FISCAL SHOCKS AND REAL EXCHANGE RATE DYNAMICS: SOME EVIDENCE FOR LATIN AMERICA

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Abstract

This paper analyses the effects of fiscal shocks using a two-country macroeconomic model for output, labour input, government spending and relative prices which provides the orthogonality restrictions for obtaining the structural shocks. Dynamic simulation techniques are then applied, in particular to shed light on the possible effects of fiscal imbalances on the real exchange rate in the case of six Latin American countries. Using quarterly data over the period 1980-2006, we find that in a majority of cases fiscal shocks are the main driving force of real exchange rate fluctuations.

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1. Introduction

Since the seminal contribution of Krugman (1979), it is well known among international economists that most of the Latin American (LA) countries suffered speculative attacks on their currencies from international investors mainly because of inconsistencies between domestic macroeconomic policies and the adopted exchange rate regime. In turn, real exchange rate misalignments have often led to macroeconomic disequilibria, and hence the correction of external imbalances might require both demand management policies and real exchange rate devaluations (see, among others, Edwards, 1988). As a result, equilibrium real exchange rates have changed over time, periods of large appreciations being followed by severe depreciations or periods of stability. Furthermore, real exchange rate variability in the LA countries over the eighties was greater than almost anywhere else in the world (Edwards, 1989), owing to debt crises that resulted in a real depreciation of the domestic currency, with frequent devaluations and inflationary episodes.

The existing literature on the sources of real exchange rate fluctuations has typically focused on the role of real demand (Enders and Hurn, 1994), monetary (Clarida and Gali, 1994; Weber, 1997) or productivity (Alexius, 2005) shocks, and has overlooked the possible effects of fiscal unbalances on countries' international competitiveness. Notable exceptions are the studies of Obstfeld (1993) and Asea and Mendoza (1994). Further, only a few studies (Chowdhury, 2004; Hoffmaister and Roldós, 2001; Rodríguez and Romero, 2007) have investigated the sources of real exchange rates fluctuations in emerging economies (and even less in LA countries), mainly assessing the relative contribution of temporary and permanent disturbances.

This paper, unlike previous studies on the exchange rate determination in emerging economies, adopts a framework which allows for a wide range of economically meaningful (structural) shocks potentially affecting the real exchange rate, including fiscal disturbances. While the effects of fiscal policy on the real exchange rate have recently been extensively investigated in the case of industrialised countries (Monacelli and Perotti, 2006; Ravn et al., 2007; Kim and Roubini, 2008), to the best of our knowledge, no studies exist for the LA countries, despite the importance of this issue for emerging economies as well. The present paper is an attempt to fill this gap.

As pointed out by Agénor et al. (2000), macroeconomic fluctuations in developing countries are related to those in industrial economies, and these linkages may have important policy implications for stabilisation and adjustment programmes (Agénor and Montiel, 1996). Accordingly, we employ a two-country macroeconomic model for output, labour input, government spending and relative prices, along the lines of the studies by Ahmed et al. (1993) and Hoffmaister and Roldós (2001), where the modelling approach to macroeconomic fluctuations developed by Blanchard and Quah (1989) is extended to an open-economy setting allowing for the possible existence of cointegration relationships among the variables of the system. The theoretical model consists of four blocks linked to each other according to a quasi-recursive scheme, and provides the orthogonality restrictions to be imposed to achieve the identification of the structural shocks. These disturbances are identified as supply-side (relative productivity and relative labour inputs) and demand-side (relative fiscal and relative preference) shocks. Their dynamic effects on the real exchange rate are then examined within a structural Vector Error Correction (VEC) framework by means of dynamic simulation and historical decomposition techniques.

Applying the same theoretical framework to a relatively homogeneous sample of countries of the same area (namely the LA region), and including the US economy in the analysis as the most appropriate proxy for foreign factors, enables one to establish whether there are empirical regularities across this set of countries, despite their historically different experiences (Ahmed, 2003). Using quarterly data over the period 1980-2006, we provide clear evidence that fiscal shocks are a key determinant of real exchange rate dynamics for most of the LA countries we consider, suggesting that appropriate fiscal policy measures are crucial to enhance the international competitiveness of these economies. By contrast, monetary factors appear to account for a relatively small fraction of the overall real exchange rate variability. These results are robust across a number of alternatives specifications of the empirical model. Further, we show that the contribution of demand (and monetary) shocks to explaining real exchange rate fluctuations increases when shorter cyclical fluctuations are taken into account.

Finally, omitting the cointegration relationships, which we show exist, is found to lead to overestimating the role of demand shocks and, most importantly, underestimating the contribution of fiscal disturbances.

The layout of the paper is as follows. Section 2 describes the econometric model. Section 3 presents the empirical results. In Section 4, dynamic simulations based on forecast error variance and historical decompositions are discussed, while robustness analysis is reviewed in Section 5. Some final remarks follow in Section 6.

1. The model

In recent years, the macro-economic effects of fiscal shocks have been extensively studied (Hemming et al., 2002), even though the current empirical evidence has mainly analysed the case of developed countries within a closed-economy setup.[1] On the other hand, the body of research on the international transmission of fiscal policy has been almost exclusively theoretical (Baxter, 1992; Bianconi and Turnovsky, 1997; Obstfeld and Rogoff, 1995), with a few exceptions for selected industrialised countries (Monacelli and Perotti, 2006; Ravn et al., 2007; Arin and Koray, 2008; Kim and Roubini, 2008).

This Section presents a two-country model which provides the theoretical framework to quantify the role of supply and demand shocks (with particular emphasis on fiscal disturbances) in explaining the fluctuations of the real exchange rate (vis-à-vis the US dollar), one of the most common indicators of international competitiveness, for the case of six LA countries (namely, Argentina, Bolivia, Brazil, Chile, Mexico and Peru). This allows us to go beyond the dichotomy between permanent/supply and transitory/demand

[1] Despite this growing empirical literature, there is still no consensus on the size of the effects of fiscal shocks on output or the real exchange rate, mostly because of the difficulties in identifying fiscal disturbances (Mountford and Uhlig, 2005). The narrative approach (Romer and Romer, 1989; Burnside et al. 2004, and Christiano, et al. 1999, among others) makes it possible to circumvent potentially controversial identifying assumptions typical of the VAR method (Mountford and Uhlig, 2005; Favero, 2002; Blanchard and Perotti, 2002, among others), but it has the drawback that these episodes could be in part anticipated or that substantial fiscal shocks, of different type or sign, could have occurred around the same time (Perotti, 2002).

shocks explored in the literature hitherto (see, among others, Chowdhury, 2004, and Rodríguez and Romero, 2007).

2.1 Economic relationships

In line with previous empirical papers on LA countries, we assume that the US economy is the relevant foreign country (Berg et al., 2002; Ahmed, 2003). Following Ahmed et al. (1993) and Hoffmaister and Roldós (2001), we rely exclusively on a quasi-recursive identification scheme based on long-run restrictions which are motivated by a simple Real Business Cycle (RBC) model with an exogenously given fiscal sector in each country. In what follows, the subscript i (j) indicates domestic (US) variables, while t indexes time. Lowercase variables stand for logarithmic transformations.

We start by defining a production function for labour (N_{st}) and capital stock (K_{st}), $Y_{st} = F(K_{st}, A_{st} \cdot N_{st})$ (Garratt et al., 2003) re-written as $\frac{Y_{st}}{N_{st}} = A_{st} \cdot f(\kappa_{st})$, where $s = i, j$, alternatively, $f(\kappa_{st}) = F(K_{st}, 1)$ is a function that satisfies the Inada conditions and $\kappa_{st} = K_{st} / (A_{st} \cdot N_{st})$ indicates the capital stock per effective labour unit. Assuming that the logarithm of the technological progress index A_{st} is given by $\ln(A_{st}) = \alpha_{1s} + \alpha_{2s}\theta_{st}^{y}$ where θ_{st}^{y} is a mean-zero $I(1)$ process and α_2's are country-specific measures of the ability to use technology, the production function becomes (in logs):

$$y_{st} = \alpha_{1s} + \ln\left[f(\kappa_{st})\right] + \alpha_{2s}\theta_{st}^{y}\ ,\ s = i, j \tag{1}$$

If the capital-labour-ratio is constant in the long run, as in Binder and Pesaran (1999), relative labour productivity can be expressed as:

$$\pi_t \equiv (y_{it} - n_{it}) - (y_{jt} - n_{jt}) = \alpha + \phi_t \tag{2}$$

where $\alpha \equiv \ln[f(\kappa_i)] - \ln[f(\kappa_j)] + (\alpha_{1i} - \alpha_{1j})$ and $\phi_t = (\alpha_{2i}\theta_{it}^{y} - \alpha_{2j}\theta_{jt}^{y})$ represents the relative technology shock.

In the long run labour inputs are expected to respond to country-specific exogenous shocks originating in the labour market and/or from permanent changes in government supply policies. Accordingly, we can write down the following functional form for both labour input levels:

$$n_{it} = \beta_{1i} + \beta_{2i}\theta^n_{it} \ , \ n_{jt} = \beta_{1j} + \beta_{2j}\theta^n_{jt}$$

where the β_1's indicate deterministic components, and the θ^n's represent idiosyncratic labour-supply disturbances. Hence, the relative employment level, n_t, can be expressed as:

$$n_t \equiv n_{it} - n_{jt} = \beta + \upsilon_t \tag{3}$$

where $\beta = (\beta_{1i} - \beta_{1j})$ and $\upsilon_t = (\beta_{2i}\theta^n_{it} - \beta_{2j}\theta^n_{jt})$ is the relative labour-supply shock.

Having defined the stochastic disturbances driving relative labour productivity and relative labour inputs, we move on to modelling the public sector of the two economies. Let $\tilde{g}$ be government size (defined as the ratio of government purchases of goods and services to output); taking the (log of) private output (the difference between total output and government spending) in the two economies, y^P, and using the approximation $\ln(1-x) \cong x$ we obtain the following relationships:

$$y_{it}{}^P = y_{it} - \tilde{g}_{it} \ , \ y_{jt}{}^P = y_{jt} - \tilde{g}_{jt} \tag{4}$$

As in Ahmed et al. (1993), the size of domestic (foreign) government depends both on domestic and foreign permanent fiscal policy shocks, the θ^{y^P} parameters, through a feedback reaction function governed by the γ_2's which measure the response to an exogenous change in the foreign (domestic) government size:

$$\tilde{g}_{it} = \gamma_{1i} + \theta^{y^P}_{it} + \gamma_{2i}\theta^{y^P}_{jt} \ , \ \tilde{g}_{jt} = \gamma_{1j} + \theta^{y^P}_{jt} + \gamma_{2j}\theta^{y^P}_{it} \tag{5}$$

where the γ_1's are constant quantities. Using equations (5) to substitute into (4) we then obtain:

$$y_{it}{}^P = \gamma_{1i} + y_{it} + \theta^{y^P}_{it} + \gamma_{2i}\theta^{y^P}_{jt} \ , \ y_{jt}{}^P = \gamma_{1j} + y_{jt} + \theta^{y^P}_{jt} + \gamma_{2j}\theta^{y^P}_{it}$$

or, in relative terms:

$$z_t \equiv y_{it}{}^P - y_{jt}{}^P = (y_{it} - y_{jt}) + (\varphi_t - \gamma)$$

where $\gamma = \gamma_{1i} - \gamma_{1j}$ and $\varphi_t = [(1-\gamma_{2i})\theta_{jt}^{y^P} - (1-\gamma_{2j})\theta_{it}^{y^P}]$ represents the relative fiscal shock.[2] Using conditions (2) and (3), we can express relative private output as a linear function of the structural shocks:

$$z_t = \alpha + \beta - \gamma + \phi_t + \upsilon_t + \varphi_t \tag{6}$$

Finally, consumers in both economies are assumed to make their consumption decisions to maximise their utility. Adopting a log-linear specification with identical preferences in the two countries, the closed-form solution is such that the (log of) relative prices, q_t, equals the marginal rate of substitution (Ahmed et al., 1993). In turn, the balanced-growth path implies that the ratio of world consumption of each good to total private output of that good is constant (d), ensuring that the following condition holds:

$$q_t = \delta + (\theta_{jt}^q - \theta_{it}^q) + (y_{it}^P - y_{jt}^P) \tag{7}$$

where the θ^q's are time-varying preference shocks entering the agents' utility function. Let $\eta_t = (\theta_{jt}^q - \theta_{it}^q)$ be the relative preference shock. Combining (6) and (7), we can express the real exchange rate as:

$$q_t = \alpha + \beta - \gamma + \delta + \phi_t + \upsilon_t + \varphi_t + \eta_t \tag{8}$$

Equation (8) represents real exchange rate dynamics as a combination of the underlying disturbances, which are left unrestricted to encompass a large number of competing theories of real exchange rate determination. Choosing a theory rather than another is thus an empirical issue to be decided by the data. Suppose, for instance, that supply-side shocks dominate the dynamics of the q_t variable. This would support empirically the Harrod-Balassa-Samuelson (HBS) view of real exchange rate

[2] Note that in our model we do not consider explicitly the permanent increase in taxes implied by increases in government size determined by fiscal shocks, which are measured by the θ^{y^P} parameters. As Kneller et al. (1999) point out, empirical studies that include only government expenditure and no taxes may be mis-specified. In order to avoid negative wealth effects of increased government spending (Baxter and King, 1993; Linneamann and Schabert, 2003), it is possible to assume, as in Ahmed et al. (1993), that changes in taxation have no effects on total output in the long-run. This can happen if wealth and the substitution effects of higher taxes on labour supply cancel out, as in a RBC model with a Cobb-Douglas production function, constant relative risk aversion preferences and a taxation proportional to output.

determination.[3] Consider, instead, the case where φ_t turns out to be the most relevant source of real exchange rate fluctuations. This would give empirical support to the model of Roldós (1995), within which public spending shocks can lead to permanent shifts in the real exchange rate. Next suppose that preference shocks are the main driving factor of q_t. This would be consistent with a general equilibrium, two-country model with a representative utility-maximising agent in the presence of cash-in-advance constraints (Stockman, 1980; Lucas, 1982). Clearly, any of the above-mentioned theoretical hypotheses could be a plausible explanation for the behaviour of the real exchange rate in the LA countries. However, were q_t to depend only on constant terms, this would put into question the empirical validity of the purchasing power parity (PPP) hypothesis, and would be more difficult to rationalise. Recent surveys covering this issue are Sarno and Taylor (2002) and Taylor (2006).

2.2 Steady-state of the model

We assume that the four variables (relative productivity, relative labour input, relative private output and real exchange rate) are driven by three common stochastic trends (ϕ_t, υ_t and φ_t) in the long-run. These trends evolve over time according to the following laws of motion:

$$\phi_t = \phi_{t-1} + \varepsilon_t^{\phi} = \phi_0 + \sum_{i=1}^{t} \varepsilon_i^{\phi}, \ \upsilon_t = \upsilon_{t-1} + \varepsilon_t^{\upsilon} = \upsilon_0 + \sum_{i=1}^{t} \varepsilon_i^{\upsilon}, \ \varphi_t = \varphi_{t-1} + \varepsilon_t^{\varphi} = \varphi_0 + \sum_{i=1}^{t} \varepsilon_i^{\varphi}$$

where ϕ_0, υ_0 and φ_0 denote initial conditions and the ε's are uncorrelated white-noise processes such that $E(\varepsilon_t^l) = 0$, $E(\varepsilon_t^l)^2 = \sigma_{\varepsilon^l}^2$, $E(\varepsilon_t^l \varepsilon_s^l) = 0$ for $s \neq t$, with $l = \phi, \upsilon, \varphi$. The model also contains the transitory stochastic component η_t, which is assumed to be orthogonal with respect to ε_t^{ϕ}, ε_t^{υ} and ε_t^{φ} and obeys the following law of motion:

$$\eta_t = \rho\eta_{t-1} + \varepsilon_t^{\eta} = \varepsilon_t^{\eta} / (1 - \rho L) \ , \ \rho < |1|$$

where ε_t^{η} is an uncorrelated white noise process.

[3] See Alexius (2005) for the empirical content of this paradigm for emerging and industrialised economies.

To find the steady state of the model, the initial values of all permanent shocks (ϕ_0, υ_0 and φ_0) along with the deterministic component of all the variables of the theoretical model (α, β, γ, δ) are set equal to zero. Accordingly, the steady state can be represented as follows:

$$\begin{bmatrix} \pi_t \\ n_t \\ z_t \\ q_t \end{bmatrix} = \begin{bmatrix} 1 & 0 & 0 \\ 0 & 1 & 0 \\ 1 & 1 & 1 \\ 1 & 1 & 1 \end{bmatrix} \cdot \begin{bmatrix} \phi_t \\ \upsilon_t \\ \varphi_t \end{bmatrix} \tag{9}$$

The long-run structure (9) implies that not only shocks originating from the supply-side of the economy but also demand shocks (namely, fiscal shocks) can induce permanent shifts of the steady state of the system. By contrast, relative preference shocks are assumed to have transient effects on the levels of the variables. This assumption can be rationalised in terms of the transitory nature of shocks driving demand for domestic and foreign (aggregate) goods.[4] Note that our framework allows for different representations without changes in the causality ordering of the variables and any loss in generality, with the restrictive assumption of cointegration not being strictly required to achieve identification.[5] On the other hand, testing for cointegration is a relevant empirical issue in modelling real exchange rate fluctuations (Alexius, 2005).

2.3. Identification of the structural shocks

Equations (2)-(3)-(6)-(8) represent the building-blocks to study the interactions between domestic and foreign economies. Adopting the same notation as above, we focus on the following k-dimensional VAR model in error correction form:

$$\begin{bmatrix} \Delta\pi_t \\ \Delta n_t \\ \Delta z_t \\ \Delta q_t \end{bmatrix} = c + \Pi \begin{bmatrix} \pi_{t-1} \\ n_{t-1} \\ z_{t-1} \\ q_{t-1} \end{bmatrix} + \sum_{i=1}^{p-1} \Gamma_i \begin{bmatrix} \Delta\pi_{t-i} \\ \Delta n_{t-i} \\ \Delta z_{t-i} \\ \Delta q_{t-i} \end{bmatrix} + \begin{bmatrix} u_t^{\pi} \\ u_t^{n} \\ u_t^{z} \\ u_t^{q} \end{bmatrix}, \; u_t \sim N(0, \Sigma_u) \tag{10}$$

[4] As discussed below, the data are broadly consistent with the empirical specification outlined in this Section.

[5] For instance, allowing for permanent shifts in demand between domestic and foreign goods would amount to introducing an additional stochastic trend into the system (Ahmed et al., 1993; Hoffmaister and Roldós, 2001). The model would then exhibit four common trends and no cointegration among the variables.

where c is a vector of deterministic components, the Γ's are matrices of autoregressive parameters, Δ is the first difference operator and the vector $u_t = [u_t^{\pi} \quad u_t^{n} \quad u_t^{z} \quad u_t^{q}]'$ contains the estimated residuals. Given our theoretical assumptions, we expect the long-run matrix Π to have rank one, i.e. the presence of one cointegrating vector in model (10).

Structural identification is achieved following the common trends methodology (Warne, 1993). Omitting the deterministic component, the reduced-form moving average (MA) representation of the model defines the data generating process (DGP) as a function of the initial conditions (set equal to zero for the sake of exposition) and of the reduced-form shocks u's. This is given by:

$$x_t = C\sum_{i=1}^{t} u_i + C^*(L)u_t$$

where the matrix C measures the impact of cumulated shocks to the system, $C^*(L)$ is an infinite polynomial in the lag operator L, and $u_t = [\pi_t \quad n_t \quad z_t \quad q_t]'$.

The reduced form and the structural residuals are linked through the relationship:

$$u_t = B\varepsilon_t \tag{11}$$

where B is a non-singular matrix (Warne, 1993). Hence, the structural MA representation is the following:

$$x_t = \Phi\sum_{i=1}^{t} \varepsilon_i + \Phi^*(L)\varepsilon_t \tag{12}$$

where the matrix $\Phi = CB$ represents the permanent component of the model, and the matrix polynomial $\Phi^*(L) = C^*(L)B$ the transitory or cyclical component. Structural identification allows to decompose each of the four time series into the sum of distinct components driven by structural shocks. Focusing on the real exchange rate, q_t, we have $q_t = q_t^{\phi} + q_t^{\upsilon} + q_t^{\varphi} + q_t^{\eta}$ with:

$$\begin{aligned} q_t^{\phi} &= \Phi_{41}\sum_{i=1}^{t}\varepsilon_i^{\phi} + \sum_{i=1}^{t}\Phi_{i,41}^{*}\varepsilon_i^{\phi}, \quad q_t^{\upsilon} = \Phi_{42}\sum_{i=1}^{t}\varepsilon_i^{\upsilon} + \sum_{i=1}^{t}\Phi_{i,42}^{*}\varepsilon_i^{\upsilon}, \\ q_t^{\varphi} &= \Phi_{43}\sum_{i=1}^{t}\varepsilon_i^{\varphi} + \sum_{i=1}^{t}\Phi_{i,43}^{*}\varepsilon_i^{\varphi}, \quad q_t^{T} = \sum_{i=1}^{t}\Phi_{i,44}^{*}\varepsilon_i^{\eta}, \end{aligned} \tag{13}$$

respectively, where Φ_{jk} is the element in the j-th row and k-th column in Φ, and $\Phi^*_{i,jk}$ that in the j-th row and k-th column in the matrix Φ^*_i which forms the polynomial $\Phi^*(L)$ in (12).

The decomposition (13) makes it possible to assess to what extent each of the four stochastic elements included in the model contributes to explaining the evolution of the real exchange rate (and the other variables of the system) over time. Once the model has been identified, dynamic simulations (such as forecast error variance decomposition and impulse analysis) and historical decomposition can be performed.

3. Data and estimation results

3.1. Data and preliminary analysis

Quarterly observations over the period 1980q1-2006q4 are used. Data for the nominal exchange rate (E), defined as national currency per US dollar, consumer price index (P) and real GDP (Y) are from the IMF's International Financial Statistics (IFS) database (code AE…ZF, 64…ZF and 99BVP…RZF, respectively). For Argentina and Brazil these series were obtained from Datastream. Employment levels (N), measured in thousand of employees, are taken from Datastream for all countries. Finally, the shares of government expenditure in good and services (G)[6] are from Penn World Table 6.2. When quarterly observations are not available, annual data have been interpolated to create quarterly series using the Chow and Lin (1971) method. Finally, seasonal adjustment has been carried out using TRAMO/SEATS. Private output is obtained by multiplying the level of real GDP by the share of private output calculated as $(1-G)$. The real exchange rate (Q) is defined as E times the ratio between US and domestic prices. Thus, an increase in Q means a real depreciation. All variables are expressed in constant prices (base year 2000=1). Table 1 below provides further details.

Table 1

[6] This methodology is consistent with those used, for example, in Blanchard and Perotti (2002) and Perotti (2002).

As a preliminary analysis, we performed standard ADF (Dickey and Fuller, 1979) unit root tests on (the log of) each variable. The deterministic component includes an intercept and, when statistically significant, a linear trend. The number of lags is chosen such that no residual autocorrelation is evident in the auxiliary regressions. In all cases we are unable to reject the unit root-null hypothesis at conventional nominal levels of significance. On the other hand, differencing the series appears to induce stationarity. The PP (Philips and Perron, 1988) unit root test and the KPSS (Kwiatkowski et al. 1992) stationarity test corroborate these results.[7]

Table 2

3.2. VEC model estimates

The order of autoregression of the models is chosen according to the Akaike Information Criterion (AIC), setting the maximum lag length equal to eight. The autoregression order turns out to be two for Mexico, three for Chile, four for Argentina and Brazil, five for Peru and eight for Bolivia. System misspecification tests (not reported to save space) suggest no traces of heteroscedasticity and serial correlation.[8] Departures from normality are detected in all models. However, as pointed out by Lee and Tse (1996), the maximum likelihood approach to cointegration developed by Johansen (1995) produces testing procedures which are fairly robust to the presence of non-normality.

The number of cointegration vectors is determined on the basis of the trace test statistics of Johansen (1992). Their critical values are taken from Osterwald-Lenum (1992). Table 3 presents the results. The trace test indicates the presence of one cointegration relationship in all models at the 5 percent level of significance, except in the case of Bolivia where it suggests choosing a rank of two, but a single long-run equilibrium condition at the 1 percent level. These results are broadly consistent with our

[7] Results from the PP and KPSS tests are not reported to save space, but are available from the authors upon request. As pointed out by Boschi and Girardi (2008), the actual integration properties of the real exchange rate series of the LA countries (and Mexico in particular) are likely to depend on the choice of the relevant foreign economy, which explains why standard unit root tests tend to yield mixed results.

[8] Only in the case of Bolivia are there symptoms of autocorrelation, mainly in the equations for relative productivity and relative labour services.

a priori theoretical assumptions about the existence of (at least) three common stochastic trends driving each system.[9]

Table 3

Structural residuals are then extracted from the reduced-form disturbances by imposing (at least) $k^2 = 16$ restrictions on the elements of matrix B in equation (11). A first set of constraints is obtained by assuming that the structural shocks are orthonormal: this implies $k(k+1)/2 = 10$ (non-linear) restrictions. The choice of the cointegration rank allows to distinguish transitory shocks from permanent innovations and produces additional $r(k-r) = 3$ restrictions; in our case, there are *four* zero restrictions in the 4×3 matrix in (9), producing an over-identified structure, which can be tested by means of the usual χ^2-distributed likelihood ratio (LR) tests. The statistics for Argentina, Bolivia, Chile and Mexico turn out to be 1.28, 0.68, 0.37 and 1.60 respectively; by contrast, in the case of Brazil and Peru, their value is 263.20 and 140.81, respectively. By comparing these test statistics with the critical values of a χ^2 distribution with one degree of freedom, we are unable to reject the null of the validity of the over-identifying restriction only for the first four models. Accordingly, we impose the over-identified structure in the case of Argentina, Bolivia, Chile and Mexico, while for the Brazilian and Peruvian systems we employ a just-identified structure.

4. Evidence from the baseline specification

Once structural and data-consistent identification of the VEC models is achieved, dynamic simulations as well as historical decomposition exercises are performed so as to address two main issues: first, to quantify the role played by the underlying (structural)

[9] Notice that the VAR specification considered here is model $H_1^*(r)$ in Johansen's notation, where a linear deterministic trend is implicitly allowed for, but this can be eliminated by the cointegrating relations so that the process contains no trend stationary components. The maximum eigenvalue test statistics indicate one cointegrating relationship only for three countries (Argentina, Bolivia and Peru), while in the other models (Brazil, Chile and Mexico) there is evidence of four common stochastic trends. In general, we favour the conclusions of the trace test in line with Johansen (1992), according to which the maximum eigenvalue test may produce a non-coherent testing strategy. All results are available on request.

sources in explaining the fluctuations of the variables in each country model (Section 4.1); second, to analyse the contribution of each structural shock in driving real exchange dynamics over the sample period under investigation (Section 4.2).

4.1 Sources of system-wide and variable fluctuations

We assess the relative contribution of the structural shocks in explaining macroeconomic fluctuations by means of forecast error variance decomposition (FEVD) analysis. Table 4 reports the percentage of the variance which can be attributed to each structural shock for the individual variables of the model as well as for the system as a whole (row labelled as "system") over a simulation horizon of 20 quarters. Aggregating the shocks, we also consider the role of supply shocks (ϕ and υ) and demand disturbances (φ and η).[10]

Table 4

Supply shocks are the most relevant source of macroeconomic fluctuations in all systems. Their contribution ranges from more than 70 percent in Argentina, Brazil and Mexico to around 60 percent in Bolivia. This finding is broadly consistent with the empirical evidence for developed economies.[11] A closer look at the level of individual variables shows the existence of three distinct groups of countries. The results for Argentina, Bolivia and Mexico reveal that productivity shocks are the main driving forces of relative productivity and relative private output variability, while relative labour services and the real exchange rate fluctuations are mainly governed by labour input and fiscal shocks, respectively. By contrast, while fiscal shocks still represent the main driving forces of the variability of international competitiveness in the Chilean economy,

[10] Several studies have analysed the role of demand shocks (for instance, monetary and fiscal policies) and supply disturbances (productivity and labour supply shocks or structural restructuring policies, such as tariff and trades reforms) in a closed-economy context, both at the aggregate (Blanchard and Quah, 1989; Gali, 1999, among others) and, more recently, at the disaggregate level (Chang and Hong, 2005; Busato et al., 2005), for the US economy or other developed countries. Here, instead, we focus on the aggregate fluctuations in a two-country model for emerging economies.

[11] Bergman (1996), for instance, using a bivariate VAR model for output and inflation, shows that more than one half of the macroeconomic fluctuations in the G7 countries are due to supply shocks at the typical business cycle frequency (the twenty-quarter horizon).

relative preference (labour input) disturbances turn out to drive variability in the dynamics of relative private output (productivity and labour services). Finally, in just-identified structures (Brazil and Peru), we observe that relative productivity and relative labour services fluctuations originate from productivity shocks, with labour input and fiscal shocks dominating the variability of real exchange rates and private output changes.

Focusing on the main variable of interest in our analysis, i.e. the real exchange rate, we find evidence of a difference in behaviour between over-identified and exactly identified systems: in the former class of models international competitiveness is driven by the demand-side of the economy, whilst in the latter group of countries the real exchange rate responds mostly to supply-side disturbances. Further, fiscal shocks are the main driving force of real exchange rate movements in the majority of cases (Argentina, Bolivia and Mexico), ranging from 60 to 90 percent, while they are less relevant for Chile and Peru, even though their effects are still sizeable (35 and 21 percent, respectively). Only in the case of Brazil is the contribution of this shock negligible. [12,13]

4.2. Explaining real exchange dynamics in the LA countries over the years 1980-2006

The existence of a stable long-run relationship among the variables of each model does not prevent the relative weight of structural factors from changing over time in response to complex and interrelated reciprocal influences. Hence, it could be instructive

[12] FEVD decomposition results for each simulation quarter (not reported to save space) indicate that the contribution of fiscal shocks is slightly increasing over the simulation horizon and always statistically significant according to Monte Carlo standard errors computed with 1000 replications in all models (except for the case of Brazil).

[13] The results from the impulse response analysis (not reported for the sake of brevity) suggest a close relationship between relevance of fiscal shocks as a driving source of real exchange rate fluctuations and effects of unanticipated fiscal shocks on the level of international competitiveness in the LA economies. In all models, but the one for Argentina, we find that an increase in government spending leads to a real depreciation. The sign of the response of international competitiveness to this type of shock, however, cannot be determined *ex-ante* as clearly, since it depends on a wide range of factors. On this topic, see also Arin and Koray (2008).

to examine the hypothetical time path of international competitiveness if all disturbances had been associated to only one source of shock.

Table 5 summarises the OLS estimation results obtained by regressing changes in the real exchange rate on its component driven by individual orthogonal shocks. Since structural components are mutually orthogonal by construction, the total variation of the regressand (the measure of international competitiveness) must be *fully* captured by the explanatory variables (supply shocks, ϕ and υ, and demand disturbances, φ and η).

Table 5

The results indicate that the in-sample variability of the real exchange rate is dominated by demand shocks in most of the models, with percentages ranging from 39 (Chile) to 92 (Bolivia). In particular, for five out of the six countries (Brazil being the only exception), fiscal shocks account for a considerable percentage of real exchange rate movements, ranging from one-fifth (for Peru) to four-fifth (for Mexico) of total variability. Also, notice that in most cases (Argentina, Bolivia, Mexico and Peru) the effects of fiscal impulses are stronger than those of productivity shocks. Finally, the relative importance of the temporary components (namely, preference shocks) varies across countries, being at its highest in Brazil, where it explains 43 percent of the historical variance (the effects of fiscal shocks being negligible), and in Peru, where the corresponding share is 34 percent, whilst is countries such as Mexico and Chile it is as low as 6 percent. These findings seem at odds with the conclusions in Rodriguez and Romero (2007), who find a dominant effect of the variability of transitory (real) components on the behaviour of the real exchange rate in Argentina (Brazil).

In order to check for possible shifts in the relative explanatory contribution of shocks for real exchange rate changes over the sample span, we employ the estimated models to replicate the same exercise as above over the window embracing the period from the first available observation to 1994*q*4 and then extending it by a datapoint at a time. Summary statistics (mean, standard error of the mean, minimum and maximum values) for each system are reported in Table 6.

Table 6

The results broadly confirm the previous evidence in a number of ways. First, fiscal shocks are the most relevant source of variation for real exchange rates in the over-identified models. Second, in all models, the mean values of each shock resulting from the recursive procedure are *quantitatively* very close to their full-sample counterparts and *qualitatively* similar to the results from the forecast error variance decomposition exercise. Third, the standard error of the mean, as well as the minimum-maximum range, suggest that the relative contribution of the four driving forces in explaining real exchange rate changes is almost constant over time.

The last piece of evidence concerns the relationship between structural shocks and the pattern over time of the level of international competitiveness. Figure 1 shows, for each country, the real exchange rate series purged of the deterministic part (solid line), and its component explained by the fiscal shocks (dashed line).

Figure 1

The effects of fiscal shocks in the period 1981-1986 turn out to be considerable for all the countries under examination. After this period, however, this is still the case only for Bolivia and Mexico, while in Chile and Peru long swings in the real exchange rate are only partially caused by the fiscal components. Consistently with the previous results, fiscal shocks do not appear to have significant explanatory power in the case of Brazil.

5. Further evidence

5.1. Alternative specifications

It is widely recognised that the results from structural VAR models relying on long-run restrictions may vary considerably depending on the exact specification of the empirical model. Therefore, in this Section we study the robustness of the results discussed above with respect to changes in the empirical specification of the systems.

Three alternative empirical specifications are estimated in order to investigate how the relative weights of demand shocks (and in particular fiscal shocks) vary with the frequency of the fluctuations. Specifically, we filter the data using in turn first differences, FD, the HP filter (Hodrick and Prescott, 1997), HP, and linear detrending, LD. In particular, FD series are used to isolate short cycle fluctuations, HP-filtered series for intermediate frequencies and LD series for low frequencies. We expect the role of demand shocks to decrease with the persistence of shocks.[14] Notice that all alternative specifications neglect the existence of possible cointegration relationships. Thus, our robustness checks can shed light on the consequences of ignoring the presence of long-run equilibrium relationships between the variables.

Table 7 presents the results from imposing the over-identifying long-run restriction in the three alternative empirical specifications. p -values are in square brackets.

Table 7

Overall, the long-run structure implied by our theoretical relationship of reference is not rejected by the data in nine (one at the 1 percent, one at the 5 percent and the remaining seven at the 10 percent level of significance) out of eighteen cases. In particular, the outcome from the FD specification is fully consistent with the baseline design, even though the test statistics are slightly less supportive of our economic priors. In the present context, this conclusion is not surprising since the FD specification produces loss of relevant information, in the presence of documented cointegration relationships. Notice, further, that in the LD specification we observe the rejection of the null hypothesis in all models but one (the Chilean case).

Following the same criterion as in the previous Section, we perform a FEVD analysis under the over-identified structure for the specifications where the over-identifying restriction holds, but employing the just-identified structure when the constraint imposed on the long-run matrix is rejected by the data. The simulation horizon is set equal to 20 quarters. Table 8 shows the contribution (in percentage terms) of aggregate demand shocks and fiscal shocks to the overall forecast error variance of the real exchange rate under the three alternative empirical specifications.

Table 8

As expected, in most cases the relative importance of demand shocks is stronger in the specification where the short-run cycle frequency, FD, is isolated, and decreases when a longer cyclical component is taken into account, that is when we move from the HP to the LD specification.

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[14] The FD specification is the baseline model (10) with $\Pi = 0$ and with four common trends. Such a specification is consistent with the conclusions of the maximum eigenvalue test for Brazil, Chile and Mexico.

Comparing these results to those from the VEC models, we observe that the explanatory power of demand shocks under the alternative specifications is greater than that of their counterparts from the baseline specification with cointegration, with relative preference shocks now becoming the most important source for real exchange rate fluctuations. As shown by Alexius (2005), the lack of the long run-equilibrium conditions between fundamental variables and the real exchange rate eliminates the relationship between the latter and productivity disturbances. Thus, the relative system impact of supply disturbances tends to decrease. In addition, if the long-run properties of the system are not properly taken into account, the effects of fiscal shocks are underestimated, as the relationship between government size and the dynamics of the real exchange rate is overlooked: on average, the share of overall real exchange rate variability explained by fiscal shocks is as high as 48 percent in the baseline framework, while it goes from 12 percent in the FD specification to 9 percent in the LD specification.

5.2. Extensions: the role of monetary shocks

As monetary and fiscal policy are interrelated, especially in the LA countries, we also incorporate monetary shocks allowing for monetary neutrality, i.e. no long-run real effects of money though a possibly high degree of persistence. Therefore, in order to check the sensitiveness of our findings with respect to the omission of monetary factors, we augment our baseline specification by including the following equation:

$$m_t = (m_{it} - m_{jt}) = \iota + \phi_t + \upsilon_t + \varphi_t + \eta_t + \kappa_t \tag{14}$$

where ι is a deterministic component and $\kappa_t \equiv (\kappa^m_{jt} - \kappa^m_{it})$ is the relative money shock. As in Ahmed et al. (1993), the real sector "comes first" in the money equation so as to ensure that money is neutral in the long run. The augmented model based on equations (2), (3), (6), (8) and (14) is then estimated for all LA countries under investigation.[15]

Cointegration tests yield qualitatively similar results as those reported in Table 3. Accordingly, for each LA country model we estimate a VEC system with a single

[15] Notice that because of data limitations for the money stock, the estimation horizon now embraces the period 1980-2006 only for the models of Argentina and Mexico. The starting year for the model with money for Bolivia, Brazil, Chile and Peru is 1988, 1990, 1982 and 1990, respectively.

cointegration relationship. We also take account of possible alternatives (HP filtering, first differencing and linear detrending methods) to cointegration to achieve stationarity.

In order to perform the FEVD analysis, we employ the over-identified structure for those models where the additional restriction implied by the theoretical model on the recursive structure of the long run matrix is not rejected by the data (at the 1 percent level of significance). In the remaining cases, instead, we use the just-identified structure. Table 9 reports the p-values for imposing that constraint in each country model and in the baseline model as well as in the three alternative specifications. The over-identifying restriction turns cannot be rejected in 10 out of 24 cases. Notice that most of the rejections occur in the models of Brazil and Peru and in the specifications based on linear detrending.

Table 9

In Table 10 we present the average contribution (in percentage terms) of aggregate real shocks and fiscal shocks to the overall forecast error variance of the real exchange rate for the four alternative specifications, over a simulation horizon of 20 quarters.

Table 10

The results are interesting in a number of respects. First, the average contribution of fiscal shocks is larger than that attributable to monetary disturbances, which is consistent with the evidence reported in the literature surveyed in Sarno and Taylor (2002). In the specification with cointegration, the average contribution of monetary shocks is roughly 5 percent, much less than the contribution of fiscal shocks (around one-fourth of the overall variability). This conclusion holds for all specifications and for all models (except for Chile in the VEC model and for Boliva and Peru in the three VAR models). Second, in the VAR models (HP, FD and LD), the role of fiscal shocks in explaining real exchange rate fluctuations appears to decline (around 10 percent), giving support to previous conclusions that neglecting the presence of cointegration leads to a breakdown of the linkage between government spending and the real exchange rate. Third, the role of monetary shocks tends to decrease when larger cyclical fluctuations are taken into account: on average, monetary shocks account for almost 10 percent of total variability in the specification in first differences, but only 5 percent in the LD models. Overall, our findings suggest the presence of deviations from PPP, the dynamics of real exchange rates in LA countries appearing to be driven by various factors, with the real dominating over the monetary ones. As is well known, standard supply-side theories based on HBS effects are usually found not to be fully satisfactory in explaining real exchange rate fluctuations - here we document that a major role is played by fiscal shocks as suggested by Roldós (1995), implying the importance of fiscal policy to improve the international competitiveness of these emerging economies.

6. Conclusions

This paper uses a two-country model to analyse the role of a wide class of underlying (structural) disturbances in driving real exchange rates (defined relative to the US dollar) in six LA countries (Argentina, Bolivia, Brazil, Chile, Mexico and Peru). Specifically, it considers the effects on competitiveness of relative productivity, labour, preference, monetary and fiscal shocks. The role of fiscal shocks in particular had not previously been studied in the case of the LA economies. Moreover, most of the existing literature adopts a simple modelling strategy, relying exclusively on a standard permanent/transitory (supply/demand) decomposition, which only provides partial evidence, as, by construction, it allows for only two types of shocks, ignoring the possibility of a wider class of disturbances hitting the economy (and consequently the real exchange rate as well) that also need to be investigated. Our approach, being much more general, enables us to shed new light on the driving forces of real exchange rate dynamics in emerging economies.

Therefore, our contribution to the literature on fiscal shocks is two-fold. First, we extend the methodology developed in Ahmed et al. (1993) and Hoffmaister and Roldós (2001) so as to identify fiscal shocks in a multicountry/multivariate time series context, allowing for the existence of possible cointegration relationship among the variables of the system. Second, using quarterly data over the period 1980-2006, we present some new empirical evidence for six LA countries, indicating that fiscal shocks are a key determinant of real exchange rate dynamics for most of the economies we consider, and play a more crucial role than monetary factors. Further, using alternative econometric specifications, we show that the relative importance of demand shocks (and in particular of monetary shocks) varies with the frequency of cyclical fluctuations isolated in the models. Specifically, the explanatory power of demand (monetary) shocks increases when shorter cyclical fluctuations are taken into account. Moreover, neglecting the presence of cointegration, which in fact holds in our case, amounts to overlooking the linkage between productivity and government spending and the real exchange rate. As we show, this leads to overestimating the role of demand shocks and underestimating the contribution of fiscal disturbances, putting into question the reliability of some earlier evidence, for which this criticism is relevant (see, e.g. Ahmed et al., 1993; Chowdhury, 2004; Hoffmaister and Roldós, 2001; Rodríguez and Romero, 2007).

Concerning possible extensions of the present study, a weighted average of the other LA countries and the three largest industrialised economies (the US, the euro area and Japan) could be used as the foreign economy in the analysis, instead of the US (see Boschi and Girardi, 2008). Also, capital flows could be included in the model by introducing an uncovered interest parity equation in real terms. These issues are left for future research.

References

Agenor P.R., McDermott C.J. and Prasad E.S. (2000), Macroeconomic Fluctuations in Developing Countries: Some Stylized Facts, *World Bank Economic Review*, 14, p. 251- 285.

Agénor P.R. and Montiel P.J. (1996), *Development Macroeconomics*, Princeton University Press.

Ahmed S. (2003), Sources of Economic Fluctuations in Latin America and Implications for Choice of Exchange Rate Regimes, *Journal of Development Economics*, 72, p. 181-202.

Ahmed S., Ickes B.W., Wang P. and Yoo B.S. (1993), International Business Cycle, *American Economic Review*, 83, p. 335-359.

Alexius A. (2005), Productivity Shocks and Real Exchange Rates, *Journal of Monetary Economics*, 52, p. 555-566.

Arin K.P. and Koray F. (2008), Beggar Thy Neighbour? The Transmission of Fiscal Shocks from US to Canada, *Open Economies Review*, forthcoming.

Asea P.K. and Mendoza E.G. (1994), The Balassa-Samuelson Model: A General-Equilibrium Appraisal, *Review of International Economics*, 2, p. 244-267.

Baxter M. (1992), Fiscal Policy, Specialization, and Trade in the Two-Sector Model: The Return of Ricardo?, *Journal of Political Economy*, 100, p. 713-744.

Baxter M. and King R. (1993), Fiscal Policy in General Equilibrium, *American Economic Review*, 83, p. 315-334.

Berg A., Borensztein E. and Mauro P. (2002), An Evaluation of Monetary Regime Options for Latin America, *IMF Working Paper*, 211.

Bergman M. (1996), International Evidence on the Sources of Macroeconomic Fluctuations, *European Economic Review*, 40, p. 1237-1258.

Bianconi M. and Turnovsky S.J. (1997), International Effects of Government Expenditure in Interdependent Economies, *Canadian Journal of Economics*, 30, p. 57-84.

Binder M. and Pesaran M.H. (1999), Stochastic Growth Models and their Econometric Implications, *Journal of Economic Growth*, 4, p. 139-83.

Blanchard O. and Quah D. (1989), The Dynamic Effects of Aggregate Demand and Supply Disturbances, *American Economic Review*, 79, p. 655-673.

Blanchard O. and Perotti R. (2002), An Empirical Characterization of the Dynamic Effects of Changes in Government Spending and Taxes on Output, *Quarterly Journal of Economics*, 117, p. 1329-1368.

Boschi M. and Girardi A., The Contribution of Domestic, Regional, and International Factors to Latin America's Business Cycle, Documenti di Lavoro ISAE, 105.

Burnside C., Eichembaum M. and Fisher J.D.M. (2004), Fiscal Shocks and Their Consequences, *Journal of Economic Theory*, 115, p. 89-117.

Busato F., Girardi A. and Argentiero A. (2005), Technology and Non-technology Shocks in a Two-sector Economy, Economics Working Papers, University of Aarhus, 11.

Chang Y. and Hong J.H. (2005), Do Technological Improvements in the Manufacturing Sector Raise Or Lower Employment?, *American Economic Review*, 96, p. 352-368.

Chow G. and Lin A.L. (1971), Best linear unbiased distribution and extrapolation of economic time series by related series, *Review of Economic and Statistics*, 53, p. 372-375.

Chowdhury I.S. (2004), Sources of Exchange Rate Fluctuations: Empirical Evidence from Six Emerging Market Countries, *Applied Financial Economics*, p. 697-705.

Christiano L.J., Eichenbaum M. and Evans C.L. (1999), Monetary Policy Shocks: What Have We Learned and to what End?, in: Taylor J.B. and Woodford M. (eds.), *Handbook of Macroeconomics*, p. 65-148.

Clarida R.H. and Gali J. (1994), Sources of Real Exchange Rate Fluctuations: How Important Are Nominal Shocks?, *Carnegie-Rochester Conference Series on Public Policies*, 41, p. 1-56.

Dickey D.A. and Fuller W.A. (1979), Distribution of the estimators for autoregressive time series with a unit root, *Journal of the American Statistical Association* 74, p. 427-431.

Edwards S. (1988), Real and Monetary Determinants of Real Exchange Rate Behaviour: Theory and Evidence for Developing Countries, NBER Working Paper, 2721.

Edwards S. (1989), Capital Flows, Real Exchange Rates and Capital Controls: Some Latin America Experiences, NBER Working Paper, 6800.

Enders W. and Hurn S. (1994), Theory and Tests of Generalized Purchasing-Power Parity: Common Trends and Real Exchange Rates in the Pacific Rim, *Review of International Economics*, 2, p. 179-190.

Favero C. (2002), How Do European Monetary and Fiscal Authorities Behave?, CEPR Discussion Paper, 3426.

Gali J. (1999), Technology, Employment, and the Business Cycle: Do Technology Shocks Explain Aggregate Fluctuations?, *American Economic Review*, 89, p. 249-271

Garratt A., Lee K., Pesaran M.H. and Shin Y. (2003), A Long-Run Structural Macroeconometric Model of the UK, *Economic Journal*, 113, p. 412-455.

Hemming R., M. Kell and S. Mahfouz (2002), The Effectiveness of Fiscal Policy in Stimulating Economic Activity: A Review of the Literature, IMF Working Paper, 208.

Hodrick R.J. and Prescott E.C. (1997), Postwar U.S. Business Cycles: An Empirical Investigation, *Journal of Money, Credit and Banking*, 29, p. 1-16.

Hoffmaister A.W. and Roldos J.E. (2001), The Sources of Macroeconomic Fluctuations in Developing Countries: Brazil and Korea, *Journal of Macroeconomics*, 23, p. 213-39.

Johansen S. (1992), Cointegration in Partial Systems and the Efficiency of Single-equation Analysis, *Journal of Econometrics*, 52, p. 389-402.

Johansen S. (1995), *Likelihood-Based Inference in Cointegrated Vector Autoregressive Models*, Oxford University Press, Oxford.

Kim S. and Roubini N. (2008), Twin Deficit or Twin Divergence? Fiscal Policy, Current Account, and Real Exchange Rate in the U.S., *Journal of International Economics*, 74, p. 362-383.

Kneller R., Bleaney M.F. and Gemmell N. (1999), Fiscal Policy and Growth: Evidence from OECD Countries, *Journal of Public Economics*, 74, p. 171-190.

Krugman P. (1979), A Model of Balance of Payments Crises, *Journal of Money, Credit, and Banking*, 11, p. 311-325.

Kwiatkowski D., Phillips P.C.B., Schmidt P. and Shin Y. (1992), Testing the Null of Stationarity against the Alternative of a Unit Root: How Sure Are we that Economic Time Series Have a Unit Root?, *Journal of Econometrics*, 54, p. 159-178.

Lee T. and Tse Y. (1996), Cointegration Tests with Conditional Heteroskedasticity, *Journal of Econometrics*, 73, p. 401-410.

Linnemann L. and Schabert A. (2003), Fiscal Policy in the New Neoclassical Synthesis, *Journal of Money, Credit and Banking*, 35, p. 911-929.

Lucas R.E. (1982), Interest rates and currency prices in a two-country world, *Journal of Monetary Economics*, 10, p. 335-359.

MacKinnon J.G. (1996), Numerical Distribution Functions for Unit Root and Cointegration Tests, *Journal of Applied Econometrics*, 11, p. 601-618.

Monacelli T. and Perotti R. (2006), Fiscal Policy, the Trade Balance and the Real Exchange Rate: Implications for International Risk Sharing, mimeo.

Mountford A. and Uhlig H. (2005), What Are the Effects Fiscal Policy Shocks?, SBF Discussion Paper, 39.

Obstfeld M. (1993), Model Trending Real Exchange Rates, Centre for International and Development Economics Research (CIDER) Working Papers, 11.

Obstfeld M. and Rogoff K. (1995), The Mirage of Fixed Exchange Rates, *Journal of Economic Perspectives*, 9 (4), p. 73-96.

Osterwald-Lenum M. (1992), A Note with Quantiles of the Asymptotic Distribution of the Maximum Likelihood Cointegration Rank Test Statistics," *Oxford Bulletin of Economics and Statistics*, 54, p. 461-472.

Phillips P.C.B. and Perron P. (1988) Testing for a Unit Root in Time Series Regression, *Biometrika*, 75, p. 335–346.

Perotti R. (2002), Estimating the Effects of Fiscal Policy in OECD Countries, ECB Working Paper, 168.

Ravn M., Schmitt-Grohe S. and Uribe M. (2007), Explaining the Effects of Government Spending on Consumption and the Real Exchange Rate, mimeo.

Rodríguez G. and Romero I. (2007), The Role of Permanent and Transitory Components in the Fluctuations of Latin-American Real Exchange Rates, *Applied Economics*, p. 1-10.

Roldòs J.E. (1995), Supply-side of Disinflation Programs, *IMF Staff Papers*, 42, p. 158-183.

Romer C.D. and Romer D.H. (1989), Does Monetary Policy Matter? A New Test in the Spirit of Friedman and Schwartz, in Blanchard O.J. and Fisher S.F. (eds.), *NBER Macroeconomics Annual*, MIT Press, p. 121-170.

Sarno L. and Taylor M.P. (2002), Purchasing Power Parity and the Real Exchange Rate, *IMF Staff Papers*, 49, p. 65-105.

Stockman A.C. (1980), A Theory of Exchange Rate Determination, *Journal of Political Economy*, 88, p. 673-698.

Taylor M.P. (2006), Real Exchange Rates and Purchasing Power Parity: Mean-reversion in Economic Thought, *Applied Financial Economics*, 16, p. 1-17.

Warne A. (1993), A Common Trends Model: Identification, Estimation and Inference, IIES Seminar Paper, Stockholm, Stockholm University, 555.

Weber A. (1997), Sources of Purchasing Power Disparities between the G3 Economies, *Journal of the Japanese and International Economies*, 11, p. 548-583.

Table 1 – Construction of the variables

Variable	*Definition*
Relative productivity	$\pi_t^i = (\ln Y_t^i - \ln N_t^i) - (\ln Y_t^{US} - \ln N_t^{US})$
Relative employment	$n_t^i = \ln N_t^i - \ln N_t^{US}$
Relative private output	$z_t^i = \ln[Y_t^i (1 - G_t^i)] - \ln[Y_t^{US}(1 - G_t^{US})]$
Real exchange rate	$q_t^i = \ln E_t^i + (\ln P_t^{US} - \ln P_t^i)$

Note. For each variable the superscript *i* refers to each Latin America country in turn, while the superscript *US* refers to the base country (the US economy). The subscript *t* stands for time.

Table 2 - Unit root tests

	π				*n*				*z*				*q*			
	Levels		First differences		Levels		First differences		Levels		First differences		Levels		First differences	
	DP	TS	DP	TS	DP	TS	DP	TS	DP	TS	DP	TS	DP	TS	DP	TS
Argentina	c,t	-2.16	c	-9.92	c,t	-2.42	c	-4.35	c	-1.94	.	-4.13	c,t	-3.79	c	-3.44
Bolivia	c,t	-2.08	c	-6.45	c,t	-2.01	c	-4.74	c,t	-0.32	c	-1.87	c	-1.46	.	-4.96
Brazil	c	-1.66	.	-6.44	c,t	-2.18	c	-5.19	c	-1.73	.	-3.70	c	-1.24	.	-4.63
Chile	c,t	-2.05	c	-7.94	c,t	-1.76	c	-6.59	c,t	-2.86	c	-3.63	c,t	-1.16	c	-4.45
Mexico	c,t	-1.76	c	-5.73	c,t	-2.06	c	-4.22	c	-1.42	.	-3.95	c	-2.33	.	-11.7
Peru	c,t	-1.52	c	-6.64	c,t	-0.03	c	-5.43	c	-1.49	.	-2.70	c	-1.95	.	-4.74

Note. ADF test statistics for the null hypothesis of a unit root process for the variables in the levels and in first differences are reported in columns "TS". The critical value at the 1 percent level of significance is -4.05 if a constant and a linear trend (*c*,*t*) are included in the regression, -3.49 with only a constant term (*c)* and -2.59 if no deterministic parts (-) are included. At the 5 percent level of significance these values are -3.45, -2.89 and -1.94, respectively (MacKinnon, 1996). The specification of the deterministic component is presented in the column "DP". Definitions of the variables are provided in Table 1.

Table 3 - Cointegration rank

	Lags	Rank			
		0	1	2	3
Argentina	4	66.77	28.53	11.03	1.87
Bolivia	8	69.77	31.16	11.77	0.42
Brazil	4	47.31	22.6	8.3	0.47
Chile	3	48.89	26.05	11.11	2.66
Mexico	2	49.48	27.07	11.47	0.63
Peru	5	56.06	28.68	4.64	0.74

Note. Critical values for the trace test statistics at the 95 percent for rank 0, 1, 2, 3 and 4 are 47.21, 29.68, 15.41 and 3.76, respectively, while at the 99 percent are 54.46, 35.65, 20.04 and 6.65, respectively (Osterwald-Lenum, 1992). The column "Lag" reports the number of lags included in the VAR specification suggested by the AIC.

Table 4 - Forecast error variance decomposition

	Individual shocks				*Nature of shocks*	
			Argentina			
	ϕ	υ	φ	η	*Supply*	*Demand*
$\Delta\pi$	82.80	1.68	11.87	3.65	*84.48*	*15.52*
Δn	0.96	97.46	0.37	1.21	*98.42*	*1.58*
Δz	87.24	11.96	0.69	0.11	*99.20*	*0.80*
Δq	16.13	8.99	60.69	14.19	*25.12*	*74.88*
System	*46.78*	*30.02*	*18.41*	*4.79*	*76.81*	*23.19*
			Bolivia			
	ϕ	υ	φ	η	*Supply*	*Demand*
$\Delta\pi$	50.58	5.43	34.82	9.17	*56.01*	*43.99*
Δn	7.61	91.93	0.12	0.34	*99.54*	*0.46*
Δz	70.68	6.14	19.16	4.02	*76.82*	*23.18*
Δq	3.32	0.71	90.77	5.20	*4.03*	*95.97*
System	*33.05*	*26.05*	*36.22*	*4.68*	*59.10*	*40.90*
			Brazil			
	ϕ	υ	φ	η	*Supply*	*Demand*
$\Delta\pi$	90.77	4.67	0.73	3.83	*95.44*	*4.56*
Δn	67.84	26.20	2.92	3.04	*94.04*	*5.96*
Δz	6.90	17.83	69.03	6.24	*24.73*	*75.27*
Δq	22.73	61.64	0.48	15.15	*84.37*	*15.63*
System	*47.06*	*27.59*	*18.29*	*7.07*	*74.65*	*25.36*
			Chile			
	ϕ	υ	φ	η	*Supply*	*Demand*
$\Delta\pi$	20.06	47.70	6.02	26.22	*67.76*	*32.24*
Δn	11.56	82.91	0.28	5.25	*94.47*	*5.53*
Δz	43.83	6.77	1.69	47.71	*50.60*	*49.40*
Δq	23.76	27.85	34.82	13.57	*51.61*	*48.39*
System	*24.80*	*41.31*	*10.70*	*23.19*	*66.11*	*33.89*
			Mexico			
	ϕ	υ	φ	η	*Supply*	*Demand*
$\Delta\pi$	86.20	0.65	0.97	12.18	*86.85*	*13.15*
Δn	2.49	95.96	0.07	1.48	*98.45*	*1.55*
Δz	70.92	25.75	2.52	0.81	*96.67*	*3.33*
Δq	17.52	1.63	78.35	2.50	*19.15*	*80.85*
System	*44.28*	*31.00*	*20.48*	*4.24*	*75.28*	*24.72*
			Peru			
	ϕ	υ	φ	η	*Supply*	*Demand*
$\Delta\pi$	67.42	22.32	2.86	7.40	*89.74*	*10.26*
Δn	67.32	25.31	3.78	3.59	*92.63*	*7.37*
Δz	7.35	13.03	64.80	14.82	*20.38*	*79.62*
Δq	23.35	39.07	21.05	16.53	*62.42*	*37.58*
System	*41.36*	*24.93*	*23.12*	*10.59*	*66.29*	*33.71*

Note. Average percentage contribution of each structural shock in explaining variable fluctuations over a simulation horizon of 20 quarters. ϕ, υ, φ, η indicate relative productivity, relative labour, relative fiscal and relative preference shocks, respectively. The column "Supply" is the aggregate contribution of ϕ and υ disturbances. The column "Demand" is the aggregate contribution of φ and η disturbances. The row "System" indicates the average contribution of individual shocks and aggregate disturbances, disentangled according to their nature, for the whole system.

Table 5 - Historical decomposition

	Individual shocks				*Nature of shocks*	
	ϕ	υ	φ	η	*Supply*	*Demand*
Argentina	19.07	24.65	35.35	20.93	*43.72*	*56.28*
Bolivia	5.21	2.90	73.36	18.53	*8.11*	*91.89*
Brazil	15.51	39.87	1.25	43.37	*55.38*	*44.62*
Chile	39.11	21.89	32.53	6.47	*61.00*	*39.00*
Mexico	10.98	2.04	80.85	6.13	*13.02*	*86.98*
Peru	17.82	27.11	21.02	34.05	*44.93*	*55.07*

Note. Percentage contribution of each structural shock in explaining the historical variance of the real exchange rate quarterly changes. ϕ, υ, φ, η indicate relative productivity, relative labour, relative fiscal and relative preference shocks, respectively. The column "Supply" is the aggregate contribution of ϕ and υ disturbances. The column "Demand" is the aggregate contribution of the φ and η disturbances.

Table 6 - Historical decomposition: recursive method

		Argentina		
	ϕ	υ	φ	η
Mean	19.07	22.24	38.93	19.76
Std. err. of mean	0.05	0.22	0.33	0.14
Minimum	18.23	19.65	34.92	17.80
Maximum	20.08	25.16	42.27	21.71
		Bolivia		
	ϕ	υ	φ	η
Mean	2.59	2.53	79.07	15.81
Std. err. of mean	0.11	0.03	0.36	0.24
Minimum	1.98	2.08	73.36	12.33
Maximum	5.21	2.93	83.57	18.88
		Brazil		
	ϕ	υ	φ	η
Mean	17.65	38.03	1.09	43.22
Std. err. of mean	0.16	0.17	0.01	0.17
Minimum	15.44	35.14	0.84	40.22
Maximum	20.18	41.18	1.25	46.73
		Chile		
	ϕ	υ	φ	η
Mean	39.98	22.03	31.05	6.94
Std. err. of mean	0.14	0.10	0.13	0.06
Minimum	38.04	20.77	28.73	6.10
Maximum	42.51	23.70	32.87	7.65
		Mexico		
	ϕ	υ	φ	η
Mean	10.32	1.87	82.70	5.11
Std. err. of mean	0.03	0.02	0.12	0.08
Minimum	9.87	1.66	80.85	3.93
Maximum	10.98	2.17	84.53	6.13
		Peru		
	ϕ	υ	φ	η
Mean	22.29	30.68	14.37	32.66
Std. err. of mean	0.29	0.31	0.45	0.17
Minimum	17.82	26.24	11.20	30.49
Maximum	24.31	33.18	21.31	35.56

Note. Percentage contribution of each structural shock in explaining the historical variance of the real exchange rate quarterly changes. ϕ, υ, φ, η indicate relative productivity, relative labour, relative fiscal and relative preference shocks, respectively. Summary statistics computed over simulation windows of increasing size, extended by a datapoint at a time, are reported by rows. All windows start with the first available observation, but they have different ending quarters. The smallest window covers the period up to 1994*q*4, while the largest window embraces the entire sample span.

Table 7 – Robustness analysis: model specification

	Model specification		
	FD	*HP*	*LD*
Argentina	[0.01]	[0.24]	[0.00]
Bolivia	[0.30]	[0.00]	[0.00]
Brazil	[0.00]	[0.00]	[0.00]
Chile	[0.09]	[0.54]	[0.45]
Mexico	[0.10]	[0.41]	[0.00]
Peru	[0.00]	[0.65]	[0.00]

Note. p-values from a χ^2-distributed LR over-identifying test with one degree of freedom are reported in squared brackets. FD, HP and LD indicates first differences, HP (Hodrick and Prescott, 1997) and linear detrending filters, respectively.

Table 8 – Robustness analysis: forecast error variance decompositions

	Model specification					
	FD		*HP*		*LD*	
	Demand shocks	φ	*Demand shocks*	φ	*Demand shocks*	φ
Argentina	*79.31*	2.20	*83.02*	6.62	*43.17*	14.22
Bolivia	*94.48*	21.24	*82.46*	12.26	*89.05*	7.64
Brazil	*85.26*	30.37	*42.64*	14.89	*60.52*	25.66
Chile	*79.85*	5.13	*82.16*	30.23	*45.01*	3.05
Mexico	*91.51*	7.61	*63.33*	4.18	*17.29*	2.06
Peru	*89.23*	5.52	*45.60*	0.40	*27.78*	2.06

Note. Average percentage contribution of demand and relative fiscal shocks (φ) in explaining real exchange rate fluctuations at different cyclical frequencies over a simulation horizon of 20 quarters. FD, HP and LD indicates first differences, HP (Hodrick and Prescott, 1997) and linear detrending filters, respectively.

Table 9 – Augmented model: testing for the over-identifying restriction

	Model specification			
	VEC model	VAR mode		
		FD	*HP*	*LD*
Argentina	[0.42]	[0.02]	[0.79]	[0.00]
Bolivia	[0.27]	[0.00]	[0.01]	[0.00]
Brazil	[0.00]	[0.00]	[0.00]	[0.00]
Chile	[0.07]	[0.00]	[0.29]	[0.00]
Mexico	[0.13]	[0.03]	[0.25]	[0.00]
Peru	[0.00]	[0.00]	[0.00]	[0.00]

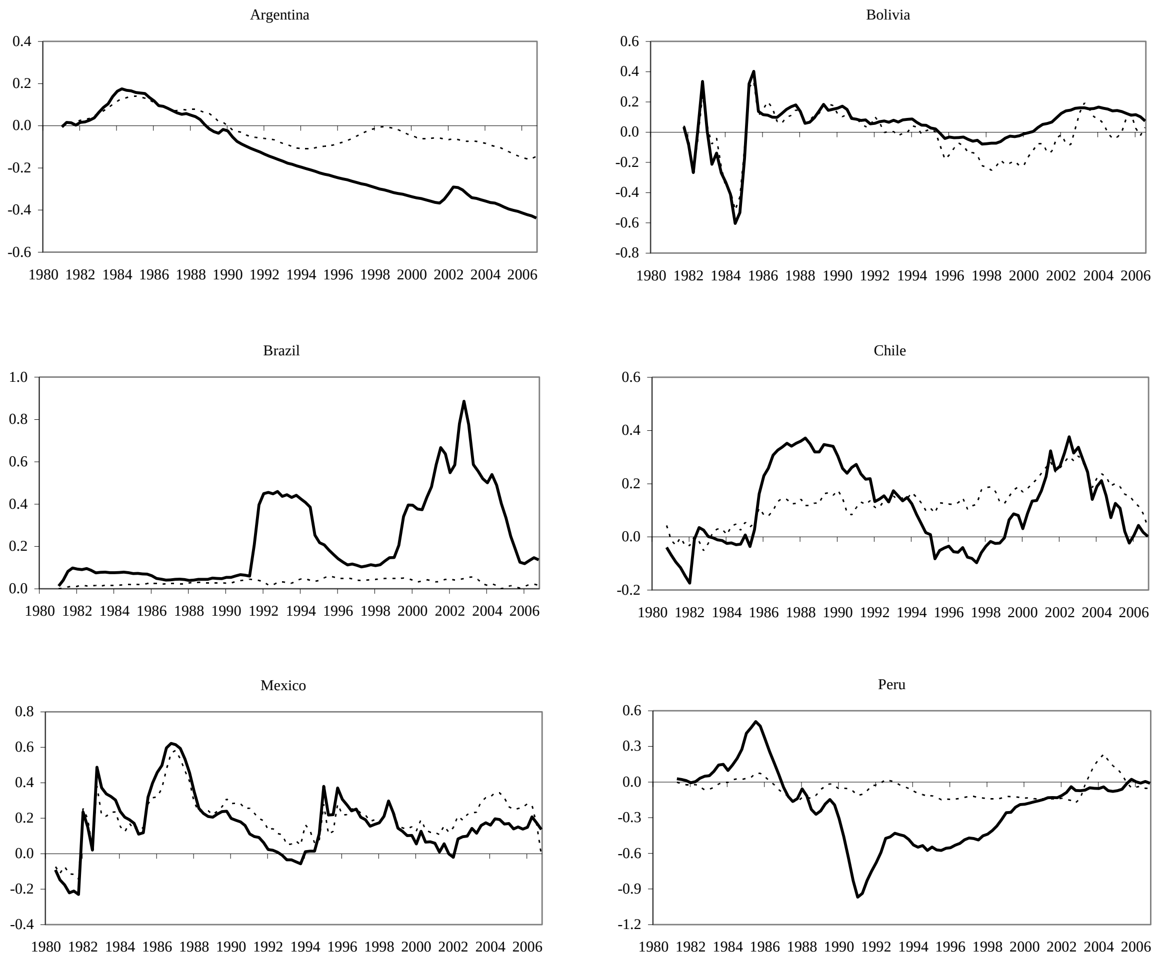
Note. p-values from a χ^2-distributed LR over-identifying test with one degree of freedom are reported in squared brackets. FD, HP and LD indicates first differences, HP (Hodrick and Prescott, 1997) and linear detrending filters, respectively. VEC model stands for baseline specification with cointegration.

Table 8 – Augmented model: forecast error variance decompositions

	Model specification							
	VEC model		VAR model					
			FD		*HP*		*LD*	
	Real shocks	φ	*Real shocks*	φ	*Real shocks*	φ	*Real shocks*	φ
Argentina	*93.02*	66.24	*97.94*	2.12	*94.79*	6.43	*99.28*	18.29
Bolivia	*98.86*	11.46	*72.31*	6.56	*88.56*	6.19	*83.48*	8.87
Brazil	*99.65*	2.46	*94.05*	39.18	*90.19*	10.72	*97.16*	37.54
Chile	*83.85*	9.98	*98.29*	1.75	*98.68*	29.32	*98.90*	8.77
Mexico	*99.52*	42.06	*98.96*	5.29	*98.00*	4.49	*99.79*	3.40
Peru	*94.21*	20.13	*80.95*	5.20	*86.87*	1.41	*87.15*	11.58

Note. Average percentage contribution of non-monetary (real) and relative fiscal shocks (φ) in explaining real exchange rate fluctuations at different cyclical frequencies over a simulation horizon of 20 quarters. FD, HP and LD indicates first differences, HP (Hodrick and Prescott, 1997) and linear detrending filters, respectively. VEC model stands for baseline specification with cointegration.

Figure 1 - Real exchange rate dynamics and the component driven by the fiscal shock



Note. In each graph, the solid line indicates the real exchange rate, while the dashed line plots its component driven by the fiscal shock.