show that if *per capita* output reflects overall economic activity and thus proxies for the mean return on domestic investments, a negative relationship between NFA and output is plausible. This is also predicted by the "postulated" asset demand functions employed by the portfolio balance approach (see Branson and Henderson (1985) for a survey) in which the effect of output on asset demands reflects the assumption that agents hold money only for transactions purposes. In such a framework assets demand depends on expectations on firms' investments profitability, which may be reasonably considered dependent on the economy business cycle stance. Thus, the real *per capita* output is a proxy for the economic climate affecting investment opportunities and returns. Fernàndez Arias (1996) considers the domestic investment climate as unobservable and estimates its effect as the residual capital flow which is not accounted for by other variables. Taylor and Sarno (1997) consider the level of real US industrial production as a "push" factor of capital flows to developing countries, although they do not include a variable reflecting the domestic investment climate. However, they implicitly assume that an improved domestic economic environment, as proxied by an increase in output, may have positive feedbacks on country creditworthiness and thus secure a larger access to international financial markets.

The arguments relating NFA positions and foreign output go the opposite way.

Real exchange rate Traditional portfolio balance models argue that a depreciation of the real exchange rate improves the trade balance and, then, the NFA position, thus implying a positive relationship between NFA and the real exchange rate. In the same direction points the fact that in many developing countries, and especially in Latin America, large and abrupt outflows of foreign capital are often associated with sudden currency devaluations. On the other hand, the exchange rate based stabilization programs adopted in Latin American countries caused large real exchange rate appreciations that in turn drove capital abroad as expectations about unsustainable current account dynamics were formed (Edwards (2001)), which implies a negative relation between NFA and the real exchange rate.

One more possible long-run relation between the NFA position and the real exchange rate relates to the "transfer effect", which is discussed in the next Subsection.

Real interest rates According to the portfolio balance approach, the demand for domestic assets responds positively to domestic rates of return. In turn, rates of return are influenced by the level and composition of net foreign assets due, for example, to a home bias in asset demand or to an upward supply of international funds. This results in a negative relationship between the NFA position and the real interest rate differential.

In emerging markets, however, the real rates of return incorporate large risk premium. Therefore, an increase in the rate of return may indeed reflect an increased riskiness perceived by international investors, and thus be associated with an outflow of capital and an increasing NFA position. If, on the other hand, the country risk premium is inversely related to the NFA position, then the negative relationship between NFA and domestic interest rates may be restored (Lane and Milesi-Ferretti (2001b)).

Finally, an increase in domestic real interest rates might reflect decreasing asset prices and thus be associated with a lower value of the outstanding stock of liabilities. This implies a positive relationship between NFA and domestic interest rates when the measurement of NFA reflects changes in the stock value, as in this study.

Regarding foreign interest rates, it is plausible to claim that their relationship with NFA positions moves along the line set by the portfolio balance approach and thus be positive.

Combining the solvency condition (1) with the asset demand relation (2) yields:

$$\frac{\tilde{F}_{t+1}}{\tilde{Y}_t} = \frac{F_{t+1}}{Y_t} = \mu_{nfa} F_{nfa} \left(\overset{\pm}{\frac{Y_t}{P_t}}, \overset{\pm}{\frac{Y_t^*}{P_t^*}}, \overset{\pm}{\frac{E_t P_t^*}{P_t}}, \overset{\pm}{\rho_{t+1}^{e}}, \overset{\pm}{\rho_{t+1}^{*e}}, t \right) \exp(\eta_{l,t+1} + \eta_{nfa,t+1}), \tag{3}$$

where $F_t = \tilde{F}_t / POP_{t-1}$. Equation (3) provides the first long-run relationship to be included in the model, with the signs of partial derivatives indicated over each argument.

2.2 Relationship between net foreign assets and the real exchange rate: the "transfer problem"

Much of the existing literature considers the Purchasing Power Parity (PPP) hypothesis as a valid long-run explanation of the real exchange rate, which, accordingly, is constant. However, the PPP is a weak empirical model and alternative theories suggest that the long-run real exchange rate is time-varying. Dating back to the classic debate between Keynes (1929) and Ohlin (1929) on the economic effects of German war reparations, i.e. the "transfer problem", some literature suggests that non-zero NFA positions are associated, in the long-run, with some degree of adjustment of the real exchange rate.

Lane and Milesi-Ferretti (2004) build a simple theoretical model explaining the long-run co-movement between the real exchange rate, net foreign assets, relative GDP, and the terms of trade. While I refer to that paper for a formal description of the model, I briefly discuss here the main reasoning.

The small open economy produces traded and non-traded goods. The agent's utility depends on an index of both goods and on labour effort, and he can invest in international real bonds. By construction, the terms of trade may influence the real exchange rate only indirectly through a wealth effect on the relative price of non-tradable goods. By taking a linear approximation around a benchmark steady-state, Lane and Milesi-Ferretti (2004) derive a negative relationship between the real exchange rate on one hand and the net foreign assets, the level of output and the terms of trade on the other.

The rationale for a role played by relative output is provided by the Balassa-Samuelson hypothesis: a rise in the output level raises the relative prices of non-tradables since productivity growth is concentrated in the tradable sector, thus leading to an appreciation of the real exchange rate, defined as $(E_t P_t^*)/P_t$, through an increase in the level of domestic prices P_t. Alternatively, an increase in wealth driven by a higher output may reduce the domestic labour supply again increasing the relative price of non-tradables.

The terms of trade affect the real exchange rate if there is home bias in the consumption of tradables: in this case an increase in the relative price of home exports leads to a rise in the relative domestic consumer price level. The terms of trade can also affect the real exchange rate via a wealth effect: an increase in the terms of trade boosts the domestic real income and has an equivalent effect to an increase in domestic output as discussed above.

The net foreign assets may influence the real exchange rate through a number of mechanisms. For example, a transfer to a country raises spending on non-tradables, thus wages increase, the export sector declines and the real exchange rate appreciates.[5]

To sum up, any factor that raises consumption of tradables reduces labour supply to the non-tradables sector through a wealth effect. This, in turn, makes the relative price of non-tradables increase and the real exchange rate appreciate.

The above discussion implies the following long-run (steady state) relationship:

$$Q_t \equiv \frac{E_t P_t^*}{P_t} = \mu_q F_q \left(\overset{-}{\frac{Y_t}{P} / \frac{Y_t^*}{P_t^*}}, \overset{-}{\frac{F_{t+1}}{Y_t}}, \overset{-}{TT_t} \right) \exp(\eta_{q,t+1}) \tag{4}$$

where TT_t denotes the terms of trade and $\eta_{q,t+1}$ is a stationary process capturing the temporary deviations from the long-run relation.[6] The signs of the partial derivatives are reported above each argument.

3 Econometric methodology

The econometric methodology consists of estimating a VECM model for each country/region; the country-specific models are then stacked in a single Global VAR of the world economy.

3.1 Econometric specification of the long-run theoretical model

Taking log-linear approximations and rearranging, the two long-run equilibrium relationships (3) and (4) can be written as follows:

$$\begin{aligned} \xi_{1,t+1} &= \pm\beta_{11} y_t \pm \beta_{12} sr_t \pm \beta_{13} lr_t \pm \beta_{14} q_t + nfa_t & (5)\\ &\quad \pm\beta_{15} y_t^* \pm \beta_{16} sr_t^* \pm \beta_{17} lr_t^* - b_{10} - b_{11} t \\ \xi_{2,t+1} &= \beta_{21}(y_t - y_t^*) + q_t + \beta_{24} nfa_t + \beta_{28} tt_t - b_{20} & (6) \end{aligned}$$

where $\xi_{j,t+1}$ for $i = 1, 2$ denotes the deviations from the equilibrium relationships,[7] $y_t = \ln(Y_t/P_t)$, $r_t = \ln \rho_{t+1}$, with sr_t and lr_t being the short- and long-term rates respectively, the latter included because of their possible effects on the long-term components of NFA, i.e. bonds and foreign direct investments,

[5] See Lane and Milesi-Ferretti (2002) for more discussion on the link between net foreign assets and the real exchange rate.

[6] Note that Lane and Milesi-Ferretti (2004) define the real exchange rate as the ratio between domestic and foreign prices, so that they specify a positive relationship between it and its main determinants, relative output, net foreign assets, and terms of trade.

[7] Söderlind and Vredin (1995) argue against considering the fluctuations around the cointegrating relationships as equilibrium errors. They show how the fluctuations can be interpreted as equilibrium relations themselves depending on the deep parameters of the model.

$q_t = \ln(E_t P_t^*/P_t)$, $nfa_t = \ln(F_{t+1}/Y_t)$, $tt_t = \ln(TT_t)$, b_{j0} for $j = 1, 2$ are constants, t is a time trend, b_{11} is the time trend coefficient, β_{jk} for $j = 1, 2$ and $k = 1, ..., 8$ are coefficients, and $*$ denotes foreign variables. Expected values are proxied by actual realizations of variables at time $t + 1$.

The long-run reduced form disturbances, $\xi_{j,t+1}$, can be written in a (2×1) vector $\boldsymbol{\xi}_t$ which in turn can be expressed as a linear combination of the variables in the system:

$$\boldsymbol{\xi}_t = \boldsymbol{\beta}' \mathbf{v}_{t-1} - \mathbf{b}_0 - \mathbf{b}_1 (t-1) \tag{7}$$

where:

$$\begin{aligned} \mathbf{v}_t &= (y_t, sr_t, lr_t, q_t, nfa_t, y_t^*, sr_t^*, lr_t^*, tt_t, p_t^o)' \\ \mathbf{b}_0 &= (b_{10}, b_{20})' \\ \mathbf{b}_1 &= (b_{11}, 0)' \\ \boldsymbol{\xi}_t &= (\xi_{1,t}, \xi_{2,t})' \end{aligned} \tag{8}$$

where the logarithm of oil price, p_t^o, is included, following PSW, in order to capture the effects of global political and economic events, and:

$$\boldsymbol{\beta}' = \begin{pmatrix} \pm\beta_{11} & \pm\beta_{12} & \pm\beta_{13} & \pm\beta_{14} & 1 & \pm\beta_{15} & \pm\beta_{16} & \pm\beta_{17} & 0 & 0 \\ \beta_{21} & 0 & 0 & 1 & \beta_{24} & -\beta_{21} & 0 & 0 & \beta_{28} & 0 \end{pmatrix} \tag{9}$$

that is, $\boldsymbol{\beta}'$ is the (2×10) matrix of parameters describing the equilibrium relationships. The first row of $\boldsymbol{\beta}'$ relates to the NFA long-run equilibrium derived from the portfolio balance theory, defined by (5), and is normalized on nfa_t; the second relates to the long-run equilibrium for the real exchange rate (the transfer effect), defined by (6), and is normalized on q_t. The over-identifying restrictions implied by the economic theory outlined in Section 2 are imposed on the cointegrating matrix.

3.2 GVAR model

There are $N+1$ countries/regions in the world economy indexed by $i = 0, 1, ...N$.[8] Consider the following VECM model:[9]

$$\begin{aligned} \Delta \mathbf{x}_{it} = \; & \mathbf{a}_{i0} + \mathbf{a}_{i1} t + \mathbf{a}_{i2} \mathbf{D}_{it} - (\mathbf{I}_{k_i} - \boldsymbol{\Phi}_i) \mathbf{x}_{i,t-1} + (\boldsymbol{\Lambda}_{i0} + \boldsymbol{\Lambda}_{i1}) \mathbf{x}_{i,t-1}^* \\ & + (\boldsymbol{\Theta}_{i0} + \boldsymbol{\Theta}_{i1}) \, tt_{i,t-1} + (\boldsymbol{\Psi}_{i0} + \boldsymbol{\Psi}_{i1}) p_{t-1}^o \\ & + \boldsymbol{\Lambda}_{i0} \Delta \mathbf{x}_{it}^* + \boldsymbol{\Theta}_{i0} \boldsymbol{\Delta} tt_{it} + \boldsymbol{\Psi}_{i0} \boldsymbol{\Delta} p_t^o + \boldsymbol{\varepsilon}_{it}, \end{aligned} \tag{10}$$

[8] $N = 5$ in this paper.

[9] The exposition refers to a VARX* of order one, as suggested by the standard information criteria and by the diagnostic tests discussed below.

where $\mathbf{x}_{it}$ is a $(k_i \times 1)$ vector of country i domestic variables, $\mathbf{x}^*_{it}$ is a $(k^*_i \times 1)$ vector of foreign variables specific to country i (to be defined below), tt_{it} and p^o_t are defined above, $\mathbf{a}_{i0}$ is a $(k_i \times 1)$ vector of fixed intercepts, $\mathbf{a}_{i1}$ is a $(k_i \times 1)$ vector of coefficients of the deterministic time trend, $\mathbf{a}_{i2}$ is a $(k_i \times m)$ matrix of coefficients of the exogenous $I(0)$ deterministic components included in the $(m \times 1)$ vector $\mathbf{D}_{it}$, $\mathbf{\Phi}_i$ is a $k_i \times k_i$ matrix of lagged coefficients, $\mathbf{\Lambda}_{ij}$, for $j = 0, 1$, are $(k_i \times k^*_i)$ matrices of coefficients associated to the foreign variables, Θ_{ij} and $\mathbf{\Psi}_{ij}$, for $j = 0, 1$, are $(k_i \times 1)$ vectors associated to the terms of trade and oil price respectively, $\boldsymbol{\varepsilon}_{it}$ is a $(k_i \times 1)$ vector of idiosyncratic, serially uncorrelated, country-specific shocks, with:

$$\boldsymbol{\varepsilon}_{it} \sim i.i.d.\,(\mathbf{0}, \mathbf{\Sigma}_{ii})$$

where $\mathbf{\Sigma}_{ii}$ is non-singular, and $i = 0, 1, ..N$, $t = 1, 2, .., T$. The GVAR model allows for non-zero contemporaneous dependence of shocks across economies via cross-country covariances:

$$\textstyle\sum_{ij} = Cov(\boldsymbol{\varepsilon}_{it}, \boldsymbol{\varepsilon}_{jt}) = E(\boldsymbol{\varepsilon}_{it}\boldsymbol{\varepsilon}'_{jt}), \text{ for } i \neq j$$

The foreign variables $\mathbf{x}^*_{it}$ are weighted averages of the variables of the rest of the world with country/region-specific weights given by trade shares, i.e. the share of country j in the total trade of country i measured in US dollars, in a given base year (1995 in this paper).[10] Thus:

$$w_{ii} = 0, \ \forall i = 0, 1, ..., N.$$

and:

$$\sum_{j=0}^{N} w_{ij} = 1, \ \forall i, j = 0, 1, ..., N.$$

Therefore, a generic foreign variable x^*_{it} is given by:

$$x^*_{it} = \sum_{j=0}^{N-1} w_{ij} x_{jt} \tag{11}$$

The foreign variables, $\mathbf{x}^*_{it}$, the terms of trade, tt_{it}, and the oil price, p^o_t, are treated as weakly exogenous (or long-run forcing), which amounts to considering each economy as small when compared to the rest of the world.

The model (10) can be rewritten as:

$$\Delta\mathbf{x}_{it} = \mathbf{a}_{i0} + \mathbf{a}_{i1}t + \mathbf{a}_{i2}\mathbf{D}_{it} - \mathbf{\Pi}_i\mathbf{v}_{i,t-1} + \mathbf{\Lambda}_{i0}\Delta\mathbf{x}^*_{it} + \mathbf{\Theta}_{i0}\mathbf{\Delta}tt_{it} + \mathbf{\Psi}_{i0}\mathbf{\Delta}p^o_t + \boldsymbol{\varepsilon}_{it} \tag{12}$$

[10] Given the focus of the present study on international capital flows, it would be interesting to weigh foreign variables using some measure of bilateral financial flows or stocks, such as, for example, investors' holdings or banks' exposures. However, apart from concerns related to data availability, this would bias the relationships in favour of certain portfolio components, while my variable being NFA I consider all kind of financial assets. Moreover, trade based weights are common practice in macro-econometric modelling (Wallis (2004)).

where:

$$\mathbf{v}_{i,t-1} = \begin{pmatrix} \mathbf{z}_{i,t-1} \\ tt_{i,t-1} \\ p^o_{t-1} \end{pmatrix} \text{ and } \mathbf{z}_{it} = \begin{pmatrix} \mathbf{x}_{it} \\ \mathbf{x}^*_{it} \end{pmatrix} \tag{13}$$

and $\mathbf{\Pi}_i = (\mathbf{A}_i - \mathbf{B}_i, -\mathbf{\Theta}_{i0} - \mathbf{\Theta}_{i1}, -\mathbf{\Psi}_{i0} - \mathbf{\Psi}_{i1})$.

The number of long-run relations is given by the rank $r_i \leq k_i$ of the $k_i \times (k_i + k_i^* + 2)$ matrix $\mathbf{\Pi}_i$. Under the assumption that $\mathbf{\Pi}_i$ is rank deficient, one can write:

$$\mathbf{\Pi}_i = \boldsymbol{\alpha}_i \boldsymbol{\beta}_i' \tag{14}$$

where $\boldsymbol{\alpha}_i$ is the $k_i \times r_i$ loading matrix and $\boldsymbol{\beta}_i'$ is the $r_i \times (k_i + k_i^* + 2)$ matrix of cointegrating vectors.

In order to avoid introducing quadratic trends in the levels of the variables when $\mathbf{\Pi}_i$ is rank-deficient, I impose the following $(k_i - r_i)$ restrictions on the trend coefficients:

$$\mathbf{a}_{i1} = \mathbf{\Pi}_i \boldsymbol{\kappa}_i$$

where $\boldsymbol{\kappa}_i$ is a $(k_i + k_i^* + 1) \times 1$ vector of fixed constants.[11] Thus, equation (12) becomes:

$$\begin{aligned} \Delta \mathbf{x}_{it} &= \mathbf{a}_{i0} + \mathbf{a}_{i2}\mathbf{D}_{it} + \mathbf{\Pi}_i\boldsymbol{\kappa}_i - \mathbf{\Pi}_i\left[\mathbf{v}_{i,t-1} - \boldsymbol{\kappa}_i(t-1)\right] + \mathbf{\Lambda}_{i0}\Delta\mathbf{x}^*_{it} \\ &\quad + \mathbf{\Theta}_{i0}\mathbf{\Delta} tt_{it} + \mathbf{\Psi}_{i0}\mathbf{\Delta} p^o_t + \boldsymbol{\varepsilon}_{it} \end{aligned} \tag{15}$$

The theoretical framework developed in Section 2 implies a set of over-identifying restrictions on the cointegration space summarized by (14), which, for country i is:

$$\boldsymbol{\xi}_{it} = \boldsymbol{\beta}_i'\mathbf{v}_{i,t-1} - \mathbf{b}_{i0} - \mathbf{b}_{i1}(t-1)$$

Imposing these restrictions on model (15) yields:

$$\Delta \mathbf{x}_{it} = \mathbf{c}_{i0} + \mathbf{c}_{i2}\mathbf{D}_{it} - \boldsymbol{\alpha}_i\boldsymbol{\xi}_{it} + \mathbf{\Lambda}_{i0}\Delta\mathbf{x}^*_{it} + \mathbf{\Theta}_{i0}\mathbf{\Delta} tt_{it} + \mathbf{\Psi}_{i0}\mathbf{\Delta} p^o_t + \boldsymbol{\varepsilon}_{it} \tag{16}$$

where $\mathbf{c}_{i0} = \mathbf{a}_{i0} + \mathbf{\Pi}_i\boldsymbol{\kappa}_i - \boldsymbol{\alpha}_i\mathbf{b}_{i0}$, $\mathbf{c}_{i2} = \mathbf{a}_{i2}$, and $\boldsymbol{\alpha}_i\mathbf{b}_{i1} = \mathbf{\Pi}_i\boldsymbol{\kappa}_i$.

Rather than estimating directly the complete system composed by the $N+1$ country-specific models (16) together with the relations (11), PSW propose to estimate the parameters of each country-specific model separately. The estimated parameters of the country-specific models are then stacked together to form a Global VAR, which in reduced form is given by:

$$\mathbf{A}_i\mathbf{z}_{it} = \mathbf{a}_{i0} + \mathbf{a}_{i1}t + \mathbf{a}_{i2}\mathbf{D}_{it} + \mathbf{B}_i\mathbf{z}_{i,t-1} + \mathbf{\Theta}_{i0}tt_{it} + \mathbf{\Theta}_{i1}tt_{i,t-1} + \mathbf{\Psi}_{i0}p^o_t + \mathbf{\Psi}_{i1}p^o_{t-1} + \boldsymbol{\varepsilon}_{it} \tag{17}$$

[11] This formulation corresponds to model IV in Pesaran *et al.* (2000).

where $\mathbf{A}_i = (\mathbf{I}_{k_i}, -\mathbf{\Lambda}_{i0})$ and $\mathbf{B}_i = (\mathbf{\Phi}_i, \mathbf{\Lambda}_{i1})$ are matrices of dimension $k_i \times (k_i + k_i^*)$ and $\mathbf{A}_i$ has a full row rank.

Now, collect all country/region-specific endogenous variables in the $k \times 1$ global vector $\mathbf{x}_t = (\mathbf{x}'_{0t}, \mathbf{x}'_{1t}, ..., \mathbf{x}'_{Nt})'$ where $k = \sum_{i=0}^{N} k_i$. Then:

$$\mathbf{z}_{it} = \mathbf{W}_i \mathbf{x}_t$$

where $\mathbf{W}_i$ is the $(k_i + k_i^*) \times k$ matrix collecting the trade weights w_{ij}, $\forall i, j = 0, 1, ...N$. Therefore, equation (17) becomes:

$$\mathbf{A}_i \mathbf{W}_i \mathbf{x}_t = \mathbf{a}_{i0} + \mathbf{a}_{i1} t + \mathbf{a}_{i2} \mathbf{D}_{it} + \mathbf{B}_i \mathbf{W}_i \mathbf{x}_{t-1} + \mathbf{\Theta}_{i0} tt_{it} + \mathbf{\Theta}_{i1} tt_{i,t-1} + \mathbf{\Psi}_{i0} p_t^o + \mathbf{\Psi}_{i1} p_{t-1}^o + \boldsymbol{\varepsilon}_{it} \tag{18}$$

Stacking the $N+1$ systems (18) yields:

$$\mathbf{G}\mathbf{x}_t = \mathbf{a}_0 + \mathbf{a}_1 t + \mathbf{a}_2 \mathbf{D}_t + \mathbf{H}\mathbf{x}_{t-1} + \mathbf{\Theta}_0 \mathbf{tt}_t + \mathbf{\Theta}_1 \mathbf{tt}_{t-1} + \mathbf{\Psi}_0 p_t^o + \mathbf{\Psi}_1 p_{t-1}^o + \boldsymbol{\varepsilon}_t \tag{19}$$

where $\mathbf{G}$ is a $k \times k$ matrix of full rank, $\mathbf{a}_h = (\mathbf{a}_{0h}, ..., \mathbf{a}_{Nh})'$ for $h = 0, 1, 2$, $\mathbf{G} = (\mathbf{A}_0 \mathbf{W}_0, ..., \mathbf{A}_N \mathbf{W}_N)'$, $\mathbf{H} = (\mathbf{B}_0 \mathbf{W}_0, ..., \mathbf{B}_N \mathbf{W}_N)'$,

$$\mathbf{\Theta}_h = \begin{pmatrix} \mathbf{\Theta}_{0h} & \mathbf{0} & \mathbf{0} & \mathbf{0} \\ \mathbf{0} & \mathbf{\Theta}_{1h} & \mathbf{0} & \mathbf{0} \\ \mathbf{0} & \mathbf{0} & . & \mathbf{0} \\ \mathbf{0} & \mathbf{0} & \mathbf{0} & \mathbf{\Theta}_{Nh} \end{pmatrix}$$

for $h = 0, 1$, $\mathbf{\Psi}_h = (\mathbf{\Psi}_{0h}, ..., \mathbf{\Psi}_{Nh})'$ for $h = 0, 1$, $\mathbf{D}_t = (\mathbf{D}_{0t}, ..., \mathbf{D}_{Nt})'$, $\mathbf{tt}_t = (\mathbf{tt}_{0t}, ..., \mathbf{tt}_{Nt})'$.

The model (19) has the following error correction form:

$$\mathbf{G}\mathbf{\Delta}\mathbf{x}_t = \mathbf{a}_0 + \mathbf{a}_1 t + \mathbf{a}_2 \mathbf{D}_t - (\mathbf{G} - \mathbf{H}, \mathbf{\Theta}_0 + \mathbf{\Theta}_1, \mathbf{\Psi}_0 + \mathbf{\Psi}_1)\mathbf{y}_{t-1} + \mathbf{\Theta}_0 \mathbf{\Delta}\mathbf{tt}_t + \mathbf{\Psi}_0 \mathbf{\Delta} p_t^o + \boldsymbol{\varepsilon}_t \tag{20}$$

where:

$$\mathbf{y}_{t-1} = \begin{pmatrix} \mathbf{x}_{t-1} \\ \mathbf{tt}_{t-1} \\ p_{t-1}^o \end{pmatrix}$$

The number of long-run relationships in the global model, determined by the rank of $(\mathbf{G} - \mathbf{H})$, cannot exceed the sum of the long-run relationships existing in the country/region-specific models (PSW).

3.2.1 Persistence profiles, impulse response functions and forecast error variance decomposition

The persistence profiles, proposed by Pesaran and Shin (1996), are the time profiles of the effects of a system-wide shock on the cointegrating relations. If the vector under investigation is indeed a cointegrating vector the value of the persistence profile is unity on impact, while it converges to zero as the horizon tends to infinity. The persistence profile is independent of the way the shocks are orthogonalized or the order of the variables and equations in the VAR model.

In the GVAR the cointegrating relations are estimated in the country/region-specific models and are thus determined in terms of the country/region variables, $\boldsymbol{\beta}_i'\mathbf{z}_{i,t-1}$; therefore, in order to compute the persistence profiles, an appropriate mapping between $\mathbf{z}_{i,t}$ and the variables in the GVAR, $\mathbf{x}_t$, is required. The persistence profile, $\mathbf{PP}$, of $\boldsymbol{\beta}_{ji}'\mathbf{z}_{i,t-1}$ with respect to a system-wide shock to $\boldsymbol{\varepsilon}_t$ is thus given by:

$$\mathbf{PP}_{\boldsymbol{\beta}_{ji}'\mathbf{z}_{i,t-1};\boldsymbol{\varepsilon}_t}(n) = \frac{\boldsymbol{\beta}_{ji}'\mathbf{W}_i F^n \mathbf{G}^{-1}\boldsymbol{\Sigma}\mathbf{G}'^{-1} F'^n \mathbf{W}_i'\boldsymbol{\beta}_{ji}}{\boldsymbol{\beta}_{ji}'\mathbf{W}_i\mathbf{G}^{-1}\boldsymbol{\Sigma}\mathbf{G}'^{-1}\mathbf{W}_i'\boldsymbol{\beta}_{ji}} \tag{21}$$

where $\boldsymbol{\beta}_{ji}$ is the j^{th} cointegrating relation in the i^{th} country ($j = 1, 2, .., r_i$; $i = 1, 2, ..., N$), n is the horizon, $\mathbf{W}_i$ is the country i trade weight matrix, $F = \mathbf{G}^{-1}\mathbf{H}$, and $\boldsymbol{\Sigma}$ is the $k \times k$ variance-covariance matrix of the shocks $\boldsymbol{\varepsilon}_t$.

Two widely used tools for analysing dynamic models are the impulse response function and the forecast error variance decomposition. In carrying out these analyses it is important to identify the shocks that hit the economy in a proper manner. This is usually accomplished by orthogonalizing the shocks or by means of the structural VAR methodology. The results of the first approach rely critically on the ordering of the variables, which is not unique. The second approach requires the imposition of a number of restrictions derived from economic theory in order to identify all the possible shocks, which does not seem feasible within a Global VAR framework.[12] For the purpose of this study the identification of shocks according to their economic nature (supply, demand, policy) is unnecessary since the focus is more on their geographic origin. As noted by PSW, in the GVAR methodology the regional identification of shocks is accomplished by conditioning the estimation of the country/region-specific models on foreign variables, which leaves only a modest cross-country correlation among the residuals of endogenous variables.[13]

Therefore, in this paper the dynamic analysis is carried out by using the Generalized Impulse Response Function (GIRF) and the Generalized Forecast Error Variance Decomposition (GFEVD) developed by Koop *et al.* (1996) and Pesaran and Shin (1998).

The GIRF amounts to shocking the l^{th} variable in the i^{th} model and integrating the effects of other shocks using a historically observed distribution of the errors. This yields:

$$\mathbf{GI}_{x;\varepsilon_{il}}(n, \sqrt{\sigma_{ii,ll}}, I_{t-1}) = E(\mathbf{x}_{t+n}/\varepsilon_{ilt} = \sqrt{\sigma_{ii,ll}}, I_{t-1}) - E(\mathbf{x}_{t+n}/I_{t-1}) \tag{22}$$

where $I_t(\mathbf{x}_t, \mathbf{x}_{t-1}, ...)$ is the information set available at time $t-1$, and $\sigma_{ii,ll}$ is

[12] Dees *et al.* (2005) identify only the US monetary shocks in a Global VAR model of 26 countries.

[13] This assumption is tested in Appendix.

the variance of ε_{ilt}. Assuming that $\boldsymbol{\varepsilon}_t$ has a multivariate normal distribution, PSW show that:

$$\begin{aligned} \psi_j(n) &= \frac{1}{\sqrt{\sigma_{ii,ll}}} F^n \mathbf{G}^{-1} \boldsymbol{\Sigma} \mathbf{s}_j, \\ n &= 0, 1, 2, ... \end{aligned} \quad (23)$$

where $\mathbf{s}_j$ is a $k \times 1$ selection vector whose j^{th} element is unity and the remaining elements are zero. Therefore, $\psi_j(n)$ measures the effect of one standard error shock to the j^{th} equation at time t on expected values of $\mathbf{x}$ at time $t+n$.

The GFEVD considers the proportion of the variance of the n-step ahead forecast error of the variable of interest which is explained by conditioning on the non-orthogonalized shocks u_{jt}, $u_{j,t+1}$, ..., $u_{j,t+n}$, for $j = 1, ... k$, while explicitly allowing for the contemporaneous correlations between these shocks and the shocks to the other equations in the system. Like the GIRF, the GFEVD is invariant to the ordering of the variables. The expression for the GFEVD is thus given by:

$$\begin{aligned} \mathbf{GFEVD}(\mathbf{x}_{(l)t;} u_{(j)t}, n) &= \frac{\sigma_{ii}^{-1} \sum_{l=0}^{n} \left(\mathbf{s}_j' F^n \mathbf{G}^{-1} \boldsymbol{\Sigma} \mathbf{s}_i \right)^2}{\sum_{l=0}^{n} \mathbf{s}_j' F^n \mathbf{G}^{-1} \boldsymbol{\Sigma} \mathbf{G}'^{-1} F'^n \mathbf{s}_i} \\ n &= 0, 1, 2, ...; l = 1, ..., k \end{aligned} \quad (24)$$

which gives the proportion of the n-step ahead forecast error variance of the l^{th} element of $\mathbf{x}_t$ accounted for by the innovations in the j^{th} element of $\mathbf{x}_t$. It is important to note that due to the non-diagonal form of the $\boldsymbol{\Sigma}$, the elements of **GFEVD** across j need not sum to unity.

3.3 Conditions for the validity of the GVAR methodology

The GVAR methodology overcomes the difficulties of estimating a large model of the world economy simultaneously by first estimating the country-specific models singularly, and then stacking the coefficients estimates in a Global VAR model for dynamic analysis purposes. Therefore, it is important to emphasize the conditions under which this estimation procedure is indeed equivalent to the simultaneous estimation of the VAR model of the world economy.

1. The global model must be dynamically stable, i.e. the eigenvalues of matrix F in equation (21) lie either on or inside the unit circle.

2. The trade weights, w_{ij}, must be such small that

$$\sum_{j=0}^{N} w_{ij}^2 \to 0, \text{ as } N \to \infty, \text{ for all } i.$$

1. The cross-dependence of the idiosyncratic shocks must be sufficiently small, so that

$$\frac{\sum_{j=0}^{N} \sigma_{ij,ls}}{N} \to 0, \text{ as } N \to \infty, \text{ for all } i, l, \text{ and } s$$

where $\sigma_{ij,ls} = cov(\varepsilon_{ilt}, \varepsilon_{jst})$ is the covariance of the l^{th} variable in country i with the s^{th} variable in country j.

These conditions amount to an econometric formalization of the economic concept of "small open economy".

4 Estimation of the GVAR model

The VECM model with long-run identifying restrictions (16) is estimated for Argentina, Brazil, Mexico, the Euro Area, and Japan. Given the importance of the USA in the global economy, it is sensible to exclude the exogenous foreign variables from this model, which is therefore estimated in unrestricted form. Data are quarterly over the period 1980:1-2003:4. The Euro Area variables are constructed as weighted averages of the corresponding time series of each country in the region, i.e. Austria, Belgium, Finland, France, Germany, Greece, Ireland, Italy, Netherlands, Portugal, and Spain, with weights given by the *per capita* PPP-GDP share of the period 1995-2000.[14] Each VECM model, except for the US one, includes a vector of domestic (endogenous) variables $\mathbf{x}_{it}$ and a vector of foreign country/region-specific variables $\mathbf{x}^*_{it}$; the latter is constructed as a weighted average of the rest of the world (relative to the domestic economy) variables, as detailed in Subsection 3.2. The matrix of trade weights is reported in Table 1, where the 1995 trade shares are reported in column by country/region. This matrix represents the link among economies in the global framework, and shows to which degree each country is involved in bilateral trade.

Table 1. Trade weights.

	Argentina	**Brazil**	**Mexico**	**Euro Area**	**Japan**	**USA**
Argentina	0	0.1648	0.0039	0.0289	0.0031	0.0122
Brazil	0.3749	0	0.0105	0.0751	0.0224	0.0418
Mexico	0.0409	0.0224	0	0.0256	0.0173	0.2185
Euro Area	0.3333	0.3732	0.0621	0	0.2811	0.344
Japan	0.0444	0.1094	0.0335	0.2713	0	0.3836
USA	0.2065	0.3301	0.89	0.5992	0.6762	0

Notes: Trade weights, computed as shares of exports and imports in 1995, are displayed in column by country/region. Each column, but not row, sums to one. Source: Direction of Trade Statistics Yearbook, IMF, 2002.

[14] See Appendix for a description of variables construction and data sources.

Following the theoretical discussion of Section 2, the vector $\mathbf{v}_{it}$ includes the variables $(y_t, sr_t, q_t, nfa_t, y_t^*, sr_t^*, lr_t^*, tt_t, p_t^o)$ for i = Argentina, Brazil and Mexico. For these countries the domestic long-term interest rate is not available over a sufficiently long sample period. For i = Euro Area and Japan, $\mathbf{v}_{it} = (y_t, sr_t, lr_t, q_t, nfa_t, y_t^*, sr_t^*, lr_t^*, tt_t, p_t^o)'$. Finally $\mathbf{z}_{USAt} = (y_t, sr_t, lr_t, q_t, nfa_t, tt_t, p_t^o)'$. All foreign specific variables, as well as the terms of trade and the oil price, are treated as weakly exogenous.

The theoretical model proposed in this paper relies on the assumption that the time series used in the analysis be I(1). Unit root tests are described and discussed in Appendix. Combining the results of all types of tests I cannot unambiguously reject the null hypothesis that all series are I(1). For each country/region a VARX$^*(p_i, q_i)$ model is estimated, with p_i, the lag order of the domestic variables, and q_i, the lag order of the foreign ('star') specific variables, set equal to one following a number of criteria discussed in Appendix. The selected lag order and the inclusion of dummy variables corresponding to outlier values of residuals is sufficient to obtain a good fit of the unrestricted and restricted models to data (see the results of specification tests in Appendix).

The long period considered in this study is likely to include a number of structural breaks for the economies under scrutiny. In particular, this can be the case for countries where financial crises and deep recession have alternated frequently over time, such as the Latin American ones. In order to ascertain the degree to which such economic instability can be a problem for the estimation of the country/region models and of the GVAR, I conduct several parameter constancy tests. The results, described in Appendix, show that though episodes of instability are revealed by the data, overall the estimated parameters can be considered stable over the sample period.

The next step consists of identifying the rank of the cointegration space of the single VECM models. Table 2 reports the Maximum eigenvalue and Trace tests statistics together with their associated 90% and 95% critical values.

Table 2. Cointegration rank statistics

Maximum eigenvalue test						
H_0	**H_1**	Argentina	Brazil	Mexico	**95%**	**90%**
r = 0	r = 1	229.90	229.69	84.71	46.66	43.66
r ≤ 1	r = 2	95.48	**26.34**	61.80	40.12	37.28
r ≤ 2	r = 3	**32.54**	14.30	**12.92**	33.26	30.54
r ≤ 3	r = 4	13.07	7.96	4.25	25.70	23.11
H_0	**H_1**	Euro Area		Japan	**95%**	**90%**
r = 0	r = 1	79.53		92.13	52.63	49.59
r ≤ 1	r = 2	64.01		80.58	46.66	43.66
r ≤ 2	r = 3	**18.84**		**30.05**	40.12	37.28
r ≤ 3	r = 4	17.38		13.01	33.26	30.54
r ≤ 4	r = 5	11.10		6.94	25.70	23.11
H_0	**H_1**		USA		**95%**	**90%**
r = 0	r = 1		59.43		43.72	40.94
r ≤ 1	r = 2		57.00		37.85	35.04
r ≤ 2	r = 3		**26.02**		31.68	29.00
r ≤ 3	r = 4		5.71		24.88	22.53
r ≤ 4	r = 5		2.10		18.08	15.82
Trace test						
H_0	**H_1**	Argentina	Brazil	Mexico	**95%**	**90%**
r = 0	r = 1	370.99	278.29	163.69	107.57	102.48
r ≤ 1	r = 2	141.09	**48.61**	78.97	76.82	72.33
r ≤ 2	r = 3	**45.61**	22.26	**17.17**	49.52	46.10
r ≤ 3	r = 4	13.07	7.96	4.25	25.70	23.11
H_0	**H_1**	Euro Area		Japan	**95%**	**90%**
r = 0	r = 1	190.86		222.72	141.17	135.76
r ≤ 1	r = 2	111.33		130.58	107.57	102.48
r ≤ 2	r = 3	**47.32**		**50.01**	76.82	72.33
r ≤ 3	r ≤ 3	28.48		19.96	49.52	46.10
r ≤ 4	r ≤ 5	11.10		6.94	25.70	23.11
H_0	**H_1**		USA		**95%**	**90%**
r = 0	r = 1		150.27		108.90	103.71
r ≤ 1	r = 2		90.83		81.20	76.68
r ≤ 2	r = 3		**33.83**		56.43	52.71
r ≤ 3	r = 4		7.81		35.37	32.51
r ≤ 4	r = 5		2.10		18.08	15.82

Notes: the last two columns report the critical values at the 95% and 90% significance level. Statistics in bold indicate acceptance of the null hypothesis at the 5% significance level.

Both tests select unambiguously cointegration rank 1 for Brazil, and 2 for all other countries. This is consistent with the theoretical framework developed in Section 2 suggesting the existence of at most two possible long-run relations among the selected variables. Accordingly, I set 1 cointegrating relationship for Brazil and 2 for all other models.

4.1 Long-run

I now turn to investigating whether the cointegrating vectors of the Latin American countries, along with Euro Area and Japan are identified in terms of the two long-run structural relationships discussed in Section 2. Given the available sample size, the size of the underlying VAR models and the number of long-run restrictions, I use bootstrap techniques to test the significance of the log-likelihood ratio (LR) statistics for jointly testing the two over-identifying restrictions.[15] The parametric bootstrap estimation of the cointegrating vectors is based on 99 replications. As detailed below, the LR test results provide a weak support for the over-identifying restrictions on the cointegration space of the VECM models. However, given the strong theoretical priors linking the NFA positions to the other variables of the model, I choose to base the main body of the long- and short-run analysis on the restricted model. Nevertheless, I am conscious that the conclusions of the paper may be at least partially driven by the assumptions imposed on the long-run structure. In order to address this problem, I provide in Section 5 a robustness check of the main results by discussing the results obtained from the unrestricted version of the model.

I describe the results for each country in turn, but given the focus of this study I will linger only briefly on Euro Area and Japan, mainly as comparison terms.

4.1.1 Argentina

The LR statistic for testing the over-identifying restrictions on the cointegration space of the Argentina's model takes the value 44.92 with an empirical p-value of 1.01%. The theoretical restrictions are thus rejected at the standard level of significance, 5%, but cannot be rejected at 1% level. The estimated long-run relations are:

$$\begin{aligned}\widehat{\xi}_{1,t+1} &= \underset{(0.24)}{0.29}\, y_t + \underset{(0.06)}{1.65}\, sr_t + \underset{(0.05)}{0.40}\, q_t + nfa \\ &\quad - \underset{(0.34)}{0.92}\, y_t^* - \underset{(0.41)}{3.81}\, sr_t^* + \underset{(3.46)}{0.87}\, lr_t^* + \underset{(0.00)}{0.01}\, t \qquad (25)\\ \widehat{\xi}_{2,t+1} &= \underset{(0.23)}{-0.96}\, (y_t - y_t^*) + q_t + \underset{(0.17)}{3.03}\, nfa_t \underset{(0.35)}{-1.36}\, tt_t \qquad (26)\end{aligned}$$

where the asymptotic standard errors are in parentheses.

Equation (25) is the long-run solvency condition for NFA positions. The coefficient of foreign *per capita* output is small but significant, while that of domestic output is insignificant. This suggests that the NFA are associated in the long-run more to foreign than domestic economic conditions. The signs support the portfolio balance approach: an increase in domestic output should lead to an increase in money demand for transactions purposes and to a decrease in

[15] The use of bootstrap methods is suggested, among others, by Gredenhoff and Jacobson (1998) who show to which extent the chi-squared tests are biased in cointegrating VAR models with small samples.

demand for foreign assets, while the inverse is true for foreign output. Following Lane and Milesi-Ferretti (2001b), this result can also be interpreted as the effect of a larger access to international financial markets following the relaxation of credit constraints. Similarly, the increase in domestic output can be associated with an improved creditworthiness on international financial markets and an increased capability of attracting foreign capital.

Both domestic and global real interest rates coefficients are large and highly significant, with the foreign rate coefficient taking a higher value than the domestic counterpart and thus supporting the view expressed, among others, by Calvo *et al.* (1993, 1996). Again, the signs are in line with the portfolio balance theory: higher domestic rates are associated with larger net liabilities, while the contrary is true for foreign short-term rates. The long-term rate coefficient is insignificant and small compared to short-term rates. This may indicate that the NFA bond and equity components, more reactive to short-term rates, are larger than the FDI component, more responsive to long-term returns. Overall, foreign conditions seem to play a larger role than domestic counterparts in determining investment decisions in Argentina in the long-run.

The second cointegrating vector, equation (26), relates to the long-run determinants of the real exchange rate. The relation between the real exchange rate and relative output is statistically significant but has the wrong sign; Lane and Milesi-Ferretti (2004) obtain similar results for developing countries, in contrast to industrial countries. Analogously, the terms of trade and the real exchange rate are positively correlated, contradicting theory. Finally, the relationship between NFA and the real exchange rate is significant and rightly signed.

4.1.2 Brazil

Since the main interest of this study is on the determinants of NFA, the rank 1 cointegration space of the Brazilian model is identified in terms of the first theoretical relation. The LR statistic takes the value 7.95 with an empirical p-value of 5.05% supporting the theoretical restriction imposed. The estimated long-run relationship follows:

$$\begin{aligned} \widehat{\xi}_{1,t+1} &= \underset{(0.26)}{0.53}\, y_t + \underset{(0.24)}{6.95}\, sr_t + \underset{(0.17)}{1.19}\, q_t + nfa_t \\ &\quad + \underset{(1.74)}{2.14}\, y_t^* - \underset{(1.00)}{4.09}\, sr_t^* + \underset{(8.93)}{3.96}\, lr_t^* - \underset{(0.01)}{0.02}\, t \end{aligned} \tag{27}$$

Equation (27) shows a negative relationship between the NFA position and domestic output, consistently with the portfolio balance theory and with the argument proposed by Lane and Milesi-Ferretti (2001b) based on a larger access to international financial markets and improved creditworthiness. The negative relationship between foreign output and NFA is insignificant. The coefficients associated with domestic and foreign short-term real rates, large and significant, are consistent with a portfolio balance argument. As in the Argentina's model, the coefficient of the long-term real interest rate is insignificant.

4.1.3 Mexico

The LR statistic for the Mexican model is 53.08 with an empirical p-value of 1.01%. Again, likewise the Argentinean model, the theoretical restrictions are rejected at the standard level of significance, but cannot be rejected at 1% level. The estimated long-run relations are:

$$\begin{aligned}
\widehat{\xi}_{1,t+1} &= \underset{(0.55)}{-0.93} y_t - \underset{(0.60)}{5.11}\, sr_t - \underset{(0.16)}{0.19}\, q_t + nfa_t \\
&\quad + \underset{(0.88)}{0.83}\, y_t^* + \underset{(3.26)}{6.11}\, sr_t^* - \underset{(3.57)}{5.26}\, lr_t^* + \underset{(0.005)}{0.0001} t \qquad (28)\\
\widehat{\xi}_{2,t+1} &= \underset{(0.70)}{-5.01}\,(y_t - y_t^*) + q_t + \underset{(0.55)}{1.44}\, nfa_t + \underset{(0.52)}{3.04}\, tt_t \qquad (29)
\end{aligned}$$

The first relation shows that the domestic output coefficient is larger than the foreign one and none is statistically significant. Unlike the Argentina's and Brazil's model, the relation between NFA and domestic output is positive. Lane and Milesi-Ferretti (2001b) suggest that an increase in domestic output relative to foreign one can be associated to a decline in the relative marginal product of capital along with a decline in domestic investment, both determining an improvement of NFA and an outflow of payments. Alternatively, an increase in domestic permanent output may be associated to higher savings and an accumulation of foreign assets. The domestic short-term interest rate coefficient is large and significant, but its sign does not support the portfolio balance theory. The reason for a positive relationship between NFA and domestic short-term rates of return may be the risk premium component reflecting the solvency concerns of international investors: an increase of interest rates following a reduced credibility of Mexican debtors can induce an outflow of capital and thus a positive relation between NFA and interest rates. Tight monetary policies adopted during crisis episodes as a measure to stop capital flights is a complementary explanation of this positive correlation. Both the foreign short- and long-term interest rates are insignificant. The long-run association of Mexican NFA with domestic factors is consistent with the findings by Fernandez-Arias (1996).

Equation (29) does not support the negative long-run correlation between the real exchange rate and relative *per capita* output, while the signs of the NFA and the terms of trade coefficients support the transfer effect theory.

4.1.4 Industrial countries

As a matter of comparison I discuss briefly the estimated cointegrating vectors of the Euro Area's and Japan's model.

Euro Area:

$$\begin{aligned}\widehat{\xi}_{1,t+1} &= \underset{(0.14)}{1.91}\,y_t - \underset{(0.68)}{4.42}\,sr_t - \underset{(0.73)}{0.68}\,lr_t + \underset{(0.03)}{0.15}\,q_t + nfa_t \\ &\quad - \underset{(0.14)}{1.48}\,y_t^* + \underset{(0.26)}{0.03}\,sr_t^* + \underset{(0.76)}{1.78}\,lr_t^* - \underset{(0.001)}{0.003}\,t \qquad (30) \\ \widehat{\xi}_{2,t+1} &= \underset{(0.71)}{11.91}\,(y_t - y_t^*) + q_t + \underset{(0.74)}{8.45}\,nfa_t - \underset{(0.40)}{0.85}\,tt_t \qquad (31)\end{aligned}$$

Japan:

$$\begin{aligned}\widehat{\xi}_{1,t+1} &= \underset{(0.37)}{1.09}\,y_t - \underset{(3.17)}{6.59}\,sr_t + \underset{(3.67)}{35.05}lr_t + \underset{(0.12)}{0.18}\,q_t + nfa_t \\ &\quad + \underset{(0.91)}{1.35}\,y_t^* - \underset{(3.39)}{4.24}\,sr_t^* + \underset{(3.44)}{0.16}\,lr_t^* - \underset{(0.0051)}{0.014}\,t \qquad (32) \\ \widehat{\xi}_{2,t+1} &= \underset{(0.29)}{0.21}\,(y_t - y_t^*) + q_t - \underset{(0.20)}{1.12}\,nfa_t + \underset{(0.10)}{1.08}\,tt_t \qquad (33)\end{aligned}$$

In both models domestic output is negatively related to the NFA position, with a statistically significant coefficient. It seems highly implausible that this is due to a larger access to international markets following higher level of output, a more sensible explanation for developing than for industrial countries. Here the portfolio balance mechanism and the effect of improved economic conditions on expected investments profitability might be at work. This reasoning is supported by the sign of the foreign output coefficient in the Euro Area's model, while that of the Japan's model is insignificant. The positive relation between short-term real interest rates and NFA is puzzling from the point of view of the portfolio balance approach. However, a possible explanation lies in the measurement of NFA that takes into account valuation changes: an increase in the real interest rate might reflect a decrease in the price of existing stocks and thus stimulate investors to acquire foreign assets. This effect is more likely to occur in the short-run when asset price changes prevail on longer-term profitability expectations. The negative relationship between long-term real interest rates and NFA for the Euro Area's model support this view. The other interest rates coefficients are insignificant. The transfer effect hypothesis is strongly supported in the Euro Area's model, except for the terms of trade. By contrast, the transfer effect is supported in the Japan's model only by the terms of trade coefficient, while the relative output coefficient is insignificant and the NFA coefficient sign is opposite to what expected.

4.1.5 Persistence profiles

The persistence profiles of both long-run relations, shown in Figures 1 and 2, converge to zero for all countries, supporting the hypothesis that the portfolio balance theory and the transfer effect indeed represent cointegrating relations.

Figure 1
Persistence profiles of the first long-run relation

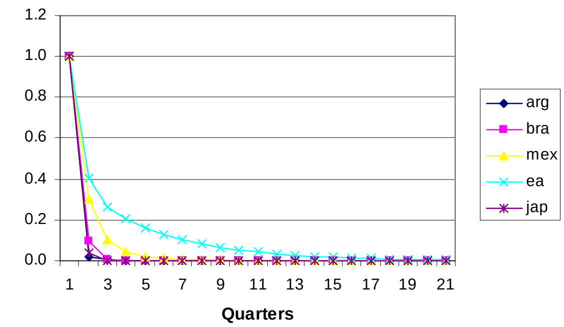
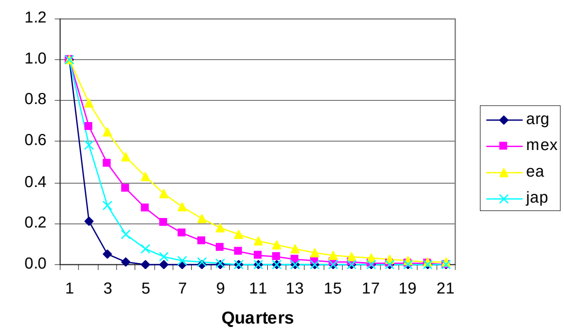


Figure 2
Persistence profiles of the second long-run relation



There is no evidence of overshooting in the profiles. The first long-run relation (portfolio balance) completes its adjustment within 5 quarters for all countries except the Euro Area, where it is very persistent and takes more than 3 years to revert to equilibrium after being shocked. This may imply that in the Euro Area any variable causing a departure from equilibrium will take a long time before being corrected, unlike in Latin America. This result is in line with what observed in this region over the last twenty years, when foreign capital has been subject to sharp reversals and when interest and exchange rates have shifted suddenly in response to finance shocks. It may also mean that international investors are less prone to bear an unbalanced portfolio with respect to fundamentals when it comes to Latin American countries, and they try to revert to fundamental-based NFA positions after a shock. This is also true for Japan, whose cointegrating relation follows the same pattern as Latin American ones. The second long-run relation shows on average a higher persistence than the first one, with the Euro Area relation being again the most sluggish. The Argentinean real exchange rate relation takes little more than one year to revert to equilibrium while the Japanese one takes more than two years. The Mexican and European relations departures from equilibrium are corrected after almost 4 years.

To sum up, the long-run analysis shows that NFA are mainly associated to external economic conditions for Argentina and Mexico, while for Brazil this is only true for output, but not interest rates. Short-term interest rates coefficients are the largest among all factors, and mostly significant. Latin American equilibrium relations appear less persistent than industrial countries' ones.

4.2 Short-run

The short-run dynamics of the model are characterized by the error correction specification of the single country/region models given in Tables 3a and 3b.

Table 3a VECM model estimates: Latin American countries

			Argentina		
Equation	Δy_t	Δsr_t	Δlr_t	Δq_t	Δnfa_t
Intercept	-0.105 (0.026)	-1.339 (0.054)	-	-0.357 (0.054)	-0.180 (0.068)
$\xi_{1\tau}$	-0.049 (0.001)	-0.556 (0.020)	-	-0.138 (0.020)	-0.015 (0.025)
$\xi_{2\tau}$	0.026 (0.009)	0.142 (0.018)	-	0.016 (0.018)	-0.157 (0.022)
Δy^*_t	-0.517 (0.161)	0.399 (0.335)	-	-0.165 (0.333)	0.675 (0.419)
Δsr^*_t	0.230 (0.052)	2.493 (0.108)	-	-0.198 (0.108)	-0.036 (0.135)
Δlr^*_t	5.906 (1.619)	-6.422 (3.357)	-	-0.342 (3.342)	0.885 (4.199)
Δtt_t	0.216 (0.040)	-0.001 (0.082)	-	-0.260 (0.082)	0.176 (0.103)
Δoil_t	-0.006 (0.027)	0.097 (0.056)	-	0.019 (0.056)	0.136 (0.070)
sc1	-0.083 (0.013)	-0.008 (0.026)	-	0.05 (0.026)	0.056 (0.033)
ARG84q3	1.046 (0.093)	0.12 (0.192)	-	0.123 (0.191)	0.561 (0.240)
ARG86q1	-0.021 (0.019)	1.204 (0.040)	-	0.137 (0.039)	-0.067 (0.049)
ARG82q3	0.076 (0.034)	-0.009 (0.070)	-	1.139 (0.070)	-0.392 (0.087)
ARG84q4	-0.046 (0.058)	-0.142 (0.121)	-	0.028 (0.120)	1.153 (0.151)
R^2	0.73	0.97	-	0.84	0.66
			Brazil		
Equation	Δy_t	Δsr_t	Δlr_t	Δq_t	Δnfa_t
Intercept	-0.002 (0.049)	1.299 (0.046)	-	0.026 (0.066)	0.000 (0.039)
$\xi_{1\tau}$	0.000 (0.005)	-0.123 (0.004)	-	-0.003 (0.006)	-0.000 (0.004)
Δy^*_t	-0.845 (0.374)	0.226 (0.345)	-	-0.666 (0.501)	0.485 (0.292)
Δsr^*_t	0.389 (0.062)	1.286 (0.057)	-	-0.307 (0.083)	0.092 (0.048)
Δlr^*_t	0.233 (1.434)	1.618 (1.322)	-	3.548 (1.922)	-1.724 (1.119)
Δtt_t	0.113 (0.032)	0.056 (0.029)	-	0.048 (0.042)	0.029 (0.025)
Δoil_t	-0.021 (0.029)	-0.067 (0.027)	-	-0.065 (0.039)	0.006 (0.022)
sc1	-0.077 (0.012)	0.045 (0.011)	-	-0.044 (0.156)	-0.024 (0.009)
sc2	0.060 (0.011)	-0.021 (0.010)	-	-0.036 (0.014)	0.036 (0.008)
BRA87q2	0.226 (0.045)	1.320 (0.042)	-	0.168 (0.061)	0.067 (0.035)
BRA82q3	0.069 (0.063)	-0.065 (0.058)	-	1.069 (0.084)	-0.354 (0.049)
BRA83q1	1.129 (0.112)	-0.078 (0.103)	-	0.026 (0.150)	0.378 (0.087)
R^2	0.79	0.97	-	0.70	0.62
			Mexico		
Equation	Δy_t	Δsr_t	Δlr_t	Δq_t	Δnfa_t
Intercept	0.008 (0.013)	0.064 (0.010)	-	0.103 (0.022)	-0.056 (0.018)
$\xi_{1\tau}$	-0.011 (0.013)	0.073 (0.010)	-	0.144 (0.022)	-0.081 (0.018)
$\xi_{2\tau}$	0.030 (0.007)	-0.010 (0.005)	-	-0.055 (0.012)	0.036 (0.010)
Δy^*_t	0.831 (0.292)	0.619 (0.237)	-	-1.065 (0.513)	0.632 (0.422)
Δsr^*_t	-0.948 (0.561)	0.696 (0.455)	-	1.666 (0.985)	-1.078 (0.810)
Δlr^*_t	1.010 (0.822)	-1.252 (0.667)	-	-2.564 (1.444)	1.058 (1.187)
Δtt_t	0.010 (0.030)	-0.056 (0.024)	-	-0.070 (0.052)	0.014 (0.043)
Δoil_t	0.011 (0.014)	-0.016 (0.011)	-	0.022 (0.025)	0.051 (0.020)
sc1	-0.070 (0.006)	0.031 (0.005)	-	-0.042 (0.011)	-0.032 (0.009)
sc2	-0.024 (0.006)	0.030 (0.005)	-	-0.018 (0.010)	-0.001 (0.008)
sc3	-0.090 (0.006)	0.015 (0.005)	-	-0.008 (0.010)	-0.033 (0.008)
MEX82q1	-0.220 (0.088)	1.066 (0.071)	-	-1.266 (0.307)	0.307 (0.127)
MEX82q3	-0.007 (0.040)	0.025 (0.032)	-	1.113 (0.070)	-0.494 (0.058)
R^2	0.84	0.85	-	0.82	0.71

Notes: estimated asymptotic standard errors are in brackets.

Table 3b VECM model estimates: Developed countries

Euro Area					
Equation	Δy_t	Δsr_t	Δlr_t	Δq_t	Δnfa_t
Intercept	0.202 (0.038)	-0.191 (0.022)	-0.136 (0.031)	-0.154 (0.210)	0.022 (0.032)
$\xi_{1\tau}$	-0.116 (0.022)	0.111 (0.013)	0.079 (0.018)	0.087 (0.121)	-0.014 (0.018)
$\xi_{2\tau}$	0.010 (0.004)	-0.015 (0.002)	-0.009 (0.003)	-0.027 (0.020)	-0.002 (0.003)
Δy^*_t	-0.173 (0.082)	0.087 (0.049)	0.173 (0.067)	0.253 (0.456)	-0.024 (0.069)
Δsr^*_t	0.017 (0.027)	0.036 (0.016)	0.001 (0.022)	-0.159 (0.152)	-0.016 (0.023)
Δlr^*_t	0.380 (0.194)	0.322 (0.115)	0.395 (0.158)	-0.431 (1.079)	-0.298 (0.163)
Δtt_t	0.082 (0.062)	0.071 (0.037)	0.046 (0.051)	-2.565 (0.348)	0.041 (0.052)
Δoil_t	0.017 (0.005)	-0.003 (0.003)	-0.006 (0.004)	-0.083 (0.026)	0.006 (0.004)
sc1	-0.018 (0.002)	0.004 (0.001)	0.005 (0.002)	-0.009 (0.012)	-0.006 (0.002)
sc2	-0.006 (0.002)	0.002 (0.001)	0.005 (0.001)	0.009 (0.010)	0.001 (0.001)
sc3	-0.019 (0.002)	0.004 (0.001)	0.003 (0.001)	-0.004 (0.009)	-0.002 (0.001)
EA87q1	-0.251 (0.066)	0.005 (0.039)	0.075 (0.054)	-0.526 (0.370)	1.089 (0.370)
EA80q4	0.027 (0.174)	1.027 (0.104)	0.872 (0.142)	-2.243 (0.971)	0.038 (0.146)
EA83q4	1.051 (0.078)	-0.120 (0.046)	-0.100 (0.063)	-0.346 (0.433)	0.151 (0.065)
R^2	0.85	0.72	0.56	0.45	0.84

Japan					
Equation	Δy_t	Δsr_t	Δlr_t	Δq_t	Δnfa_t
Intercept	-0.063 (0.063)	0.263 (0.048)	0.386 (0.033)	-0.582 (0.297)	-0.028 (0.069)
$\xi_{1\tau}$	0.005 (0.006)	-0.025 (0.005)	-0.037 (0.003)	0.051 (0.028)	0.003 (0.007)
$\xi_{2\tau}$	-0.058 (0.017)	0.006 (0.013)	0.001 (0.009)	-0.256 (0.080)	0.003 (0.018)
Δy^*_t	0.008 (0.155)	-0.132 (0.119)	-0.154 (0.082)	-0.195 (0.736)	0.064 (0.736)
Δsr^*_t	-0.264 (0.168)	0.008 (0.129)	0.112 (0.089)	-1.229 (0.797)	-0.386 (0.185)
Δlr^*_t	-0.015 (0.319)	0.214 (0.245)	-0.152 (0.170)	0.814 (1.514)	0.437 (0.352)
Δtt_t	-0.023 (0.031)	0.055 (0.024)	0.042 (0.017)	-0.754 (0.148)	-0.089 (0.034)
Δoil_t	0.017 (0.008)	0.000 (0.006)	-0.000 (0.004)	-0.051 (0.039)	-0.015 (0.009)
sc1	0.002 (0.003)	-0.012 (0.002)	-0.012 (0.002)	0.002 (0.016)	0.022 (0.004)
sc3	0.002 (0.003)	-0.006 (0.002)	-0.007 (0.001)	-0.006 (0.014)	-0.003 (0.003)
JAP85q1	0.033 (0.070)	0.042 (0.054)	0.016 (0.037)	0.495 (0.333)	1.031 (0.077)
R^2	0.19	0.64	0.85	0.35	0.78

USA					
Equation	Δy_t	Δsr_t	Δlr_t	Δq_t	Δnfa_t
Intercept	-0.131 (0.059)	-0.071 (0.040)	0.157 (0.033)	-0.690 (0.244)	-0.011 (0.005)
$\xi_{1\tau}$	0.002 (0.001)	0.001 (0.001)	-0.002 (0.000)	0.010 (0.003)	-0.000 (0.000)
$\xi_{2\tau}$	-0.001 (0.001)	-0.003 (0.001)	-0.003 (0.000)	0.001 (0.003)	0.000 (0.000)
Δtt_t	0.133 (0.070)	-0.009 (0.047)	-0.049 (0.039)	-0.604 (0.286)	-0.006 (0.006)
Δoil_t	0.007 (0.008)	-0.003 (0.005)	-0.005 (0.005)	-0.054 (0.033)	-0.000 (0.001)
sc1	-0.002 (0.002)	-0.001 (0.001)	-0.001 (0.001)	-0.014 (0.008)	-0.012 (0.000)
US83q1	0.084 (0.077)	0.075 (0.052)	0.031 (0.043)	0.408 (0.317)	1.073 (0.006)
US82q1	0.767 (1.114)	0.299 (0.745)	-1.073 (0.620)	-1.307 (4.565)	1.162 (0.090)
R^2	0.10	0.33	0.54	0.15	0.99

Notes: estimated asymptotic standard errors are in brackets.

The estimated coefficients of the error correction terms (also known as loading coefficients) are statistically significant in most equations showing the existence of strong interactions and feedbacks among the model variables. Only in Mexico, however, the loading coefficient of the first error correction term of the NFA equation is statistically significant. This suggests that NFA rarely bear the adjustment following perturbations to the first long-run equilibrium relationship. Since the loading coefficients of the NFA equations are small in developed countries' models too, it is difficult to argue that their insignificance in Latin American countries be due to restricted access of emerging economies to international financial markets. This implies that frictions of diverse nature still impede capital movements to some extent, despite the increasing world financial integration.

In most equations the loading coefficients of short-term real interest rates and the real exchange rate are statistically significant and large, suggesting that these factors are in charge of providing most of the adjustment required to bring the system back to the equilibrium.

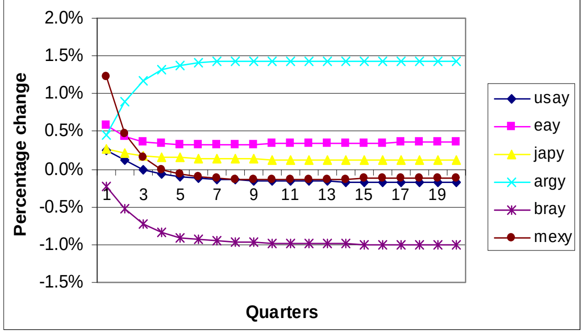
4.2.1 Impulse response functions

In this Subsection, I analyse the effect of shocks to domestic and foreign variables by examining the time profile of generalized impulse response functions (GIRF). Specifically I focus on the responses of Latin American countries' NFA positions to one standard deviation shocks to *per capita* output and interest rates.

I emphasize that although this approach does not allow for the interpretation of disturbances according to their "structural" economic nature, nevertheless it is particularly suited for the analysis of the transmission of shocks across regions. In fact, since the country-specific models are estimated conditional on weakly exogenous foreign variables, it remains only a modest degree of correlations among the same shocks across different regions. This amounts to the identification of shocks according to their geographical origin.

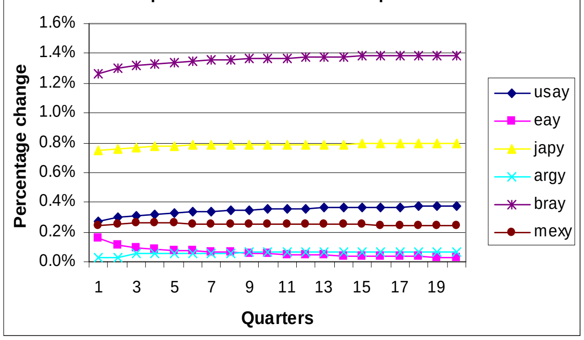
Shocks to *per capita* output equations Figure 3 shows that Argentina's NFA are much more sensitive to domestic output shocks than to foreign ones, at least in the medium-run.

Figure 3
Response of Argentina's NFA to output shocks



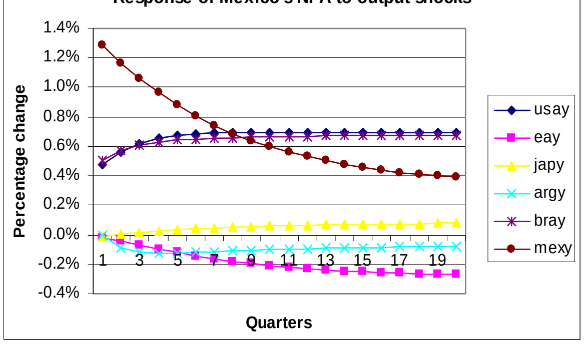
A one standard deviation positive shock to domestic output produces, on impact, a 0.5% increase of NFA, with a cumulative effect slightly below 1.5% after 6 quarters. The Brazilian output has a cumulative negative effect of 1% after 10 quarters, while the rest of foreign output shocks exerts an effect comprised between -0.1% and 0.5%, the only remarkable exception being the Mexican shock (1.2%). Among the industrial countries, the European output has the largest effect ranging from 0.6% on impact to 0.3% after 5 years. Figure 4 draws a broadly similar picture for Brazil.

Figure 4
Response of Brazil's NFA to output shocks



A shock to domestic output, in fact, produces the largest response both on impact (1.2%) and in the medium term (1.4%). The Japanese output shock has the second largest effect over all horizons, while other impulses produce a response ranging between 0% and 0.4%, with Argentina and the Euro Area generating the smallest responses. Figure 5 reveals some differences in the behaviour of Mexico's NFA with respect to other Latin American countries.

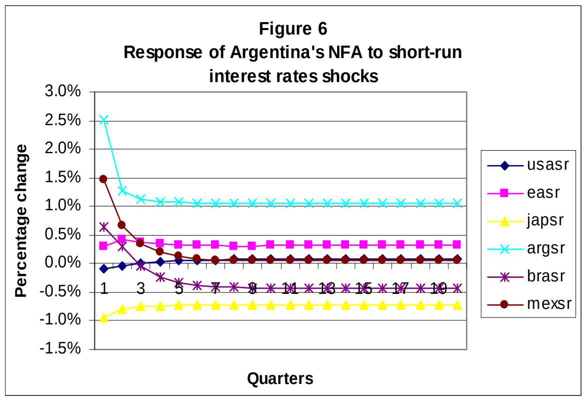
Figure 5
Response of Mexico's NFA to output shocks



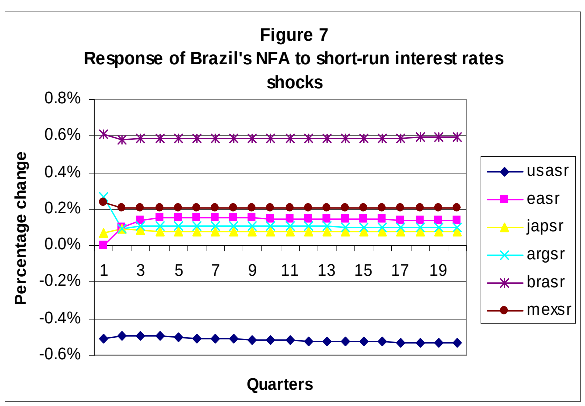
Over a 2 years horizon domestic output exerts the largest effect, with a 1.3% impact. However, beyond that point the responses to US and Brazilian output shocks become larger and cumulate to 0.7% in the medium-run. The other responses remain small over all horizons, the largest being that to the European shock.

To summarize, domestic output shocks' effects are much larger than foreign ones for Argentina and Brazil, while this is true for Mexico only over a 2 years horizon. The European, Japanese, and US outputs produce the largest responses respectively in Argentina, Brazil and Mexico.

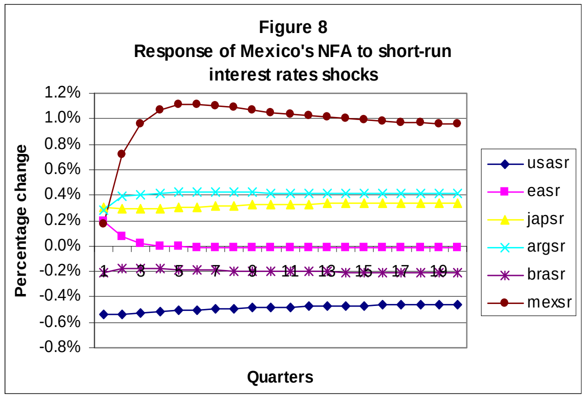
Shocks to short-term real interest rate equations Figure 6 shows that the largest response of Argentina's NFA is to its own interest rate shock, both on impact (2.5%) and in the medium-term (1% after 6 quarters).



A typical positive shock to domestic short-term interest rate turns Argentina to a net debtor and thus drives capital out. Positive shocks to Brazil's and Mexico's short-term rates have on impact the same qualitative effects (outflow of capital) on Argentina. Among the foreign shocks, the Japanese interest rate generates the largest response, while the response to the US shock is negligible at all horizons. Also Brazil's NFA (Figure 7) react mostly to domestic interest rates (0.6% over all horizons), though the response is weaker than Argentina's one.



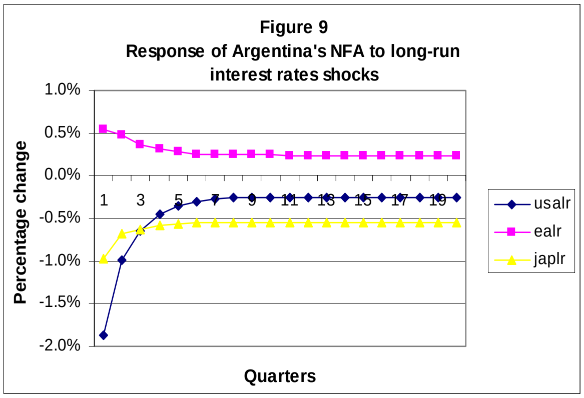
The largest foreign influence is exerted by US rates, that produce an average negative (capital inflow) position of 0.5%. Figure 8 reveals a similar response of Mexico's NFA to interest rates disturbances.

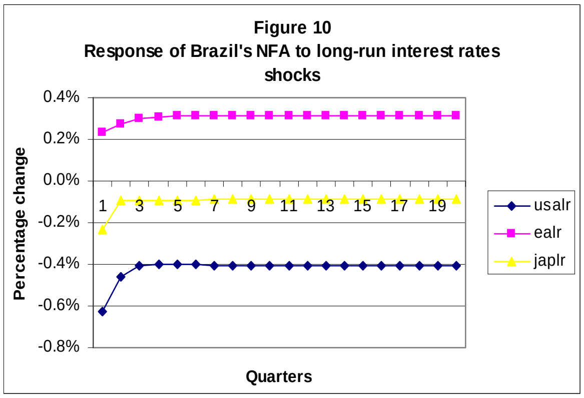


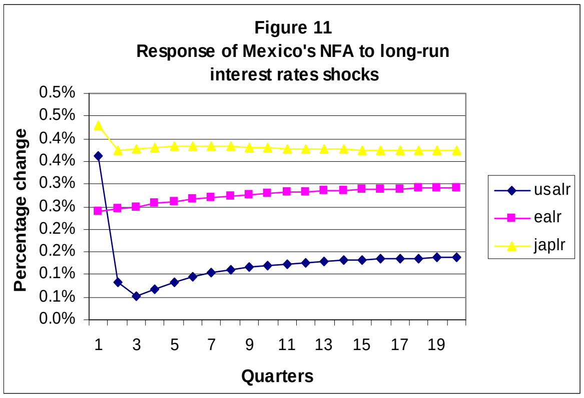
The domestic shock dominates by far all others from the second quarter on, with a medium-term positive response reaching 1%. The most important foreign shock is the US one, averaging 0.5% at all horizons. To sum up, domestic interest rates shocks generate larger responses of Latin American NFA positions compared to foreign ones. USA's interest rates are the most important foreign shocks for Brazil and Mexico, while the Japanese rates produce the largest response in Argentina.

Agénor (1998) argues that in many developing countries expectations of lower inflation (typically associated with stabilization policies and liberalized financial markets) have been followed by higher domestic nominal interest rates and capital inflows in order to accommodate sharp increases in domestic real money balances. This does not appear to be the mechanism underlying the relationship between the short-term interest rate and capital flows for the countries considered in this paper. In fact, given the way the real interest rates are constructed, lower inflationary expectations and higher nominal interest rates should lead to higher real interest rates and, thus, to capital inflows, but the results discussed above point to the opposite direction: an increase in domestic rates drives capital out. This could be due to the fact that the higher interest rates are mainly associated with higher risk premiums and financial distress risk. This interpretation is in line with the negative relation between Latin American NFA and industrial countries' interest rates, which, on the other hand, seems at odds with the evidence presented by Fernàndez-Arias (1996), Calvo *et al.* (1993, 1996), and Frankel and Okongwu (1996).

Shocks to long-term real interest rate equations Figure 9, 10, and 11 display the responses of Latin American countries' NFA to foreign long-term real interest rates shocks.

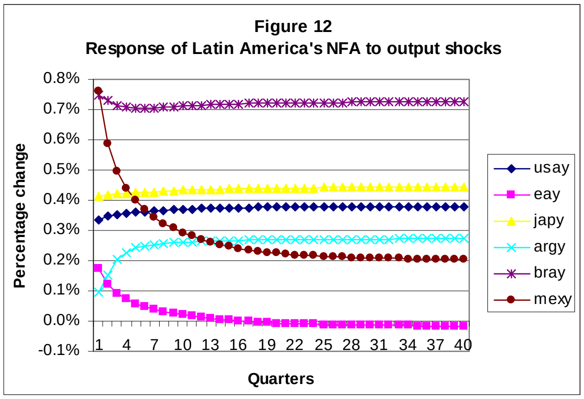


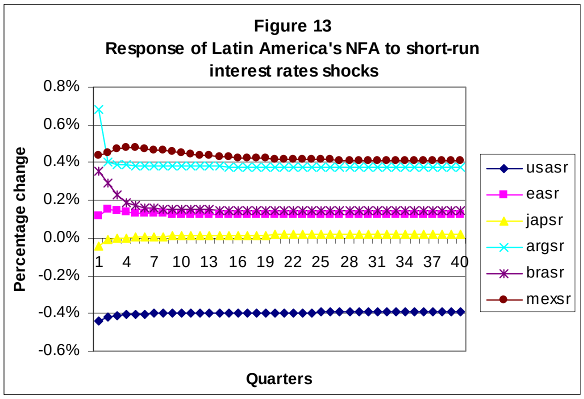


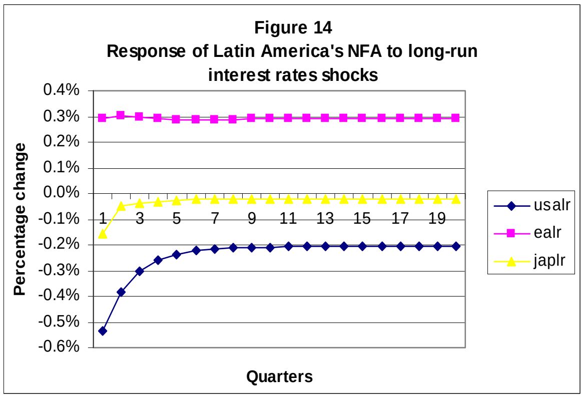


The effects are noticeable in all countries, and especially in Argentina on which US and Japanese rates exert the largest effect on impact. Somehow counter intuitively, the US rates have the smallest effect on Mexican NFA from the second quarter onwards.

Regional responses Figures 12, 13, and 14 show the effects of output and interest rates' shocks on the regional NFA. The responses are computed as weighted averages of single Latin American countries' responses with weights given by the shares of Purchasing Power Parity GDP over the period 1995-2000.







Among the industrial countries, the largest effect is exerted by the Japanese and US output, respectively, while the European output affects the regional NFA only on impact. The Brazilian output, however, has larger effects than industrial economies both on impact and in the long-term. The US short-term real interest rates exert the largest influence among the industrial countries, but the effect of Argentinean and Brazilian rates are of the same magnitude. The US long-term real interest rates have by far the largest effect on impact, but after only 3 quarters they are outweighed by European rates. It is interesting to note that European and US real interest rates, both short- and long-term, exert opposite influence on Latin American NFA: an increase in the former makes capital outflow, while an increase in the latter raise regional liabilities.

Symmetry of responses Table 4 reports the cross-country correlation coefficients of Latin American NFA responses to selected external shocks.

Table 4. Correlations of NFA responses.

Shock	**Arg/Bra**	**Arg/Mex**	**Bra/Mex**
usay	-0.91	-0.99	0.84
eay	0.50	0.34	0.97
japy	-0.97	-0.93	0.98
usasr	-0.79	0.91	-0.97
easr	-0.03	0.18	-0.94
japsr	0.20	0.51	-0.41
usalr	0.95	-0.67	-0.87
ealr	-0.96	-0.86	0.74
japlr	0.96	-0.88	-0.97

Notes: cross-country correlations of the responses of NFA to external shocks.

Large and positive correlation coefficients imply that, when hit by external shocks, the NFA positions tend to react independently of domestic specific conditions. The opposite is true for large and negative coefficients, as well as for small coefficients. The pattern of correlation is not clear cut. For example, the large and positive correlation coefficients of the responses of Brazilian and Mexican NFA positions to industrial countries output shocks (first three figures of the last column of Table 4) denote a symmetric effect of these shocks on Brazil and Mexico. However, the large negative coefficients in the same column show that the NFA position in Brazil and Mexico reacts asymmetrically to short-term real interest rates shocks in US and Euro Area, and to long-term real interest rates shocks in US and Japan. Argentinean and Mexican NFA react asymmetrically to US and Japanese output and long-term rates shocks, and symmetrically to the same countries short-term rates shocks. Finally, Argentinean and Brazilian responses to output shocks are negatively correlated for US and Japan, and positively correlated for the Euro Area. The opposite is true for the responses to long-term rates shocks, while the effect of short-term real interest rates is milder.

4.2.2 Forecast error variance decomposition

Tables 5, 6, and 7 report the generalized forecast error variance decomposition (GFEVD) of Latin American countries NFA positions in terms of internal and external factors.

Table 5. Generalized variance decomposition of the forecast error of Argentina's NFA

Period	External shocks by country					Domestic shocks				All external shocks	All domestic shocks
	USA	EA	JAP	BRA	MEX	ARGy	ARGsr	ARGrer	ARGnfa		
1	0.71	0.08	0.11	1.35	0.53	0.14	4.82	9.11	97.55	2.78	111.62
4	0.61	0.06	0.17	1.39	0.10	0.64	5.30	16.60	87.56	2.34	110.11
8	0.50	0.05	0.22	1.75	0.08	1.19	5.81	23.35	75.00	2.60	105.35
12	0.42	0.04	0.26	2.02	0.07	1.55	6.14	27.78	66.40	2.82	101.88
20	0.33	0.04	0.30	2.37	0.05	1.99	6.54	33.21	55.85	3.09	97.60
40	0.23	0.02	0.35	2.77	0.03	2.48	6.99	39.30	44.01	3.41	92.78

Notes: percentage of the k-step ahead forecast error variance explained by the shock on the corresponding column. Percentages do not sum to 100 due to non-zero covariance between the shocks.

Table 6. Generalized variance decomposition of the forecast error of Brazil's NFA

Period	External shocks by country					Domestic shocks				All external shocks	All domestic shocks
	USA	EA	JAP	ARG	MEX	BRAY	BRAsr	BRArer	BRAnfa		
1	3.63	0.20	2.53	65.82	4.42	32.80	6.04	70.41	97.59	76.61	206.84
4	3.69	0.24	2.74	62.10	4.75	34.25	5.92	67.36	97.84	73.52	205.38
8	3.77	0.26	2.89	60.95	4.85	35.10	5.88	65.99	97.87	72.73	204.83
12	3.83	0.28	2.99	60.27	4.88	35.62	5.85	65.28	97.85	72.25	204.61
20	3.89	0.29	3.11	59.42	4.90	36.26	5.82	64.48	97.81	71.61	204.38
40	3.97	0.31	3.23	58.44	4.90	36.98	5.79	63.56	97.75	70.87	204.08

Notes: percentage of the k-step ahead forecast error variance explained by the shock on the corresponding column. Percentages do not sum to 100 due to non-zero covariance between the shocks.

Table 7. Generalized variance decomposition of the forecast error of Mexico's NFA

Period	External shocks					Domestic shocks				All external shocks	All domestic shocks
	USA	EA	JAP	ARG	BRA	MEXy	MEXsr	MEXrer	MEXnfa		
1	0.75	0.14	0.17	21.32	16.56	10.24	1.21	28.83	90.97	38.93	131.26
4	0.91	0.14	0.21	28.70	19.86	8.22	3.57	23.17	82.20	49.82	117.17
8	1.00	0.17	0.33	29.31	20.99	6.24	4.47	19.63	76.54	51.81	106.88
12	1.05	0.22	0.44	28.57	21.35	4.97	4.62	17.74	73.58	51.64	100.91
20	1.10	0.31	0.58	27.14	21.57	3.55	4.53	15.69	70.33	50.71	94.10
40	1.15	0.42	0.75	25.44	21.67	2.22	4.29	13.70	67.15	49.42	87.36

Notes: percentage of the k-step ahead forecast error variance explained by the shock on the corresponding column. Percentages do not sum to 100 due to non-zero covariance between the shocks.

Each entry shows the proportion of the NFA forecast error variance explained by conditioning on contemporaneous and expected future values of selected factors at different quarterly horizons. Note that due to the positive correlations existing among shocks the variance proportions do not sum to 100.

The striking result common to all three countries is that domestic shocks explain by far the largest proportion of the NFA forecast error variance. The influence of external shocks on Argentinean NFA increases over time, though not monotonically, reaching 3.41% after 10 years, while internal shocks effect, though decreasing steadily, account for almost all of NFA variability. Among external shocks, Brazilian variables are increasingly predominant, while US ones prevail on the other developed countries at most horizons.

For Brazil's NFA, domestic shocks explain an even larger proportion of forecast error variance (almost 98% at all horizons). Among external shocks, Argentinean and Mexican variables, respectively, dominate, while US and Japan's factors follow closely.

As for Mexican NFA, the influence of domestic factors decreases over time while, conversely, that of foreign factors increases. A large proportion of variability is accounted for by Argentina and Brazil, respectively, while, as expected, US variables come next.

It is worth noting that in all Latin American countries the real exchange rate exerts by far the largest influence on the NFA variability, apart from NFA own shock. The reason may be twofold. Firstly, the exchange rate-based stabilization programs undertaken by many Latin American countries over the last two decades caused a real appreciation following the adoption of a fixed nominal exchange rate, mainly due to inflation inertia (see Edwards (2001)). When hit by negative macroeconomic shocks, Latin American countries, already suffering from real exchange rate overvaluation, proved unwilling to bear the political consequences of the necessary macroeconomic tightening. As a result the stabilization programs lost their credibility to international investors, which in turn disinvested their assets. Secondly, and strictly related to the argument just developed, one of the main factor affecting the international investors' portfolio choice in emerging markets is the exchange rate risk component of interest rates: volatile exchange rates, as reflected in the real exchange rate with inflation inertia, will result in volatile capital flows. This is consistent with the evidence provided by the recurrent currency and financial crises characterizing the recent economic history in Latin America.

The GFEVD, supporting the evidence provided by the GIRF analysis, shows that domestic factors account for almost all the variability of net foreign capital in Latin America. It is important to note that, quite surprisingly, Japanese variables are more important than European ones for Latin American NFA at all horizons, despite trade linkages privilege Europe. This is better seen by looking at Table 8 that reports the effect of external and internal shocks on the NFA of Latin America as a region.

Table 8. Generalized variance decomposition of the forecast error of regional NFA

Period	**External shocks**			**Domestic shocks**				**All external shocks**	**All domestic shocks**
	USA	EA	JAP	y	sr	rer	nfa		
1	2.17	0.15	1.33	19.57	4.26	45.82	95.45	3.65	165.11
4	2.23	0.17	1.46	19.73	5.05	43.86	90.93	3.86	159.57
8	2.28	0.20	1.58	19.61	5.41	43.28	86.82	4.06	155.12
12	2.31	0.22	1.67	19.53	5.51	43.13	84.29	4.20	152.46
20	2.34	0.25	1.78	19.47	5.54	43.07	81.29	4.38	149.37
40	2.38	0.29	1.91	19.49	5.53	43.09	78.07	4.58	146.17

Notes: percentage of the k-step ahead forecast error variance explained by the shock on the corresponding column. Percentages do not sum to 100 due to non-zero covariance between the shocks.

Among industrial countries, US exerts the largest influence on the regional NFA variability, while Japan comes next with an explained proportion much larger than the European one. The most important domestic factors are, apart from the NFA themselves, the real exchange rate and real output, respectively, while interest rates account for a modest proportion at all horizons. The decreasing importance of domestic compared to foreign shocks over time is consistent with the larger influence of external factors highlighted by the long-run analysis of Subsection 4.1.

4.2.3 Contagion

The literature on the international transmission of financial crises in emerging markets focuses on three kinds of channels: common shocks (mainly from industrial countries), investors' behaviour, trade linkages. The GVAR approach allows some reflections on the subject. The pair-wise correlations of NFA equations residuals of Latin American country-specific models are 0.22 between Argentina and Brazil, 0.08 between Argentina and Mexico, and 0.20 between Brazil and Mexico, only the first one being statistically significant at the conventional level. Thus, after controlling for other factors (either internal and external), the residual co-movement of NFA in Latin America appears negligible. This supports the hypothesis of common shocks and international investors' behaviour as main driving forces of contagion.

Table 9 reports the proportion of the forecast error variance of Latin American countries' NFA due to shocks to the neighbours' NFA.

Table 9. Generalized variance decomposition of the forecast error of Latin American NFA

Period	**Argentina's NFA**			**Brazil's NFA**			**Mexico's NFA**		
	ARG	BRA	MEX	ARG	BRA	MEX	ARG	BRA	MEX
1	97.55	0.66	0.08	57.54	97.59	2.63	5.56	5.82	90.97
4	87.56	0.51	0.07	53.99	97.84	2.85	7.89	6.57	82.20
8	75.00	0.44	0.06	52.80	97.87	2.93	6.92	6.72	76.54
12	66.40	0.40	0.05	52.06	97.85	2.96	5.81	6.73	73.58
20	55.85	0.36	0.04	51.10	97.81	2.99	4.36	6.68	70.33
40	44.01	0.31	0.03	50.02	97.75	3.01	2.88	6.59	67.15

Notes: percentage of the k-step ahead forecast error variance of the Latin American NFA (first row, in bold) explained by the shock on the other regional NFA. Percentages do not sum to 100 due to non-zero covariance between the shocks.

The contribution of Brazilian and Mexican typical financial markets shocks to Argentinean NFA is negligible. This is not surprising since it is well known that the good performance of Argentina's currency board over most of the nineties prevented the Mexican and Brazilian crises of 1994 and 1998, respectively, from being transmitted to Argentina. On the other hand, shocks to Argentina's foreign capital positions explain a large part of the Brazilian NFA forecast error variance, while shocks originated in Mexico are much less important.[16] A noticeable contribution is given by Brazilian shocks to the Mexican NFA variability.

4.2.4 Finance-based exchange rate regime choice

The debate on the optimal exchange rate regime choice for emerging markets is concerned with the effects of financial shocks and financial integration on the adoption of common currencies or pegged exchange rates. At present there is no clear theory of how financial variables co-movements relate to the choice of an exchange rate regime and to optimum currency areas. Nevertheless, if there were evidence that Latin American countries are routinely hit by large, common financial shocks (such as sudden loss of appetite for Latin American financial assets, regardless of fundamentals), one might think that monetary policy should react similarly across countries. In that case, the adoption of a common currency (whether a regional currency or the US dollar) could be a sensible option. However, as noted above, the correlation of NFA residuals of country-specific models is very low and statistically insignificant for 2 out of 3 Latin American countries, which thus do not appear to be hit by the same financial shocks. On the other hand, as shown in Table 4, the responses of NFA to some external shocks are highly correlated across countries, implying that Latin American NFA react symmetrically. This supports the hypothesis that common international investors are important determinants of the pattern

[16] Boschi (2005) argues that though the cross-country correlation of relevant financial prices increased in the aftermath of the Argentina's crisis of 2001-2002, evidence is against the contagion hypothesis once the analysis controls for heteroscedasticity.

of NFA in the region; at the same time it suggests that a common currency would not be much disturbed by investors' behaviour, unless idiosyncratic shocks become more relevant in the future.

However, the evidence provided by the GFEVD suggests that a typical shock to the real exchange rate accounts for a large part of the NFA variability. If the interpretation proposed above, based on the negative effects of exchange rate-based stabilization programs and of recurrent currency crises on investors' expectations, is correct, the adoption of a flexible nominal exchange rate regime would be more consistent with stable expectations by foreign investors. Granger causality tests, both the 1-step and the multi-step version (see Lütkepohl and Burda (1997)) point overall to the same direction: the real exchange rate is Granger causal for the NFA in Argentina and Mexico, while the direction of causality is unclear in Brazil.[17]

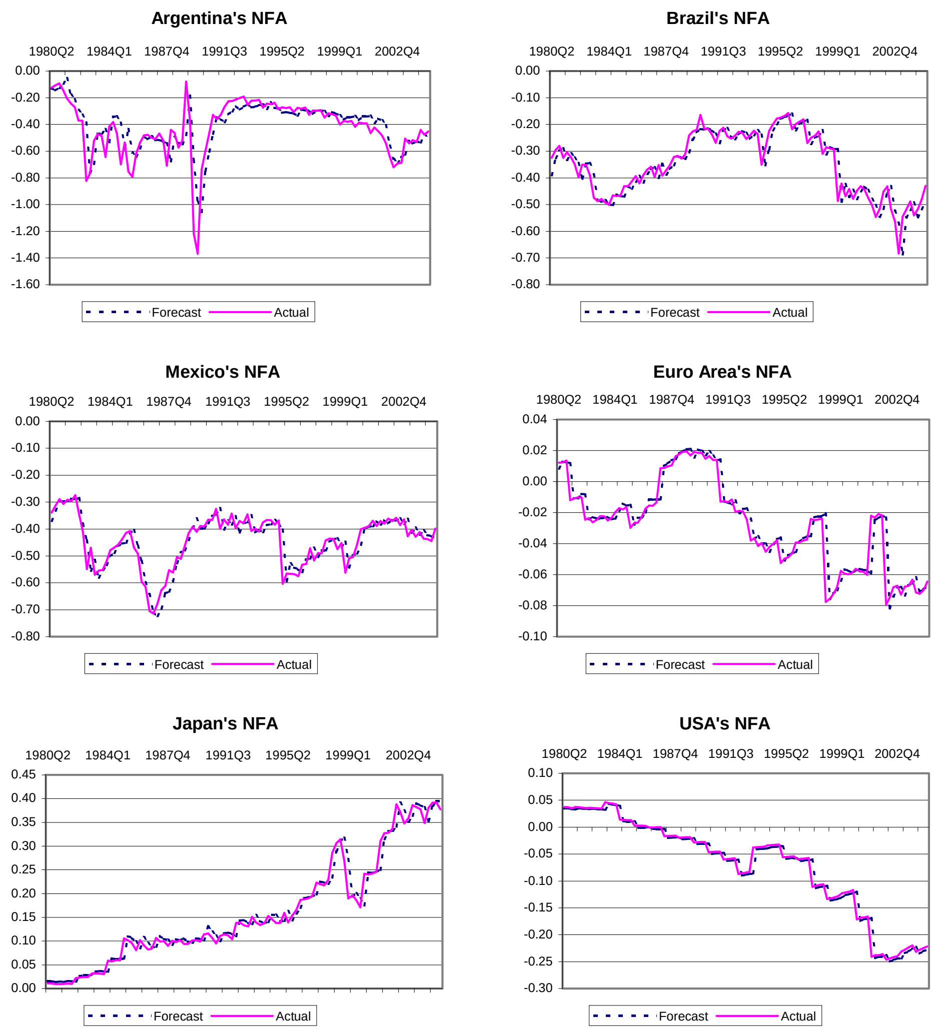
Towards the same conclusion leads the large and significant long-run relationship between the real exchange rate and NFA suggesting a preference for nominal exchange rate flexibility in order to allow the real adjustment to take place as smoothly as possible.

4.3 Forecasting performance

Figure 15 shows the forecasts, both in-sample and out-of-sample, and the actual values of NFA series. The out-of-sample forecast is implemented over the quarters 2004:1-2004:4. The GVAR estimates track remarkably well the actual data series.

[17] Granger causality tests results are available on request.

Fig. 15. Forecast and actual values of NFA



In order to evaluate quantitatively the out-of-sample forecasting performance of the GVAR, the Root Mean Squared Error (RMSE) for the NFA series obtained from the GVAR are compared to those obtained from a random walk with drift (RWWD) model. Both figures are reported in Table 10.

Table 10. RMSE of the GVAR and the RWWD

	usanfa	argnfa	branfa	mexnfa	eanfa	japnfa	Average
GVAR	0.0069	0.0555	0.0438	0.0258	0.0052	0.0172	0.0257
RWWD	0.0067	0.0527	0.0426	0.0263	0.0052	0.0176	0.0252

Notes: RMSE of the out-of-sample forecast over the period 2004:1 - 2004:4.

The RMSE from the GVAR is lower than that from the RWWD for the NFA of Mexico, Euro Area and Japan. Further, the average RMSE is computed across all variables and periods for both in-sample and out-of-sample forecasts. The former is very satisfactory and supports the model specification since it produces an average RMSE of 0.08 compared to 0.11 given by the RWWD model. The out-of-sample RMSE is as small as 0.0068, but the random walk does better with a RMSE of 0.0056.

4.4 Unrestricted GVAR

The analysis above is based on the restricted GVAR model. The results of the LR test on the over-identifying restrictions of the cointegration space, however, provide only weak support for these restrictions, even after controlling for small sample properties of the estimation through bootstrap techniques. Nevertheless, given the strong theoretical prior in favour of the long-run relations linking the NFA and the real exchange rate to the other variables, the main analysis is conducted with the restricted model. However, in order to check for the possibility that the main results discussed in the previous Sections be driven by the long-run structure imposed on the model, I estimate an unrestricted version of the GVAR and carry out the same short-run analysis based on GIRF and GFEVD.[18]

The GIRF for Argentina's NFA in the unrestricted model are somewhat different from the restricted one. The response to the domestic output shock is smaller and has opposite sign, while that to the Euro Area's shock doubles. The responses to the European short- and long-term interest rates change sign, while those to all other interest rates remain unaffected. The GIRF for Brazil are largely the same, while those for Mexico change only slightly, apart from the response to domestic output which changes its sign in the unrestricted model. As a result, the GIRF for Latin America as a region remain largely the same except for the responses to Argentinean and Mexican output shocks. Remarkably, the responses of each country's NFA to the neighboroughs' NFA shocks are exactly the same.

The GFEVD results are virtually the same in the unrestricted as in the restricted model, the only slight difference being the reduced importance of the Japanese variables in explaining Argentina's NFA in the unrestricted GVAR. It is important to note, however, that the results on the relative importance

[18] Figures and tables with the unrestricted estimation results, though unreported to save space, are available from the author on request.

of foreign and domestic factors are very robust; similarly, the large impact of Japanese macroeconomic factors on Latin American NFA is strongly confirmed. Finally, the results on contagion, with causality mainly running from Argentina to Brazil, and the Mexican NFA only modestly affected by the other countries' shocks, hold largely unchanged in the unrestricted model.

Therefore, the main results regarding the relative importance of domestic and foreign factors, as well as specific industrial countries' influence, remain unchanged when the unrestricted model is considered. Similarly, the forecasting performance of the unrestricted GVAR compared to the RWWD is the same.

5 Conclusion

In this paper I use a mixed macro-econometric modelling strategy based on GLPS (2000, 2003) and PSW (2004a) to understand the main determinants of foreign capital in Argentina, Brazil, and Mexico. The global economy includes also three industrial countries/regions, namely USA, Euro Area, and Japan. I first estimate six VECM models, one for each country/region, and then I embed the estimated parameters in a Global VAR. Two long-run restrictions derived from economic theory are tested and imposed in order to identify the cointegration space of all VECM models except the US one. The theoretical framework is based on the portfolio balance approach to the NFA position and the "transfer effect" theory of the real exchange rate.

The long-run analysis shows that foreign factors are relatively more important than domestic ones in the long-run equilibrium behaviour of capital flows. NFA are associated to a larger extent to foreign output rather than to domestic output in Argentina and Brazil, while the coefficients of domestic and foreign output are statistically insignificant in the Mexican model. The coefficient of domestic interest rate is larger for Brazil, while it is smaller for Argentina and Mexico, when compared to foreign rates. However, on the whole the short-term interest rates have the largest coefficient among all factors affecting the NFA position.

The short-run analysis conducted through the Generalized Impulse Response Functions and Forecast Error Variance Decomposition supports strongly the hypothesis that domestic shocks are predominantly responsible for the short-run dynamics of Latin American NFA. Moreover, Japanese macroeconomic factors exert an important role in determining the short-run variability of capital flows to Latin America. While most of the literature on capital flows to emerging markets concentrate on the role played by the US, this paper shows that it is crucial to distinguish between different sources of external disturbances.

Once controlled for common macroeconomic variables, the cross-country correlation of idiosyncratic shocks to NFA is negligible. This supports the view that contagion in these countries depends on common external factors or investors' behaviour rather than the transmission of financial shocks.

An important implication of these results for the choice of the exchange rate regime is that, given the large proportion of NFA forecast error variance explained by the real exchange rate, the adoption of a flexible nominal exchange rate would allow for a smaller variability of external financing. To the same conclusion points the long-run relation between the real exchange rate and the NFA positions.

6 Appendix

6.1 Integration properties of the series

Tables A1a and A1b report the ADF test statistics for the levels and first differences of all variables. Specifically, Table A1a presents the ADF statistics with lag order selected according to the AIC, while the modified AIC proposed by Ng and Perron (2001), which takes into account the size distortion of ordinary AIC, is used in the ADF test of Table A1b. Together, the two tests cannot reject the hypothesis that most variables are integrated of order 1. Argentina and Brazil's domestic and foreign short-run interest rates appear to be I(0) with both tests, along with Euro Area's foreign short-run interest rates.

Table A1a. ADF unit root test statistics (based on AIC order selection)

	Argentina	**Brazil**	**Mexico**	**Euro Area**	**Japan**	**USA**
y	-4.39	-2.34	-2.58	-2.75	-1.43	-2.71
Δy	-4.46	-2.61	-3.48	-3.18	-3.35	-3.86
sr	-4.32	-7.38	-2.74	-1.96	-1.70	-1.25
Δr	-6.62	-6.61	-5.85	-9.33	-12.78	-6.80
lr	-	-	-	-2.77	-1.74	-1.77
Δlr	-	-	-	-5.80	-13.18	-5.66
q	-2.43	-2.59	-3.65	-2.89	-2.12	-2.56
Δq	-9.91	-8.42	-4.27	-6.89	-7.71	-3.11
nfa	-3.13	-1.00	-3.45	-2.69	-1.61	-2.43
Δnfa	-6.22	-7.67	-4.41	-10.12	-4.18	-3.80
y*	-2.50	-2.78	-3.13	-3.13	-3.91	-1.95
Δy*	-2.82	-4.78	-4.08	-3.80	-3.75	-3.59
sr*	-7.29	-3.86	-2.29	-11.25	-2.52	-3.69
Δsr*	-6.58	-15.43	-12.37	-8.66	-9.81	-5.93
lr*	-1.21	-1.13	-1.77	-1.36	-2.77	-1.95
Δlr*	-5.81	-5.91	-5.54	-6.11	-5.28	-12.48
tt	-3.42	-2.06	-2.56	-1.66	-2.25	-4.00
Δtt	-4.65	-15.40	-3.00	-9.20	-7.35	-5.59
oil	-2.72					
Δoil	-5.61					

Notes: The ADF statistics are based on univariate AR(p) models in the levels with p chosen according to the AIC, with a maximum lag order of 10. The sample period is 1980:1-2003:4. The regressions for all the level variables include an intercept and a linear trend with the exception of interest rates whose underlying regressions include only an intercept. The 95% critical value for regressions with trend is -3.46 and for regressions without trend -2.90.

Table A1b. ADF unit root test statistics (based on modified AIC order selection)

	Argentina	Brazil	Mexico	Euro Area	Japan	USA
y	-1.88	-1.66	-2.58	-2.22	-0.46	-2.30
Δy	-3.39	-12.57	-3.73	-2.93	-3.35	-5.19
sr	-4.04	-10.74	-2.70	-0.88	-1.70	-1.22
Δr	-18.58	-16.49	-12.95	-15.10	-28.63	-12.72
lr	-	-	-	-1.67	-1.50	-1.77
Δlr	-	-	-	-15.34	-29.06	-12.38
q	-2.43	-2.22	-2.29	-2.52	-1.56	-1.51
Δq	-6.53	-8.42	-4.27	-4.28	-4.05	-3.11
nfa	-2.68	-0.64	-2.28	-2.69	-2.71	-2.21
Δnfa	-8.85	-2.39	-5.11	-10.12	-8.94	-3.80
y*	-2.03	-2.78	-3.13	-2.23	-3.03	-1.95
Δy*	-13.71	-2.98	-5.05	-3.64	-5.09	-5.66
sr*	-10.73	-3.86	-1.29	-3.77	-0.74	-2.71
Δsr*	-16.46	-18.57	-15.19	-17.98	-15.98	-15.67
lr*	-1.21	-1.13	-1.17	-1.38	-1.38	-1.09
Δlr*	-13.78	-13.82	-12.38	-15.80	-12.61	-22.35
tt	-2.44	-2.06	-2.56	-1.66	-1.59	-2.96
Δtt	-13.19	-4.76	-14.11	-4.25	-4.81	-9.17
oil	-1.75					
Δoil	-8.43					

Notes: The ADF statistics are based on univariate AR(p) models in the levels with p chosen according to the modified AIC, with a maximum lag order of 10. The sample period is 1980:1-2003:4. The regressions for all the level variables include an intercept and a linear trend with the exception of interest rates whose underlying regressions include only an intercept. The 95% critical value for regressions with trend is -3.46 and for regressions without trend -2.90.

The sample period considered in this study includes many episodes of financial and currency crises and, more generally, economic distress both in the industrial and less developed countries. Due to these events, the time series can present structural breaks and in turn this can distort the unit root tests. In fact, instead of non stationary processes with unit roots, the series could be better characterized as stationary processes fluctuating around a deterministic trend with structural breaks. In order to take into account this possibility, I perform the ADF unit root tests with breaks proposed by Saikkonen and Lütkepohl (2002) and Lanne *et al.* (2002, 2003).[19] The test is based on a procedure that estimates first the deterministic component of the series, then subtracts it from the original series and finally runs the ADF unit root test on the modified series. The distribution under the null hypothesis is non-standard, thus I use the critical values provided by Lanne *et al.* (2002). The test (see Tables A2a and A2b) suggests that Argentina and Brazil's domestic and foreign interest rates and Euro Area's short-term foreign interest rates are I(1).

[19] The ADF unit root tests with breaks are performed using the econometric package JMulTi.

Table A2a. ADF unit root tests with breaks statistics

	y	sr	lr	q	nfa
Argentina					
Suggested break date	1994 Q2	1990 Q3		1989 Q2	1989 Q3
Test statistic	-5.25 [4]	-1.34 [2]		-0.95 [0]	-1.24 [2]
Critical value at 5%	-3.03	-2.88		-3.03	-3.03
Brazil					
Suggested break date	1990 Q2	1990 Q1		1989 Q2	1999 Q1
Test statistic	-2.80 [0]	-1.32 [2]		-3.28 [1]	-2.06 [0]
Critical value at 5%	-3.03	-2.88		-3.03 (-3.55 at 1%)	-3.03
Mexico					
Suggested break date	1995 Q2	1988 Q1		1995 Q1	1995 Q1
Test statistic	-2.06 [4]	-1.90 [12]		-1.85 [3]	-2.83 [3]
Critical value at 5%	-3.03	-2.88		-3.03	-3.03
Euro Area					
Suggested break date	1993 Q1	1981 Q4	1981 Q4	1991 Q2	2002 Q1
Test statistic	-1.71 [4]	-1.22 [3]	-1.83 [3]	-2.15 [1]	-2.01 [0]
Critical value at 5%	-3.03	-2.88	-2.88	-3.03	-3.03
Japan					
Suggested break date	1989 Q2	1986 Q4	1987 Q1	1998 Q4	1999 Q1
Test statistic	-1.79 [3]	-2.26 [3]	-2.91 [3]	-1.51 [1]	-1.35 [1]
Critical value at 5%	-3.03	-2.88	-2.88 (-3.48 at 1%)	-3.03	-3.03
USA					
Suggested break date	1981 Q2	1981 Q3	1981 Q3	1982 Q3	2001 Q1
Test statistic	-2.55 [2]	-2.29 [2]	-2.22 [2] not AIC	-1.29 [0]	-2.38 [0]
Critical value at 5%	-3.03	-2.88	-2.88	-3.03	-3.03
	y*	**sr***	**lr***	**tt**	**oil**
Argentina					
Suggested break date	1990 Q2	1990 Q1	1981 Q4	1989 Q2	-
Test statistic	-2.36 [0]	-1.29 [2]	-1.69 [3]	-2.99 [8]	-
Critical value at 5%	-3.03	-2.88	-2.88	-3.03	-
Brazil					
Suggested break date	1994 Q2	1990 Q3	1986 Q4	1989 Q2	-
Test statistic	-2.40 [6]	-1.44 [2]	-3.72 [5]	-1.75 [1]	-
Critical value at 5%	-3.03	-2.88	-2.88	-3.03	-
Mexico					
Suggested break date	1981 Q2	1990 Q2	1981 Q3	1986 Q2	-
Test statistic	-2.43 [2]	-2.74 [2]	-2.11 [2]	-2.17 [4]	-
Critical value at 5%	-3.03	-2.88	-2.88	-3.03	-
Euro Area					
Suggested break date	1990 Q2	1990 Q1	1986 Q2	1988 Q3	-
Test statistic	-1.82 [2]	-0.52 [0]	-6.14 [4]	-1.05 [0]	-
Critical value at 5%	-3.03	-2.88	-2.88	-3.03	-
Japan					
Suggested break date	1990 Q1	1986 Q3	1981 Q3	1986 Q2	-
Test statistic	-1.97 [2]	-3.12 [3]	-1.94 [2]	-2.95 [1]	-
Critical value at 5%	-3.03	-2.88 (-3.48 at 1%)	-2.88	-3.03	-
USA					
Suggested break date	1995 Q2	1990 Q1	1981 Q4	1981 Q3	1990 Q3
Test statistic	-1.49 [0]	-1.17 [0]	-1.56 [3]	-3.55 [1]	-2.09 [4]
Critical value at 5%	-3.03	-2.88	-2.88	-3.03 (-3.55 at 1%)	-3.03

Notes: the regressions include an intercept and a linear trend for all variables with the exception of interest rates whose underlying regression include only an intercept. The lag order, selected according to the AIC with a maximum lag order of 10, is reported in square brackets.

Table A2b. ADF unit root tests with breaks statistics

	Δ**y**	Δ**sr**	Δ**lr**	Δ**q**	Δ**nfa**
Argentina					
Suggested break date	1994 Q2	1990 Q3	-	1989 Q2	1989 Q3
Test statistic	-5.49 [10]	-3.15 [9]	-	-7.38 [0]	-3.46 [4]
Critical value at 5%	-2.88	-2.88	-	-2.88	-2.88
Brazil					
Suggested break date	1990 Q2	1994 Q1	-	1989 Q2	1999 Q1
Test statistic	-3.02 [7]	-3.82 [8]	-	-8.26 [0]	-7.76 [2]
Critical value at 5%	-2.88	-2.88	-	-2.88	-2.88
Mexico					
Suggested break date	1995 Q2	1988 Q1	-	1995 Q1	1995 Q1
Test statistic	-4.43 [6]	-3.08 [10]	-	-4.88 [7]	-4.62 [10]
Critical value at 5%	-2.88	-2.88	-	-2.88	-2.88
Euro Area					
Suggested break date	1993 Q1	1998 Q1	1998 Q1	1992 Q3	2002 Q1
Test statistic	-3.08 [3]	-8.48 [2]	-9.01 [2]	-6.35 [0]	-10.02 [0]
Critical value at 5%	-2.88	-2.88	-2.88	-2.88	-2.88
Japan					
Suggested break date	1989 Q2	1989 Q1	1997 Q3	1995 Q2	1999 Q1
Test statistic	-2.98 [2]	-4.68 [10]	-5.98 [4]	-7.16 [0]	-2.19 [10]
Critical value at 5%	-2.88	-2.88	-2.88	-2.88	-2.88
USA					
Suggested break date	1991 Q1	2001 Q4	1986 Q2	1995 Q2	2001 Q1
Test statistic	-3.69 [9]	-4.81 [10]	-6.89 [5]	-3.11 [3]	-10.31 [0]
Critical value at 5%	-2.88	-2.88	-2.88	-2.88	-2.88

	Δ**y***	Δ**sr***	Δ**lr***	Δ**tt**	Δ**oil**
Argentina					
Suggested break date	1990 Q2	1990 Q1	1986 Q2	1989 Q2	-
Test statistic	-3.17 [7]	-4.97 [6]	-5.57 [6]	-6.22 [6]	-
Critical value at 5%	-2.88	-2.88	-2.88	-2.88	-
Brazil					
Suggested break date	1994 Q2	1990 Q3	1986 Q2	1997 Q2	-
Test statistic	-4.47 [5]	-3.17 [9]	-5.63 [6]	-13.30 [0]	-
Critical value at 5%	-2.88	-2.88	-2.88	-2.88	-
Mexico					
Suggested break date	1982 Q1	1990 Q2	1986 Q2	1985 Q4	-
Test statistic	-4.21 [2]	-6.45 [2]	-6.73 [5]	-3.08 [8]	-
Critical value at 5%	-2.88	-2.88	-2.88	-2.88	-
Euro Area					
Suggested break date	1990 Q2	1990 Q3	1986 Q2	1988 Q3	-
Test statistic	-4.30 [8]	-3.72 [8]	-6.40 [6]	-2.43 [9]	-
Critical value at 5%	-2.88	-2.88	-2.88	-2.88	-
Japan					
Suggested break date	1990 Q1	1990 Q2	1986 Q2	1986 Q2	-
Test statistic	-4.30 [5]	-6.64 [2]	-6.02 [6]	-7.96 [0]	-
Critical value at 5%	-2.88	-2.88	-2.88	-2.88	-
USA					
Suggested break date	2001 Q2	1988 Q1	1997 Q3	1990 Q4	1986 Q1
Test statistic	-3.13 [7]	-5.37 [6]	-11.27 [2]	-7.49 [1]	-5.84 [4]
Critical value at 5%	-2.88	-2.88	-2.88	-2.88	-2.88

Notes: the regressions do not include an intercept and a linear trend. The lag order, selected according to the AIC with a maximum lag order of 10, is reported in square brackets.

6.2 Order selection, specification, and parameters stability tests

I choose p_i, the lag order of the domestic variables, by comparing the results of various selection criteria, namely the Akaike information criterion (AIC), the Schwarz Bayesian criterion (SBC) and the log-likelihood ratio statistic (LR) adjusted to take into account small sample problems, starting from a maximum lag order of 4. The results are reported in Table A3.

Table A3. Test statistics for selecting the lag order of the endogenous (domestic) variables in the VARX$^*(p_i,q_i)$ model.

		Argentina		
Order (p_i)	AIC	SBC		Adjusted LR test
4	**396.0**	254.8		
3	386.7	265.7	$\chi 2(16)$ =	35.1713[.004]
2	385.7	284.8	$\chi 2(32)$ =	58.8953[.003]
1	381.3	**300.6**	$\chi 2(48)$ =	87.3067[.000]
0	148.3	87.8	$\chi 2(64)$ =	433.6722[.000]
		Brazil		
Order (p_i)	AIC	SBC		Adjusted LR test
4	529.8	388.6		
3	529.0	407.9	$\chi 2(16)$ =	23.4035[.103]
2	**537.1**	436.2	$\chi 2(32)$ =	34.3963[.354]
1	535.8	**455.1**	$\chi 2(48)$ =	**58.3991[.145]**
0	239.1	178.6	$\chi 2(64)$ =	493.5271[.000]
		Mexico		
Order (p_i)	AIC	SBC		Adjusted LR test
4	861.3	720.1		
3	862.7	741.6	$\chi 2(16)$ =	20.3793[.204]
2	**868.0**	767.2	$\chi 2(32)$ =	35.1906[.320]
1	867.9	**787.2**	$\chi 2(48)$ =	**57.6519[.160]**
0	635.1	574.6	$\chi 2(64)$ =	403.7881[.000]
		Euro Area		
Order (p_i)	AIC	SBC		Adjusted LR test
4	**1550.8**	1342.8		
3	1545.1	1368.6	$\chi 2(25)$ =	39.3925[.034]
2	1545.1	1400.1	$\chi 2(50)$ =	71.4694[.025]
1	1538.5	**1425.1**	$\chi 2(75)$ =	111.9507[.004]
0	1226.3	1144.3	$\chi 2(100)$ =	544.4913[.000]
		Japan		
Order (p_i)	AIC	SBC		Adjusted LR test
4	1364.3	1175.1		
3	1367.5	1209.9	$\chi 2(25)$ =	29.3600[.249]
2	**1374.4**	1248.3	$\chi 2(50)$ =	53.7989[.331]
1	1372.1	**1277.5**	$\chi 2(75)$ =	**90.5587[.106]**
0	1145.7	1082.6	$\chi 2(100)$ =	429.4398[.000]
		USA		
Order (p_i)	AIC	SBC		Adjusted LR test
4	1505.2	1335.0		
3	1512.0	1373.3	$\chi 2(25)$ =	25.6769[.425]
2	**1519.0**	1411.8	$\chi 2(50)$ =	51.1332[.429]
1	1513.1	**1437.4**	$\chi 2(75)$ =	**94.8525[.061]**
0	1129.4	1085.3	$\chi 2(100)$ =	672.2486[.000]

Notes: statistics in bold indicate the order selected by the relevant criterion/test. Unrestricted VARs are estimated with foreign variables treated as exogenous.

The SBC unambiguously selects the order 1 for all models; the AIC selects the order 4 for Argentina and Euro Area and the order 2 for Brazil, Mexico, Japan, and USA; finally, the LR selects an order higher than 4 for Argentina and Euro Area and 1 for Brazil, Mexico, Japan, and USA. According to the above results, and taking into account the limited sample size compared to the number of unknown parameters in each VARX* model, p_i is set equal to 1. This choice is comforted by the fact that the SBC estimates the lag order consistently, while the AIC does not (Lütkephol (1993), p. 383). In order to choose q_i, the lag order of the foreign ('star') specific variables, I run for each country/region an unrestricted VAR in which the foreign variables are treated as endogenous.[20] The SBC criterion selects invariantly a lag order of one. Basing on this evidence, considering data limitations, and following PSW, I set q_i equal to one in all models.

The specification tests are conducted on both the unrestricted and the restricted version of each country/region model. Univariate specification tests for the unrestricted models (Table A4) show that the null hypothesis of no serial correlation is rejected only for the output equation of the Argentina's model, while the null of normality is rejected for 3 equations: the Japanese long-term real interest rate, the US output and the US NFA.

[20] These results are unreported to save space.

Table A4. Unrestricted models - Univariate specification tests statistics.

	Δy	**Δstr**	**Δltr**	**Δq**	**Δnfa**
Argentina					
Serial Correlation F(4,78)	4.41 [0.003]**	1.03 [0.399]	-	0.68 [0.609]	1.74 [0.149]
Normality $\chi^2(2)$	1.48 [0.478]	1.14 [0.564]	-	3.26 [0.196]	6.75 [0.034]
Heteroscedasticity F(1,93)	5.54 [0.021]*	0.83 [0.366]	-	2.01 [0.159]	1.20 [0.277]
Brazil					
Serial Correlation F(4,79)	1.03 [0.397]	0.18 [0.950]	-	1.27 [0.291]	0.94 [0.445]
Normality $\chi^2(2)$	1.86 [0.395]	4.67 [0.097]	-	4.55 [0.103]	2.10 [0.349]
Heteroscedasticity F(1,93)	2.10 [0.151]	0.15 [0.700]	-	0.70 [0.406]	1.89 [0.173]
Mexico					
Serial Correlation F(4,78)	1.97 [0.107]	0.80 [0.526]	-	1.78 [0.142]	1.67 [0.165]
Normality $\chi^2(2)$	0.80 [0.670]	0.05 [0.976]	-	4.34 [0.114]	2.03 [0.362]
Heteroscedasticity F(1,93)	1.48 [0.226]	0.15 [0.698]	-	0.42 [0.519]	0.44 [0.508]
Euro Area					
Serial Correlation F(4,77)	0.63 [0.646]	0.49 [0.746]	0.74 [0.569]	2.59 [0.043]	0.48 [0.749]
Normality $\chi^2(2)$	0.41 [0.816]	3.61 [0.164]	0.18 [0.912]	1.44 [0.487]	4.10 [0.129]
Heteroscedasticity F(1,93)	2.37 [0.127]	1.40 [0.240]	0.88 [0.351]	1.69 [0.197]	0.45 [0.502]
Japan					
Serial Correlation F(4,80)	1.46 [0.222]	0.54 [0.707]	0.73 [0.573]	0.86 [0.493]	0.63 [0.641]
Normality $\chi^2(2)$	0.83 [0.660]	1.99 [0.369]	8.31 [0.016]*	1.34 [0.513]	3.00 [0.223]
Heteroscedasticity F(1,93)	0.11 [0.741]	2.13 [0.148]	0.56 [0.455]	0.42 [0.518]	0.77 [0.382]
USA					
Serial Correlation F(4,83)	0.55 [0.699]	0.21 [0.930]	0.37 [0.826]	1.58 [0.188]	2.17 [0.080]
Normality $\chi^2(2)$	7.88 [0.019]*	0.41 [0.816]	1.68 [0.432]	0.05 [0.978]	10.43 [0.005]**
Heteroscedasticity F(1,93)	2.47 [0.119]	0.73 [0.395]	0.33 [0.568]	0.12 [0.726]	0.29 [0.592]

*Notes: the figures in square brackets are probability values associated with test statistics. * and ** denote statistical significance at the 5% and the 1% respectively.*

The univariate F test rejects the null of homoscedasticity only for Argentina's output. Multivariate tests statistics for unrestricted models, reported in Table A5, show that there is no evidence of autocorrelation, while the null of normality is rejected for Mexico at the 5% level, and for Brazil and USA at the 1%.

Table A5. Unrestricted models - Multivariate specification tests statistics.

Argentina		
Vector Portmanteau (11):		226.61
Vector Normality test:	$\chi 2(8)$ =	13.58 [0.0935]
Vector hetero test:	*F(190,324) =*	0.75 [0.9845]
Brazil		
Vector Portmanteau(11):		174.45
Vector Normality test:	$\chi 2(8)$ =	28.30 [0.0004]**
Vector hetero test:	*F(200,318) =*	0.79 [0.9689]
Mexico		
Vector Portmanteau(11):		200.79
Vector Normality test:	$\chi 2(8)$ =	19.03 [0.0147]*
Vector hetero test:	*F(200,283) =*	0.59 [1.0000]
Euro Area		
Vector Portmanteau(11):		268.493
Vector Normality test:	$\chi 2(10)$ =	11.78 [0.3001]
Vector hetero test:	*F(330,333) =*	0.43 [1.0000]
Japan		
Vector Portmanteau(11):		216.489
Vector Normality test:	$\chi 2(10)$ =	15.71 [0.1082]
Vector hetero test:	*F(330,408) =*	0.55 [1.0000]
USA		
Vector Portmanteau(11):		286.319
Vector Normality test:	$\chi 2(10)$ =	25.86 [0.0039]**
Vector hetero test:	*F(240,519) =*	1.04 [0.3553]

*Notes: the figures in square brackets are probability values associated with test statistics. * and ** denote statistical significance at the 5% and the 1% respectively.*

Finally, there is no evidence of heteroscedasticity. Table A6 collects specification tests statistics for the models with restricted cointegration space.

Table A6. Restricted models - Multivariate specification tests statistics.

	Argentina	
Serial correlation	*LM(16)* =	30.17 [0.0171]*
Normality:	*W(8)* =	12.77 [0.1201]
	Brazil	
Serial correlation	*LM(16)* =	23.56 [0.0995]
Normality:	*W(8)* =	18.66 [0.0168]*
	Mexico	
Serial correlation	*LM(16)* =	22.91 [0.1160]
Normality:	*W(8)* =	116.50 [0.0000]**
	Euro Area	
Serial correlation	*LM(25)* =	40.46 [0.0262]*
Normality:	*W(10)* =	14.82 [0.1389]
	Japan	
Serial correlation	*LM(25)* =	47.43 [0.0044]**
Normality:	*W(10)* =	16.96 [0.0754]
	USA	
Serial correlation	*LM(25)* =	35.53 [0.0791]
Normality:	*W(10)* =	22.76 [0.0117]

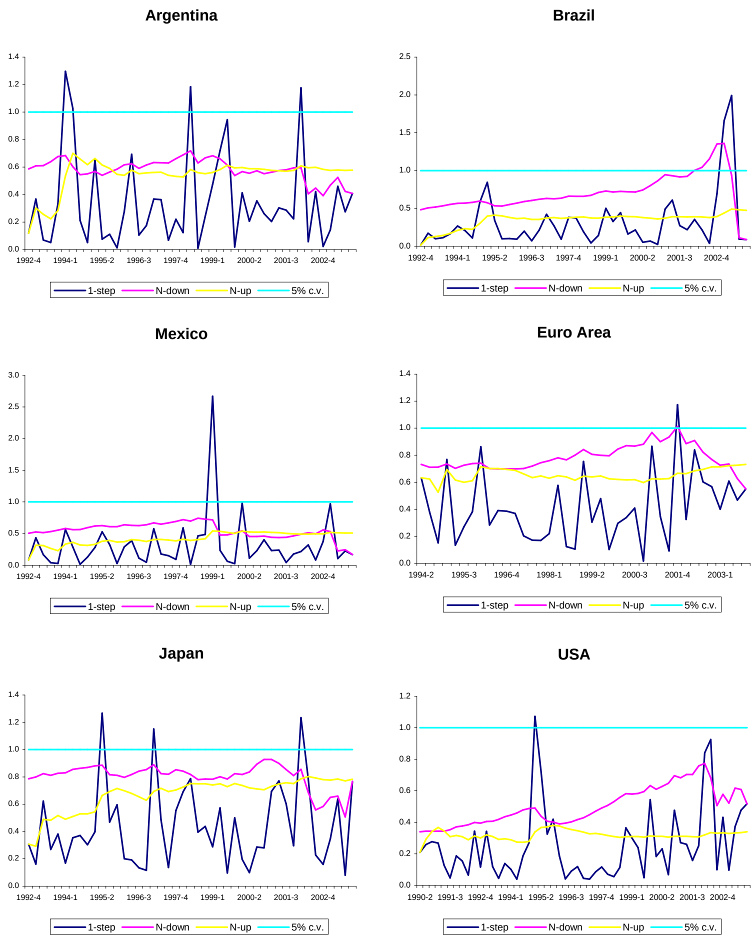
*Notes: the figures in square brackets are probability values associated with test statistics. * and ** denote statistical significance at the 5% and the 1% respectively. Serial correlation tests the null of serial uncorrelated residuals against the alternative that residuals follow a VAR(1).*

The null of serial correlation is rejected at the 5% level for Argentina and Euro Area, and at the 1% level for Japan. The null of normality is rejected only for Brazil (at 5%) and Mexico (at 1%).

In order to check parameters stability, I conduct a number of tests on either unrestricted and restricted versions of each model. Figure A1 displays multivariate *1-step*, *Break-point* (*N down-step*), and *Forecast* (*N up-step*) Chow tests results for the unrestricted VECM models. F-test statistics are normalized to the corresponding 5% level critical values (indicated by the horizontal line).[21]

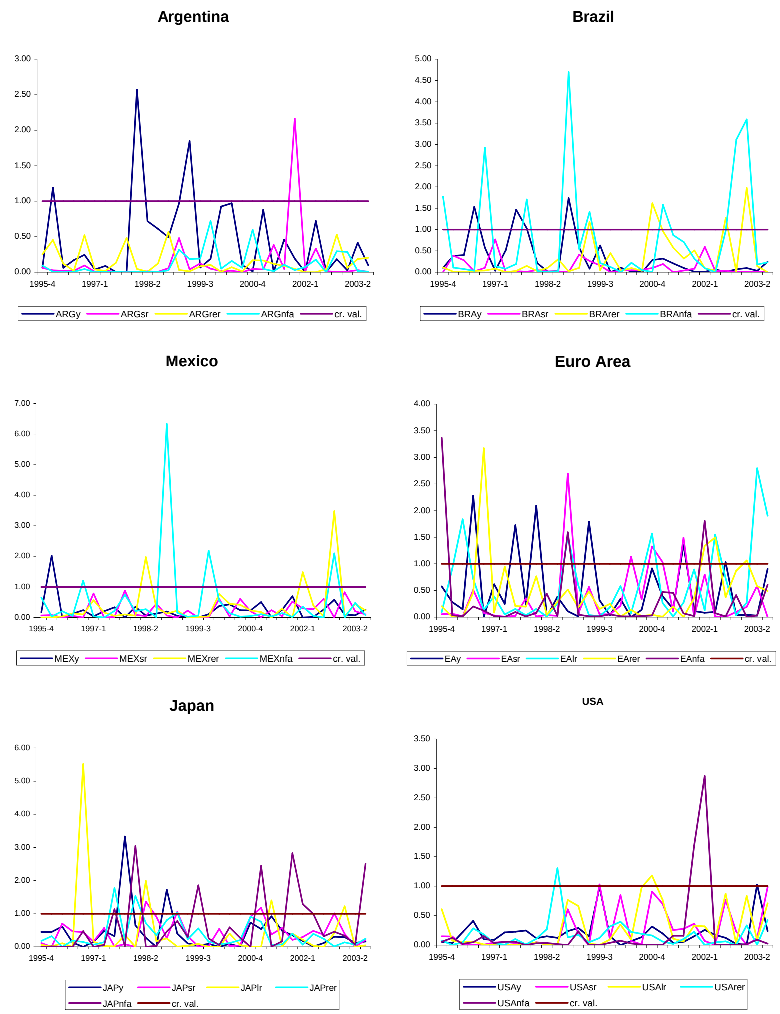
[21] See Doornik and Hendry (2001), ch. 15 for details.

Fig. A1. Chow stability tests on unrestricted models



The parameters appear stable, except for some quarters according to the *1-step* test in all models, and for the *Break-point* test for the Brazil's model. The Chow *Forecast* LR-test on the restricted models parameters (Figure A2), on the other hand, reveal some instability that is more pronounced for the output equations of the Argentina's and Euro Area's model, and the NFA of all models except Argentina.

Fig. A2. Chow 1-step ahead forecast LR-tests with 95% normalized critical values



Notes: test of the null hypothesis that the equation short-run coefficients are constant for date t relative to date t+1

Given the macroeconomic turbulence that has characterized all these countries over the sample span, such results are not surprising. However, it is comforting that the parameters are stable over most of the period and instability is limited to few, though relevant, peaks. Dealing with this problem would need to resort to different modelling econometric techniques capable of taking into account breaks in the underlying process, such as Markov-switching models, but this is out of the scope of this study.

Finally, Table A7 collects the LSTR1 F-statistics which test the null hypothesis of parameter constancy for the cointegration relations against a logistic STR model of order one for the same set of parameters (Teräsvirta (1998)).

Table A7. Restricted models - Cointegration relations parameters constancy tests.

	Argentina	Brazil	Mexico
	F (6,72)	*F (3,78)*	*F (6,73)*
y	1.25 [0.2910]	0.59 [0.6238]	1.06 [0.3970]
sr	0.31 [0.9299]	0.62 [0.6049]	0.70 [0.6469]
lr	-	-	-
q	0.14 [0.9898]	0.41 [0.7482]	1.40 [0.2265]
nfa	0.89 [0.5096]	0.39 [0.7584]	0.73 [0.6235]
	Euro Area	**Japan**	**USA**
	F (6,73)	*F (6,76)*	*F (6,79)*
y	1.34 [0.2499]	1.99 [0.0772]	1.37 [0.2357]
sr	0.38 [0.8865]	0.05 [0.9994]	0.90 [0.5003]
lr	0.19 [0.9773]	0.65 [0.6907]	0.51 [0.7987]
q	1.04 [0.4059]	0.60 [0.7330]	0.60 [0.7322]
nfa	0.72 [0.6338]	1.07 [0.3894]	0.79 [0.5775]

Notes: F-tests for parameter constancy against LSTR1 model for cointegration relations (3rd order Taylor expansion).

There is no evidence of instability for this set of parameters.

6.3 Eigenvalues of F and cross-correlation of idiosyncratic shocks

Among the conditions for the validity of the whole GVAR approach is that the eigenvalues of the matrix $F = \mathbf{G}^{-1}\mathbf{H}$ be either on or inside the unit circle (see Section 3.3). Indeed, none of the eigenvalues of F lies outside the unit circle (unreported), and the number of unitary roots is 10, i.e. less than the sum of cointegrating relationships existing in the country/region specific models (this result depends on the choice of the trade weight matrix $\mathbf{W}_i$ - see PSW).

Another condition is that the cross dependence of idiosyncratic shocks be weak. The basic idea is that conditioning the estimation of country/region-specific VECM models on foreign variables considered as proxies of "common" global factors will leave only a modest degree of correlation of the remaining shocks across countries/regions. This is also important if we were to interpret the disturbances in the GIRF analysis as "geographically structural": an external shock is truly external if its contemporaneous correlation with internal shocks is weak.

A simple way to verify these claims is by computing the contemporaneous correlation of residuals across different country-specific models for each equation. Table A8 reports such correlation coefficients, computed as averages of the correlation coefficients between the residuals of each equation (variable) with all other countries/regions equations residuals.

Table A8. Average cross-section correlations of residuals.

	Argentina	Brazil	Mexico	Euro Area	Japan	USA
y	0.00	0.00	0.00	-0.02	0.03	0.01
	[-0.04]	[-0.03]	[-0.01]	[-0.17]	[0.27]	[0.14]
sr	-0.03	0.00	0.00	0.01	-0.01	-0.02
	[-0.24]	[0.01]	[0.01]	[0.06]	[-0.09]	[-0.18]
lr	-	-	-	0.01	0.00	0.01
	-	-	-	[0.08]	[-0.01]	[0.07]
q	0.00	-0.03	-0.05	-0.02	-0.03	-0.01
	[0.03]	[-0.31]	[-0.52]	[-0.22]	[-0.30]	[-0.13]
nfa	0.01	0.02	0.02	0.01	-0.04	-0.04
	[0.10]	[0.20]	[0.22]	[0.05]	[-0.38]	[-0.36]

Notes: each entry is the average correlation of the residual of the equation on the corresponding row for the country/region on the corresponding column with all other countries/regions endogenous variables residuals. Two-tailed t-test statistics with 93 d.f.are in square brackets. The null hypothesis is no correlation. The 5% critical value is 1.99.

All coefficients are very small. A two-tailed t-test rejects the hypothesis that these coefficients are significantly different from zero at the conventional level. Thus, the model seems to be successful in capturing the effect of common factors driving domestic variables.

6.4 Testing weak exogeneity

The whole analysis in this paper relies on the crucial assumption that foreign variables, as well as terms of trade and oil price, are weakly exogenous in the country/region-specific VECM models. This means that foreign variables are unaffected by deviations from the long-run equilibrium. The weak exogeneity assumption can be tested in two ways. The first one considers all variables as endogenous and then runs a LR test that the relevant rows of the loading matrix $\boldsymbol{\alpha}_i$ in equation (16) be zero. This is the right procedure when economic theory is uninformative as to whether the relevant variables are weakly exogenous or not.

The second procedure is preferable when economic theory is informative about the variables long-run behaviour. In my case the small open economy assumption that justifies treating foreign variables as weakly exogenous is rather cogent. Thus, following PSW, I test the joint significance of the error correction term in auxiliary equations of the country/region-specific foreign variables, $\mathbf{x}_{it}^*$. Specifically, this alternative procedure, originally proposed by Johansen (1992), requires to carry out the following regression for each l^{th} element of country i vector of foreign variables, $\mathbf{x}_{it}^*$:

$$\Delta x_{il,t}^{*} = \mu_{il} + \sum_{j=1}^{2} \gamma_{ijl} ECM_{i,t-1}^{j} + \boldsymbol{\varphi}_{il}' \Delta \mathbf{v}_{i,t-1} + \zeta_{il,t}$$

where μ_{il} is a constant, $ECM_{i,t-1}^{j}$, $j = 1, 2$ are the estimated error correction terms corresponding to the r_i cointegrating relations found in the i^{th} model, $\boldsymbol{\varphi}_{il}$ is a vector of coefficients, $\mathbf{v}_{i,t-1}$ is defined by (13), and $\zeta_{il,t}$ is the residual. Then, an F test of the joint hypotheses that $\gamma_{ijl} = 0$, $j = 1, 2$ is carried out. Table A9 reports the results.

Table A9. F statistics for testing the weak exogeneity of the country-specific foreign variables, terms of trade, and oil prices.

	y*	sr*	lr*	tt	oil
Argentina					
F(2,82)	6.53 [0.0023]**	0.10 [0.9070]	2.07 [0.1328]	1.35 [0.2644]	3.53 [0.0339]*
Brazil					
F(1,83)	0.12 [0.7283]	1.66 [0.2013]	0.36 [0.5515]	1.55 [0.2171]	0.40 [0.5280]
Mexico					
F(2,82)	0.69 [0.5041]	2.89 [0.0615]	0.64 [0.5307]	5.49 [0.0058]**	1.88 [0.1588]
Euro Area					
F(2,81)	0.70 [0.5011]	1.34 [0.2669]	0.38 [0.6849]	4.58 [0.0131]*	2.81 [0.0660]
Japan					
F(2,79)	1.73 [0.1835]	0.67 [0.5165]	0.87 [0.4208]	3.25 [0.0437]*	1.07 [0.3480]
USA					
F(2,81)	-	-	-	3.02 [0.0540]	3.38 [0.0388]*

*Notes: the figures in square brackets are probability values associated with test statistics. * and ** denote statistical significance at the 5% and the 1% respectively.*

The weak exogeneity assumption is rejected in the model of Argentina for output and the oil price, in the models of Mexico, Euro Area and Japan for the terms of trade, and in the model of US for the oil price. Most of the test statistics are significant at the 5% level of significance, but not at the 1%. Given the overall statistical support and the strong theoretical prior in favour of the weak exogeneity hypothesis, I estimate the country/specific models with foreign variables, terms of trade and oil price treated as weakly exogenous. The most questionable hypothesis is that concerning oil price in the US model. In consideration of the many geopolitical factors affecting the oil price over the sample period, I interpret the result for the US model as mild evidence in support of the null hypothesis of weak exogeneity.

6.5 Data sources and variables construction

Net Foreign Assets (F) The NFA series is a quarterly interpolation of the "Net external position" annual series provided by Lane and Milesi-Ferretti

(2006).

Population (POP) The source is the IFS database. The code is 99Z..ZF.... Available annual data are interpolated linearly.

Nominal Output (Y^{NC}) The series is the volume of GDP in billions of national currency. It is taken from IFS for all countries except for Brazil. The code is 99B./CZF.... The series for Brazil is obtained from IPEADATA.

Output (Y) The source for all countries, except Brazil, is the IFS database. The code is ..99BVP/RZF.. (2000=100). The quarterly data for Argentina's GDP volume index are only available from 1993:1; the series is extended backward using the rates of growth of the GDP index series provided by Oxford Economic Forecasting. The GDP index of Brazil is obtained by deflating (with the CPI) the GDP volume in billions of national currency provided by IPEADATA.

Price index (P) The source is the IFS' Consumer Prices Index (CPI), which code is 64...ZF... (2000=100).

Exchange rates (E) The source is the IFS' series of National Currency per US Dollar, with code ..RF.ZF... except fo Mexico for which the series ..WF.ZF... is used.

Nominal short-term interest rates (R) The series is the Money Market Rate or equivalent (code 60B..ZF...) from the IFS.

Nominal long-term interest rates (R^L) The series is the Government Bond Yield or equivalent (code 61...ZF...) taken from the IFS. The data are not available for Argentina, Brazil, and Mexico.

Export prices (P^{ex}) The series is the Export Unit Values or Export Price Index taken from the IFS for all countries with the exception of Argentina and Mexico. The code is 74..DZF... or 76...ZF.... Data for Argentina and Mexico are provided from Oxford Economic Forecasts.

Import prices (P^{imp}) The series is the Import Unit Values or Import Price Index from IFS for all countries with the exception of Argentina and Mexico. The code is 75..DZF... or 76.X.ZF.... Data for Argentina and Mexico are provided from Oxford Economic Forecasts.

Oil price (P^o) The series is the price of Brent from IFS, with code 11276AAZZF....

The variables used in the econometric exercise are constructed from the series above as follows.

$y = \ln[100 \cdot (Y/POP)/POP_{2000}]$;
$r = 0.25 \cdot \ln(1 + R_t/100) - \ln(P_{t+1}/P_t)$;
$r^L = 0.25 \cdot \ln(1 + R_t^L/100) - \ln(P_{t+1}/P_t)$;
$q = \ln(100 \cdot E/E_{2000}) - \ln(P)$;
$nfa = F/(Y^{NC}/E)$;
$y_i^* = \sum_{j=0}^{N-1} w_{ij} y_j$;
$r_i^* = \sum_{j=0}^{N-1} w_{ij} r_j$;
$r_i^{L*} = \sum_{j=0}^{N-1} w_{ij} r_j^L$
$tt = \ln(TT)$ where $TT = P^{ex}/P^{imp}$;
$p^o = \ln(100 \cdot P^o/P_{2000}^o)$.

The Euro Area variables are constructed as weighted averages of the corresponding series of Austria, Belgium, Finland, France, Germany, Greece, Ireland, Italy, Netherlands, Portugal, and Spain. The weights are each country's mean shares of the Euro Area's real GDP in PPP over the period 1995-2000. The real GDP in PPP series are obtained from the World Bank's World Development Indicators 2002.

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