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An econometric contribution to the intertemporal approach of the current account

Wagner Piazza Gaglianone*
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Abstract

This paper investigates an intertemporal optimization model to analyze the current account through Campbell & Shiller's (1987) approach. In this setup, a Wald test is conducted to analyze a set of restrictions imposed to a VAR, used to forecast the current account for a set of countries. We focused on three estimation procedures: OLS, SUR and the two-way error decomposition of Fuller & Battese (1974). We also propose an original note on Granger causality, which is a necessary condition to perform the Wald test. Theoretical results show that, in the presence of global shocks, OLS and SUR estimators might lead to a biased covariance matrix, with serious implications to the validation of the model. A Monte Carlo simulation confirms these findings and indicates the Fuller & Battese procedure in the presence of global shocks. An empirical exercise for the G-7 countries is also provided, and the results of the Wald test substantially change with different estimation techniques. In addition, global shocks can account up to 40% of the total residuals of the G-7. The model is not rejected for Canada, in sharp contrast to the literature, since the previous results might be seriously biased, due to the existence of global shocks.

Keywords: current account, capital mobility, error decomposition, common shocks.
JEL Classification: C31, E21, F32, F47.

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1 Introduction

The current account can be used by domestic residents to smooth consumption by borrowing from or lending to the rest of the world. Several authors have analyzed the open economy model, initially proposed by Sachs (1982) and later detailed by Obstfeld & Rogoff (1994), with a theoretical framework that defines the optimal current account from the agents' intertemporal optimization problem, supposing that agents can freely smooth consumption in the presence of shocks. The comparison of this optimal value with the observed current account allows us to test for consumption optimality.

This approach is encompassed by several classes of small open economy models,[1] and the most basic version is the present value model (PVM) of the current account. Although the literature of PVMs is relatively extensive, the following papers should be mentioned (suggesting an overall rejection of the model for developed countries): Sheffrin & Woo (1990) perform a study of the current account of Belgium, Canada, Denmark and UK. The results indicate a rejection of the model for Denmark, Canada and UK, whereas the PVM could not be rejected for Belgium. Otto (1992) tests the PVM for the USA and Canada, and rejects the model in both countries. Ghosh (1995) investigates the current account of 5 major industrialized countries: USA, Canada, Japan, Germany and UK, and the results suggest rejection of the model in all countries, except for the USA.

On the other hand, some papers document results supporting the PVM, in contrast to the previous findings, such as Ghosh & Ostry (1995) that test it for 45 developing countries and do not reject it for about 2/3 of the countries. Hussein & Mello (1999) also test the PVM for some developing countries (Chile, Greece, Ireland, Israel, Malaysia, Mexico, South Africa, South Korea and Venezuela), and find evidences to support the PVM. In the same line, Agénor et al. (1999) focus on the current account of France, concluding that the PVM holds and the analyzed country was perfectly able to smooth consumption.[2]

Notwithstanding the lack of consensus on the macroeconomic front, what happens on the "econometric side"? Is it possible that an inappropriate econometric technique leads to wrong conclusions regarding the rejection of the PVM? Unlike the mentioned literature, the objective of this paper is to provide an econometric approach to the current account debate. The methodology generally adopted in the literature to analyze the PVM was initially proposed by Campbell & Shiller (1987), and consists of estimating an unrestricted VAR,

[1] For recent developments regarding small open economy models see Grohé & Uribe (2003). In addition, see Chinn & Prasad (2003), which provide an empirical characterization of the determinants of current account for a large sample of industrial and developing countries. See also Aguiar & Gopinath (2006), which develop a quantitative model of debt and default in a small open economy. Finally, see Obstfeld & Rogoff (1996) and Bergin (2003) for a good discussion about new open economy literature and its empirical dimension.

[2] In order to deepen the debate, several authors also proposed extensions to the standard PVM model. A short list includes Ghosh & Ostry (1997), which consider precautionary saving, Gruber (2000) includes habit formation, Bergin & Sheffrin (2000) allow for a time-varying world interest rate and consider tradable and non-tradable goods, İşcan (2002) modifies the basic model introducing durables and also nontraded goods. More recently, Nason & Rogers (2006) propose a real business cycle (RBC) model, which nests the basic PVM, including non-separable preferences, shocks to fiscal policy and world interest rate, and imperfect capital mobility, explanations broadly presented in the literature for the rejection of the PVM. According to Nason & Rogers (2006), although each suspect matters in some way, none is capable to completely improve the fit of the model to the data.

whose parameters are used in the construction of the optimal current account, and perform a Wald test to investigate a set of restrictions imposed to the VAR, testing whether the optimal current account equals the observed series.

However, the presence of common shocks in the econometric model can play a crucial role, and is widely recommended in the literature to explain business cycles fluctuations. For instance, Centoni et al. (2003) investigate whether co-movements observed in the international business cycles are the consequences of common shocks or common transmission mechanisms. Similarly to most studies (such as King et al. 1991), Centoni et al. (2003) confirm that permanent shocks are the main source of the business cycles, accounting for a 50% effect in a panel of European countries. The authors also show that the domestic component is responsible for most of the business cycle effects of transitory shocks for all the G-7 countries, whereas the foreign component dominates the cyclical variability that is due to permanent shocks in France, Germany and Italy.[3]

This way, seems to exist a consensus in the literature regarding a common world component that might partially explain current account fluctuations. This common (or global) shock is ignored in the OLS estimation (widely used in the literature), but could be considered in a SUR approach. In fact, along this paper we stress the fact that in the estimation process an econometrician might consider a set of countries separately (OLS) as well as jointly (e.g., SUR), in order to capture contemporaneous correlations of the residuals of the VAR. However, due to the possible finite sample bias of the OLS and SUR covariance matrices (see Driscoll & Kraay, 1998), we also investigate the two-way error decomposition of Fuller & Battese (1974), hereafter FB, which can properly treat the existence of common shocks in the estimation process.

Therefore, we aim to contribute to the current account debate by investigating the estimation of a PVM through three different techniques (OLS, SUR and FB). In addition, we propose a quite original note on Granger causality, which is showed to be a necessary condition to perform the Wald test of Campbell & Shiller (1987). In addition, we present some theoretical results to show that (in the presence of common shocks) OLS and SUR estimators might produce a biased covariance matrix, with serious implications to the validation of the model.

A small Monte Carlo simulation confirms these findings and indicates the FB procedure in the presence of global shocks. We also provide an empirical exercise for the G-7 countries, and (indeed) the results substantially change with different estimation techniques. In addition, global shocks can account up to 40% of the total residuals of the G-7, confirming the importance of such shocks in the estimation process. The model is not rejected for Canada, in sharp contrast to the literature, since the previous results might be seriously biased, due to the existence of global shocks.

[3] In the same sense, Canova & Dellas (1993) document that after 1973 the presence of common disturbances, such as the first oil shock, plays a role in accounting for international output co-movements. Glick & Rogoff (1995) study the current account response to different productivity shocks in the G-7 countries, based on a structural model including global and country-specific shocks. Furthermore, Canova & Marrinan (1998), which investigate the generation and transmission of international cycles in a multicountry model with production and consumption interdependencies, argue that a common component to the shocks and of production interdependencies appear to be crucial in matching the data.

This paper is structured in the following way: Section 2 provides an overview of the macroeconomic model of the current account and discusses some econometric techniques that might be used in the estimation process. Section 3 presents the results of an empirical exercise for the G-7 countries, and Section 4 presents our main conclusions.

2 Methodology

2.1 Present Value Model

The Present Value Model (PVM) adopted to analyze the intertemporal optimization problem of a representative agent is based on Sachs (1982), considering the perfect capital mobility hypothesis across countries. In this context, countries save through flows of capital in their current accounts, according to their expectations of future changes in net output. Thus, the current account is used as an instrument of consumption smoothing against possible shocks to the economy, and can be expressed by

$$CA_t = B_{t+1} - B_t = Y_t + rB_t - I_t - G_t - C_t \tag{1}$$

where B_t represents foreign assets, Y_t gross domestic product (GDP), r the world interest rate, I_t total investment, G_t the government's expenses and C_t aggregated consumption.

The consumption path, related to the dynamics of the current account, can be divided into two components: the trend term, generated by the difference between the world interest rate and the rate of time preference, and the smoothing component, related to the expectations of changes in permanent income. This paper only studies the second component effect, by isolating from the current account, the trend component in consumption. Thus, the optimal current account (only associated with the consumption smoothing term) is given by

$$CA_t^* = Y_t + rB_t - I_t - G_t - \theta C_t \tag{2}$$

where θ is a parameter that removes the trend component in consumption.[4] The net output Z_t , also known in the literature as national cash flow, is defined by

$$Z_t \equiv Y_t - I_t - G_t \tag{3}$$

Substituting the optimal consumption expression in equation (2), it can be shown that the present value relationship between the current account and the future changes in net output is given by (see Ghosh & Ostry (1995) for further details):

$$CA_t^* = -\sum_{j=1}^{\infty} (\frac{1}{1+r})^j E_t(\Delta Z_{t+j} \mid R_t) \tag{4}$$

where R_t is the agent's information set. It should be mentioned that the main assumptions of the model are time-separable preferences, zero depreciation of capital, and complete asset markets. A quadratic form is also adopted for the utility function, without precautionary saving effects (see Ghosh & Ostry, 1997).

[4] The tilt parameter (θ) is not equal to one whenever the rate of time preference differs from the world interest rate.

According to equation (4), the optimal current account is equal to minus the present value of the expected changes in net output. For instance, the representative agent will increase its current account, accumulating foreign assets, if a future decrease in income is expected, and vice-versa.

2.2 Econometric Model

The econometric model is based on the methodology developed by Campbell & Shiller (1987), which suggest an alternative way to verify a PVM when the involved variables are stationary. The idea is to test a set of restrictions imposed to a Vector Auto Regression (VAR), used to forecast the current account through equation (4). The advantage of this approach is that, although the econometrician does not observe the agent's information set, this framework allows us to summarize all the relevant information through the variables used in the construction of the VAR.

However, to apply this methodology, the VAR must be stationary. Hence, the first empirical implication is to verify whether ΔZ_t is a weakly stationary variable. The current account (in level) must also be a stationary variable, since it can be written as a lineal combination of stationary variables (via equation (4)). The stationarity of these variables can be checked later by unit root tests. Campbell & Shiller (1987) argue that series represented by a Vector Error Correction Model (VECM) can be rewritten as an unrestricted VAR. Thus, consider the following VAR representation:[5]

$$\begin{bmatrix} \Delta Z_t^i \\ CA_t^i \end{bmatrix} = \begin{bmatrix} a^i(L) & b^i(L) \\ c^i(L) & d^i(L) \end{bmatrix} \times \begin{bmatrix} \Delta Z_{t-1}^i \\ CA_{t-1}^i \end{bmatrix} + \begin{bmatrix} \mu_1^{i*} \\ \mu_2^{i*} \end{bmatrix} + \begin{bmatrix} \varepsilon_{1t}^i \\ \varepsilon_{2t}^i \end{bmatrix} \tag{5}$$

where the index i represents the analyzed country and $a^i(L)$, $b^i(L)$, $c^i(L)$ and $d^i(L)$ are polynomials of order p. Hence, the estimation of the VAR must be preceded by the estimation of θ, which occurs in the cointegration analysis between C_t and $(Y_t + rB_t - I_t - G_t)$. The model VAR(p) can be described as a VAR(1), in the following way:

$$\begin{bmatrix} \Delta Z_t^i \\ \vdots \\ \vdots \\ \Delta Z_{t-p+1}^i \\ CA_t^i \\ \vdots \\ \vdots \\ CA_{t-p+1}^i \end{bmatrix} = \begin{bmatrix} a_1^i & \cdots & a_p^i & b_1^i & \cdots & b_p^i \\ 1 & & & & & \\ & \ddots & & & 0 & \\ & & 1 & & & \\ c_1^i & \cdots & c_p^i & d_1^i & \cdots & d_p^i \\ & & & 1 & & \\ & 0 & & & \ddots & \\ & & & & & 1 \end{bmatrix} \begin{bmatrix} \Delta Z_{t-1}^i \\ \vdots \\ \vdots \\ \Delta Z_{t-p}^i \\ CA_{t-1}^i \\ \vdots \\ \vdots \\ CA_{t-p}^i \end{bmatrix} + \begin{bmatrix} \mu_1^{i*} \\ 0 \\ \vdots \\ 0 \\ \mu_2^{i*} \\ 0 \\ \vdots \\ 0 \end{bmatrix} + \begin{bmatrix} \varepsilon_{1t}^i \\ 0 \\ \vdots \\ 0 \\ \varepsilon_{2t}^i \\ 0 \\ \vdots \\ 0 \end{bmatrix} \tag{6}$$

or, in a compact form:

$$X_t = AX_{t-1} + \mu^* + \varepsilon_t \tag{7}$$

[5] It should be mentioned that, hereafter, CA_t will be constructed considering the parameter θ, to remove the trend component in consumption, as it follows: $CA_t = Y_t + rB_t - I_t - G_t - \theta C_t$.

where $X_t \equiv \begin{bmatrix} \Delta Z_t^i & \cdots & \Delta Z_{t-p+1}^i & CA_t^i & \cdots & CA_{t-p+1}^i \end{bmatrix}'$, A is the companion matrix, μ^* represents a vector of intercepts, and ε_t is a vector that contains the residuals. The VAR(1) is stationary by assumption, and the equation (7) can be rewritten removing the vector of means μ:

$$(X_t - \mu) = A(X_{t-1} - \mu) + \varepsilon_t \tag{8}$$

where $\mu^* = (I - A)\mu$. The forecast of the model j periods ahead is given by

$$E[(X_{t+j} - \mu \mid H_t) = A^j (X_t - \mu) \tag{9}$$

where H_t is the econometrician's information set (composed of current and past values of CA and ΔZ), contained in the agent's information set R_t. Define h' as a vector with $2p$ null elements, except the first: $h' = \begin{bmatrix} 1 & 0 & \ldots & 0 \end{bmatrix}$. Then, one can select ΔZ_t in the vector X_t, in the following way:

$$\Delta Z_t = h' X_t \therefore \Delta Z_{t+j} = h' X_{t+j} \therefore (\Delta Z_{t+j} - \mu_{\Delta Z}) = h'(X_{t+j} - \mu) \tag{10}$$

where the vector μ contains the means $\mu_{\Delta Z}$ and μ_{CA^*}. Thus, applying the conditional expectation in the previous expression, it follows that:

$$E[(\Delta Z_{t+j} - \mu_{\Delta Z}) \mid H_t] = E[h'(X_{t+j} - \mu) \mid H_t] = h' E[(X_{t+j} - \mu) \mid H_t] = h' A^j (X_t - \mu) \tag{11}$$

where the last equality comes from equation (9). In order to calculate the optimal current account CA_t^*, one can take expectations of equation (4):

$$E(CA_t^* \mid H_t) = CA_t^* = -\sum_{j=1}^{\infty} (\frac{1}{1+r})^j E(\Delta Z_{t+j} \mid H_t) \tag{12}$$

The first equality comes from the fact that CA_t^* is contained in H_t, and the second is given by the law of iterated expectations ($H_t \subseteq R_t$). Applying the unconditional expectation in the previous expression:

$$E(CA_t^*) = -\sum_{j=1}^{\infty} (\frac{1}{1+r})^j E(\Delta Z_{t+j}) \therefore \mu_{CA^*} = -\sum_{j=1}^{\infty} (\frac{1}{1+r})^j \mu_{\Delta Z} \tag{13}$$

Combining equation (12) with equation (13), it follows that:

$$(CA_t^* - \mu_{CA^*}) = -\sum_{j=1}^{\infty} (\frac{1}{1+r})^j E(\Delta Z_{t+j} - \mu_{\Delta Z} \mid H_t) \tag{14}$$

Applying the expression (11) in the equation above:

$$(CA_t^* - \mu_{CA^*}) = -\sum_{j=1}^{\infty} (\frac{1}{1+r})^j h' A^j (X_t - \mu) = -h'(\frac{A}{1+r})(I - \frac{A}{1+r})^{-1} (X_t - \mu) \tag{15}$$

where the last equality is due to the convergence of an infinite sum, since the variables ΔZ_t and CA_t are stationary. Rewriting the previous equation in a simplified form:

$$(CA_t^* - \mu_{CA^*}) = K(X_t - \mu) \tag{16}$$

$$K = -h'(\frac{A}{1+r})(I - \frac{A}{1+r})^{-1} \tag{17}$$

where the vector K is derived from the world interest rate r and the matrix A. To formally test the model, one can analyze the null hypothesis $(CA_t^* - \mu_{CA^*}) = (CA_t - \mu_{CA})$. Define g' as a vector with $2p$ null elements, except the $(p+1)th$ element, that assumes a unit value. Thus, under the null hypothesis, it follows that:

$$(CA_t^* - \mu_{CA^*}) = (CA_t - \mu_{CA}) = g'(X_t - \mu) \tag{18}$$

Combining equations (16) and (18), the model can be formally tested through a set of restrictions imposed to the coefficients of the VAR:

$$g'(X_t - \mu) = -h'(\frac{A}{1+r})(I - \frac{A}{1+r})^{-1}(X_t - \mu) \therefore g'(I - \frac{A}{1+r}) = -h'(\frac{A}{1+r}) \tag{19}$$

Applying the structure of matrix A into equation (19), the following restrictions[6] can be derived:

$$\begin{aligned} a_i &= c_i \ \ ; i = 1...p \\ b_i &= d_i \ \ ; i = 2...p \\ b_1 &= d_1 - (1+r) \end{aligned} \tag{20}$$

Another important implication of the model is that the current account Granger-cause changes in net output, or in other words, CA_t helps to forecast ΔZ_t. This causality can be tested by means of the statistical significance of the $b(L)$ coefficients. Therefore, the implications of the intertemporal optimization model, according to Otto (1992), can be summarized by:[7]

1. Verifying the stationarity of CA_t and ΔZ_t, through unit root tests;
2. Checking if CA_t Granger-cause ΔZ_t;
3. Analyzing the cointegration between C_t and $(Y_t + rB_t - I_t - G_t)$, and calculating the parameter θ;
4. Formally investigating, by means of a Wald test, the equality of the optimal and observed current accounts, given by restrictions (20).

2.3 A note on Granger Causality and Wald Tests

The optimal current account is generated from the vector K (see expressions (16) and (17)), which depends on matrix A and the world interest rate r. However, it should be noted that an estimated coefficient for matrix A could not be statistically significant. These results could seriously compromise the subsequent optimal current account analysis, as it follows.

The Granger causality between the current account and net output (CA_t Granger-cause ΔZ_t) is a primordial implication of the theoretical model, and as argued before, can be alternatively tested through the significance of the $b(L)$ coefficients.[8] Moreover, if this implication is not empirically observed, the model

[6] These restrictions can be verified by a Wald test.

[7] It is in fact a set of testable implications of the PVM. Therefore, the statistical acceptance of the model occurs only if all of these implications could be verified.

[8] Presented in equation (5).

should be rejected irrespective of any other results, since equation (4) is the theoretical foundation of the whole study. In this case, the current account could not help to predict variations in net output, suggesting that the agents are badly described by the model. Thus, one should not construct the optimal current account and perform a comparison with the observed series. Unfortunately, this is done in several papers presented in the literature.

To study this topic more carefully, a simple VAR(1) is initially presented. The Granger causality between CA_t and ΔZ_t, in this case, will be determined by the statistical significance of the b_1 coefficient.

$$\begin{bmatrix} \Delta Z_t \\ CA_t \end{bmatrix} = \begin{bmatrix} a_1 & b_1 \\ c_1 & d_1 \end{bmatrix} \times \begin{bmatrix} \Delta Z_{t-1} \\ CA_{t-1} \end{bmatrix} + \begin{bmatrix} \mu_1 \\ \mu_2 \end{bmatrix} + \begin{bmatrix} \varepsilon_{1t} \\ \varepsilon_{2t} \end{bmatrix}. \tag{21}$$

In this case, the VAR is represented in a compact form by $X_t = AX_{t-1} + \mu + \varepsilon_t$ and, after some algebraic manipulations, the vector K takes the form:

$$K = -h'(\frac{A}{1+r})(I_2 - \frac{A}{1+r})^{-1} = \begin{bmatrix} \alpha & \beta \end{bmatrix}, \tag{22}$$

where

$$\alpha = \frac{-a_1(1+r-d_1) - b_1c_1}{(1+2r-d_1+r^2-rd_1-a_1-a_1r+a_1d_1-b_1c_1)}, \tag{23}$$

$$\beta = \frac{-a_1b_1 - b_1(1+r-a_1)}{(1+2r-d_1+r^2-rd_1-a_1-a_1r+a_1d_1-b_1c_1)}. \tag{24}$$

If the Granger causality is rejected by the data (e.g., b_1 is not significant), then equation (25) indicates that $\beta = 0$, or in other words, CA_t^* is not a function of CA_t. In this case, the optimal current account would be given by

$$(CA_t^* - \mu_{CA^*}) = K(X_t - \mu) = \begin{bmatrix} \alpha & 0 \end{bmatrix} \begin{bmatrix} \Delta Z_t - \mu_{\Delta Z} \\ CA_t - \mu_{CA} \end{bmatrix} = \alpha\,(\Delta Z_t - \mu_{\Delta Z}). \tag{25}$$

Hence, if $\beta = 0$ the null hypothesis $(CA_t^* - \mu_{CA}^*) = (CA_t - \mu_{CA})$ is always rejected, since under Ho β should be equal to one (and α should be zero). A further analysis of the vector K for a VAR(2) is presented in appendix, in a similar way. The generalization of this cautionary note for a VAR(p) is straightforward, and can be summarized by Proposition 1. According to Hamilton (1994), in the context of a bivariate VAR(p), if one of the two variables does not Granger-cause the other, then the companion matrix is lower triangular (e.g., $b(L) = 0$). Thus, the β_i coefficients of the vector K $(i = 1, \ldots, p)$ are always zero, because of the algebraic structure of the vector, as also detailed in appendix.

Proposition 1 *Consider the VAR representation (5) of the intertemporal model of current account. The Granger causality from the current account (CA_t) to the first difference of the net output (ΔZ_t) is a necessary condition to perform the Wald test and verify the validation of the model, i.e., if the $b(L)$ coefficients of the VAR(p) model are not statistically significant, then, the Wald test is not applicable and the model should be rejected.*

Proof. See Appendix. ■

Therefore, if the Granger causality could not be confirmed by the data set, neither a Wald test should be performed nor the optimal current account should be generated, since the basic assumption of the model is not verified,[9] as summarized in table 1.

Table 1 - A note on Granger causality and Wald tests

Result of Granger causality	Wald test	Model	Conclusion
CA_t not Granger-cause ΔZ_t	not applicable (β=0)	rejected	model cannot generate CA_t^*(*)
CA_t Granger-cause ΔZ_t	rejects Ho ($\beta \neq 1$)	rejected	$CA_t^* \neq CA_t$ (**)
	does not reject Ho (β=1)	not rejected	$CA_t^* = CA_t$ (***)

Notes: (*) indicates that CA_t^* only depends on ΔZ, instead of CA_t

(**) suggests that agents do not smooth consumption;

(***) means that agents perfectly smooth consumption.

2.4 Estimation Method

2.4.1 SUR estimation

The VAR model (5) is usually estimated in the literature, equation-by-equation, using OLS. However, the Seemingly Unrelated Regressions (SUR) technique, originally developed by Zellner (1962), can also be adopted, since it is based on a Generalized Least Squares (GLS) estimation applied to a system of equations as a whole, in which the data for several countries is examined simultaneously. The joint estimation is given by stacking the system of equations that compose the VAR (for each country $i = 1, \ldots, N$) in the following way:

$$\begin{bmatrix} \Delta Z_t^i \\ CA_t^i \end{bmatrix} = \begin{bmatrix} a^i(L) & b^i(L) \\ c^i(L) & d^i(L) \end{bmatrix} \times \begin{bmatrix} \Delta Z_{t-1}^i \\ CA_{t-1}^i \end{bmatrix} + \begin{bmatrix} \varepsilon_{1t}^i \\ \varepsilon_{2t}^i \end{bmatrix} \tag{26}$$

Then, define $Y_t^i = \begin{pmatrix} \Delta Z_t^i \\ CA_t^i \end{pmatrix}_{(2\times 1)}$, $X_t^i = \begin{pmatrix} \Delta Z_{t-1}^i \\ \vdots \\ \Delta Z_{t-p_i}^i \\ CA_{t-1}^i \\ \vdots \\ CA_{t-p_i}^i \end{pmatrix}_{(2p_i \times 1)}$, $\varepsilon_t^i = \begin{pmatrix} \varepsilon_{1t}^i \\ \varepsilon_{2t}^i \end{pmatrix}_{(2\times 1)}$

and $\beta_i' = \begin{pmatrix} a_1^i & \ldots & a_{p_i}^i & b_1^i & \ldots & b_{p_i}^i & c_1^i & \ldots & c_{p_i}^i & d_1^i & \ldots & d_{p_i}^i \end{pmatrix}_{(1\times 4p_i)}$

[9] Recall Ghosh & Ostry (1995) results, in which the authors test the PVM for 45 developing countries and do not reject it for 29 countries. However, a careful analysis of the tests reveals that only 25 countries (from the entire set of countries) in fact support the Granger causality implication (at 5% level). This way, the paper should conclude that (at most) in only 18 countries (instead of 29) the model could not be rejected, since only 18 countries indeed exhibit good results for both the Wald and Granger causality tests.

This way, the VAR for a country i can be represented by

$$\left(Y_t^i\right)_{(2 \text{ x } 1)} = \begin{pmatrix} X_t^{1\prime} & 0 \\ 0 & X_t^{1\prime} \end{pmatrix}_{(2 \text{ x } 4\text{p}_i)} \begin{pmatrix} \beta_i \end{pmatrix}_{(4\text{p}_i \text{ x } 1)} + \begin{pmatrix} \varepsilon_t^i \end{pmatrix}_{(2 \text{ x } 1)} \tag{27}$$

Furthermore, the system of equations for a given set of N countries can be expressed by

$$\begin{pmatrix} Y_t^1 \\ \vdots \\ Y_t^N \end{pmatrix}_{(2\text{N x } 1)} = \begin{pmatrix} X_t^{1\prime} & 0 & \dots & & 0 \\ 0 & X_t^{1\prime} & & & \\ \vdots & & \ddots & 0 & \vdots \\ & & 0 & X_t^{N\prime} & 0 \\ 0 & & \dots & 0 & X_t^{N\prime} \end{pmatrix}_{(2\text{N x 4P})} \begin{pmatrix} \beta_1 \\ \vdots \\ \beta_N \end{pmatrix}_{(4\text{P x } 1)} + \begin{pmatrix} \varepsilon_t^1 \\ \vdots \\ \varepsilon_t^N \end{pmatrix}_{(2\text{N x } 1)} \tag{28}$$

or in a compact form $Y = X\beta + \varepsilon$. The name SUR comes from the fact that each equation in the previous system has its own vector of coefficients, which might suggest that the equations are unrelated. Nevertheless, correlation across the errors in different equations can provide links that can be exploited in estimation. It should be noted that $\sum_{i=1}^{N} p_i = P$, where p_i is the number of lags of the VAR, for a country i. The residuals ε have mean zero and are serially uncorrelated, with covariance matrix given by $E(\varepsilon\varepsilon') = \sigma^2\Omega$. Hence, the GLS estimator of β and its variance-covariance matrix are given by

$$\widetilde{\beta} = \left(X'\Omega^{-1}X\right)^{-1} X'\Omega^{-1}Y \tag{29}$$

$$E(\widetilde{\beta} - \beta)(\widetilde{\beta} - \beta)' = \sigma^2 \left(X'\Omega^{-1}X\right)^{-1} \tag{30}$$

In general, the $(N \times N)$ matrix Ω is unknown and the last expression cannot be directly applied. However, $\widetilde{\beta}$ can be calculated by an estimate of the ijth element of Ω, given by

$$\widehat{w}_{ij} = \frac{e_i' e_j}{T}, \qquad \text{where } i, j = 1, \dots, N \tag{31}$$

where e_i is a $(T \times 1)$ vector containing the residuals of the ith equation estimated by OLS. In this case, a feasible SUR estimator of β is obtained as

$$\widetilde{\beta}^* = \left(X'\widehat{\Omega}^{-1}X\right)^{-1} X'\widehat{\Omega}^{-1}Y \tag{32}$$

$$E(\widetilde{\beta}^* - \beta)(\widetilde{\beta}^* - \beta)' = \sigma^2 \left(X'\widehat{\Omega}^{-1}X\right)^{-1} \tag{33}$$

The OLS estimator, on the other hand, is given by

$$\widehat{\beta} = \left(X'X\right)^{-1} X'Y \tag{34}$$

$$E(\widehat{\beta} - \beta)(\widehat{\beta} - \beta)' = \sigma^2 \left(X'X\right)^{-1} X'\Omega X \left(X'X\right)^{-1} \tag{35}$$

The difference between their variance-covariance matrices is a positive semidefinite matrix, and can be expressed by

$$E(\widehat{\beta} - \beta)(\widehat{\beta} - \beta)' - E(\widetilde{\beta} - \beta)(\widetilde{\beta} - \beta)' = \sigma^2 \eta\Omega\eta' \tag{36}$$

where $\eta = \left(X'X\right)^{-1} X' - \left(X'\Omega^{-1}X\right)^{-1} X'\Omega^{-1}$, indicating the gain in efficiency of SUR estimators in comparison to the OLS counterpart.

2.4.2 Caveats of SUR estimation

The SUR estimator generally exhibits a good performance when N is small relative to T, but in fact becomes not feasible when $T < (N+1)/2$. Even when the SUR model is correctly specified, its performance might be poor due to a very large number of free parameters to be estimated, in comparison to the time dimension. In other words, as N becomes large for a fixed value of T, the estimated covariance matrix becomes "nearly" singular, introducing a bias into the standard error estimates. This way, the finite sample performance of the OLS and SUR estimators deteriorates rapidly as the size of the cross-sectional dimension increases.

Driscoll & Kraay (1998) investigate finite-sample properties of variance estimators, concluding that both OLS and SUR estimators indeed exhibit substantial downward finite sample bias, even for moderate values of cross-sectional dependence, and are outperformed by a spatial correlation consistent estimator proposed in their article, based on the nonparametric technique of Newey & West (1987) and Andrews (1991).

The main idea is to obtain consistent estimates of the $N \times N$ matrix of cross-sectional correlations by averaging over the time dimension. This way, the estimated cross-sectional covariance matrix can be used to construct standard errors, which are robust to the presence of spatial correlation. Driscoll & Kraay (1998)'s approach, in contrast to SUR, might be applicable in situations such as cross-country panel data models with a relatively large number of countries.

In this paper, however, we focus on a panel model with small N and large T, but in order to deal with possible finite sample bias of the covariance matrix, we also investigate a two-way error decomposition (next described) that can properly deal with cross-country correlations.

2.4.3 Fuller & Battese (1974) and the two-way error decomposition

The performance of any estimation procedure depends on the statistical characteristics of the error components in the model. In this section, we adopt the Fuller & Battese (1974) method to consider individual and time-specific random effects into the error disturbances, in which parameters can efficiently be estimated by using a feasible GLS framework.

In dynamic panel models, the presence of lagged dependent variables might lead to a non-zero correlation between regressors and error term. This could render OLS estimator for a dynamic error-component model to be biased and inconsistent (see Baltagi, 2001, p. 130), due to the correlation between the lagged dependent variable and the individual specific effect. In addition, a feasible GLS estimator for the random-effects model under the assumption of independence between the effects and explanatory variables would also be biased. In these cases (with large N and short T), Andersen & Hsiao (1981) suggests first differencing the model to get rid of the individual effect. On a different approach, but still in a framework of dynamic models with large N and short T, Holtz-Eakin et alli (1988) investigate panel VAR (PVAR) models, in order to provide more flexibility to the VAR modeling for panel data. See also Hsiao (2003, p. 70,107) for further details.

In this paper, due to the specific structure of our VAR, we take a different route. Since we are interested here in weakly stationary variables, in a random-effects model with short N and large T, we apply the Fuller

& Battese (1974) approach to our system of equations (28). These authors establish sufficient conditions for a feasible GLS estimator to be unbiased and exhibit the same asymptotic properties of the GLS estimator in a crossed-error model, i.e., in which an error decomposition is considered to allow for individual effects that are constant over cross sections or time periods. To do so, initially consider the stacked model $Y = X\beta + \varepsilon$, from the system of equations (28), where $Y = (y_{1,1}; y_{1,2}; ...; y_{1,T}; ...; y_{2N,T})$; $X = (x_{1,1}; x_{1,2}; ...; x_{1,T}; ...; x_{2N,T})$; β and $x_{i,t}$ are $p \times 1$ vectors. The Fuller & Battese (1974) two-way random error decomposition is given by $\varepsilon_{i,t} = v_i + e_t + \epsilon_{i,t}$, in which $E(\varepsilon\varepsilon' \mid X) \equiv \Omega$.

Thus, the model is a variance components model, with the variance components $\sigma^2_\epsilon; \sigma^2_v; \sigma^2_e$ to be estimated. A crucial implication of such a specification is that the effects are not correlated with the regressors. For random effects models, the estimation method is a feasible generalized least squares (FGLS) procedure that involves estimating the variance components in the first stage and using the estimated variance covariance matrix thus obtained to apply generalized least squares (GLS) to the data. It is also assumed that $E(v_i) = 0; E(v_i^2) = \sigma^2_v$; $E(v_i v_j) = 0, \forall i \neq j; v_i$ is uncorrelated with $\epsilon_{i,t}, \forall i, t$; and also $E(e_t) = 0; E(e_t^2) = \sigma^2_e$; $E(e_t e_s) = 0, \forall t \neq s; e_t$ is uncorrelated with v_i and $\epsilon_{i,t}, \forall i, t$.[10]

Contrary to Wallace & Hussain (1969) or Swamy & Arora (1972), Fuller & Battese (1974) also consider the case in which σ^2_v and/or σ^2_e are equal to zero.[11] The estimators for the variance components are obtained by the fitting-of-constants method, with the provision that any negative variance components is set to zero for parameter estimation purposes. First, the least square residuals are defined by: $\widehat{\varepsilon} = M_{12}(I - X[X'M_{12}X]^{-1}X'M_{12}]Y$; $\widehat{v} = (M_{12} + M_{1.})(I - X[X'(M_{12} + M_{1.})X]^{-1}X'(M_{12} + M_{1.})]Y$; $\widehat{e} = (M_{12} + M_{.2})(I - X[X'(M_{12} + M_{.2})X]^{-1}X'(M_{12} + M_{.2})]Y$. Next, Fuller & Battese compute the unbiased estimators for the variance components: $\widehat{\sigma}^2_\epsilon = \frac{\widehat{\varepsilon}\widehat{\varepsilon}'}{(N-1)(T-1)-\delta_1}$; $\widehat{\sigma}^2_v = \frac{\widehat{v}\widehat{v}' - [T(N-1)-\delta_2]\widehat{\sigma}^2_\epsilon}{T(N-1)-T\phi_1}$; $\widehat{\sigma}^2_e = \frac{\widehat{e}\widehat{e}' - [N(T-1)-\delta_3]\widehat{\sigma}^2_\epsilon}{N(T-1)-N\phi_2}$ where $\delta_1 \equiv rank(X'M_{12}X)$; $\delta_2 \equiv rank(X'M_{.2}X)$; $\delta_3 \equiv rank(X'M_{1.}X)$; $\phi_1 \equiv tr\{[X'(M_{12} + M_{1.})X]^{-1}X'M_{1.}X\}$; $\phi_2 \equiv tr\{[X'(M_{12} + M_{.2})X]^{-1}X'M_{.2}X\}$.

Once the component variances have been estimated, we form an estimator of the composite residual covariance, and then GLS transform the dependent and regressor data. The respective GLS estimator is given by $\widehat{\beta}_{FB} = (X'\Omega^{-1}X)^{-1}X'\Omega^{-1}Y$ and, thus, the FB estimator is the related feasible GLS estimator $\widehat{\widehat{\beta}}_{FB}$, in which Ω is estimated through $\widehat{\sigma}^2_\epsilon, \widehat{\sigma}^2_v$, and $\widehat{\sigma}^2_e$. Fuller & Battese (1974) show that their estimator is consistent,

[10] The authors also define the following mutually orthogonal, symmetric and idempotent matrices $M_{..} = \frac{J_{2NT}}{2NT}$; $M_{1.} = \frac{I_{2N} \otimes J_T}{T} - M_{..}$; $M_{.2} = \frac{J_{2N} \otimes I_T}{2N} - M_{..}$; $M_{12} = I_{2NT} - \frac{I_{2N} \otimes J_T}{T} - \frac{J_{2N} \otimes I_T}{2N} + M_{..}$; where I_{2N} and I_T are identity matrices of order $2NT$ and T, respectively; and J_{2N} and J_T are $(2N \times 2N)$ and $(T \times T)$ matrices having all elements equal to one. The covariance matrix $\Omega = \sigma^2_\epsilon I_{2NT} + \sigma^2_v(I_{2N} \otimes J_T) + \sigma^2_e(J_{2N} \otimes I_T)$ can be expressed by $\Omega \equiv \sigma^2_\epsilon M_{12} + (\sigma^2_\epsilon + T\sigma^2_v)M_{1.} + (\sigma^2_\epsilon + 2N\sigma^2_e)M_{.2} + (\sigma^2_\epsilon + T\sigma^2_v + 2N\sigma^2_e)M_{..}$, or even, $\Omega = \gamma_1 M_{12} + \gamma_2 M_{1.} + \gamma_3 M_{.2} + \gamma_4 M_{..}$, where $\gamma_1 \equiv \sigma^2_\epsilon$; $\gamma_2 \equiv (\sigma^2_\epsilon + T\sigma^2_v)$; $\gamma_3 \equiv (\sigma^2_\epsilon + 2N\sigma^2_e)$; $\gamma_4 \equiv (\sigma^2_\epsilon + T\sigma^2_v + 2N\sigma^2_e)$.

[11] Baltagi (1981) performed a Monte Carlo study on a single regression equation with two-way error component disturbances and studied the properties of several estimators, including OLS and six feasible GLS estimators: Fuller & Battese (1974), Swamy-Arora (1972), Wallace & Hussain (1969), among others. The results suggest that OLS standard errors are biased and all FGLS are asymptotically efficient and performed relatively well in finite samples, making it difficult to choose among them. The methods differ only in the specifications estimated in evaluating the residuals: The Swamy-Arora estimator of the component variances uses residuals from the within (fixed effect) and between (means) regressions, while the Wallace-Hussain estimator uses only OLS residuals. In general, they provide similar answers, especially in large samples. Additional details on random effects models are provided in Baltagi (2001).

unbiased and asymptotically equivalent to the GLS estimator. In addition, the estimated covariance matrix of coefficients is unbiased, since it is based on unbiased and consistent estimators $\widehat{\sigma}_{\epsilon}^2, \widehat{\sigma}_{v}^2, \widehat{\sigma}_{e}^2$.

Table 2 - Comparison of OLS, SUR and FB covariance matrix of residuals

$E(\varepsilon_{i,t}\varepsilon_{j,s} \mid X)$	OLS	SUR	FB
$i = j; t = s$	σ_i^2	σ_i^2	$(\sigma_v^2 + \sigma_e^2 + \sigma_\epsilon^2)$
$i \neq j; t = s$	0	$\sigma_{i,j}^2$	σ_e^2
$i = j; t \neq s$	0	0	σ_v^2
$i \neq j; t \neq s$	0	0	0

In the SUR approach, the covariance structure allows for conditional correlation between the contemporaneous residuals for cross-section, but restricts residuals in different periods to be uncorrelated. On the other hand, following the argument of Wooldridge (2002, p. 259), rather than depending on $N(N+1)/2$ variances and covariances, as would be the case in a SUR analysis, Ω of the Fuller & Battese (1974) approach only depends on three parameters, $\sigma_\epsilon^2, \sigma_v^2, \sigma_e^2$, regardless of the size of N. This parsimonious feature might be useful for a large panel model, with $N, T \to \infty$. See Baltagi (1980), which investigates a SUR model with error components, and also Pesaran and Smith (1995) and Phillips and Moon (1999) for panel data with large T and N. [12] We next show some important results of the OLS and SUR estimators for the system of equations (28), under the Fuller & Battese error decomposition.

Definition 3: Define the bias on the estimated covariance matrix by: $B \equiv Var(\beta) - E(\widehat{Var}(\beta) \mid X)$;

Definition 4: Define $\Psi \equiv 2N\sigma_e^2/(\sigma_\epsilon^2 + T\sigma_v^2 + 2N\sigma_e^2)$;

Assumption A1: $(X'X)$ and $(X'JX)$ are positive definite matrices, where $J \equiv J_{2N} \otimes I_T$;

Assumption A2: (i) $tr[(X'X)^{-1}X'JX - I] > 0$; and (ii) $T > 2p$, where p is the number of lags of the VAR[13];

Proposition 2 *(OLS) Assume the Fuller & Battese (1974) two-way random error decomposition. (i) If* $(\sigma_\epsilon^2; \sigma_e^2; \sigma_v^2) > 0$ *, then,* $\widehat{\beta}_{OLS}$ *is inconsistent; (ii) if* $\sigma_v^2 = 0$ *and* $(\sigma_\epsilon^2; \sigma_e^2) > 0$*, then,* $\widehat{Var}(\beta_{OLS})$ *is biased; (iii) if* $\sigma_v^2 = 0$, $(\sigma_\epsilon^2; \sigma_e^2) > 0$*, and A1-A2 hold, then,* $diag(\frac{\partial B_{OLS}}{\partial \sigma_e^2}) > 0$*, i.e., an increase of the common shock* σ_e^2 *(and, thus,* Ψ*) induces an upward bias on all estimated (OLS) variances; and (iv) if* $(\sigma_v^2; \sigma_e^2) = 0$ *and* $\sigma_\epsilon^2 > 0$ *, then,* $\widehat{Var}(\beta_{OLS})$ *is unbiased.*

Proof. See Appendix. ■

As already expected, substituting OLS residuals instead of the true disturbances introduces bias in the corresponding estimates of the variance components and, thus, on the covariance matrix of coefficients. See

[12] For panels with large N and T, several approaches might be considered: (i) sequential limits, in which a sequential limit theory is considered; diagonal-path limits, which allows the two indexes to pass to infinity along a specific diagonal path in the two dimensional array; or joint limits, in which both indexes pass to infinity simultaneously. See Hsiao (2003, p.295) for further details, and also Phillips & Moon (1999), which provide sufficient conditions that ensures the sequential limits to be equivalent to joint limits.

[13] Recall from (28) that $\sum_{i=1}^{N} p_i = P$, where p_i is the number of lags of the VAR for a given country i. By assuming that $p_i = p$, $\forall i$, then it follows that $k \equiv 4P = 4Np$, and thus $T > 2p$ means $2NT > 4Np = k$.

Maddala (1971) and Baltagi (2001, p.35) for further details. On the other hand, if assumptions A1-A2 do not hold, then, we cannot guarantee that all estimated variances are upwarded biased. It could be the case that some variances do exhibit a positive bias, whereas others are not affected at all, or even show a negative bias.[14]

Assumption A3: $\sigma_\epsilon^2 I_{2N} + \sigma_e^2 J_{2N} - \widehat{\Sigma} > 0$;

Assumption A4: $\sigma_\epsilon^2 I_{2N} + \sigma_e^2 J_{2N} = \widehat{\Sigma}$;

Assumption A5: $\widehat{\Sigma}$ is a positive semidefinite matrix; and $\Psi \geqslant \frac{2N(2NT-k)}{(tr((X'X)^{-1}X'JX)-2NT)+2N(2NT-k)}$; where $k = 4Np$ and $J \equiv J_{2N} \otimes I_T$;

Proposition 3 *(SUR) Assume the Fuller & Battese (1974) two-way random error decomposition. (i) If* $(\sigma_\epsilon^2; \sigma_e^2; \sigma_v^2) > 0$ *, then,* $\widehat{\beta}_{SUR}$ *is inconsistent; (ii) if* $\sigma_v^2 = 0$, $(\sigma_\epsilon^2; \sigma_e^2) > 0$ *and A3 hold, then,* $\widehat{Var}(\beta_{SUR})$ *is biased. In addition, if A2(ii) and A5 also hold, then,* $(B_{OLS} - B_{SUR}) \geqslant 0$*, i.e., the bias of SUR estimated variances of* β *is not greater than the respective OLS bias; (iii) if* $\sigma_v^2 = 0$, $(\sigma_\epsilon^2; \sigma_e^2) > 0$ *and A4 hold, then,* $\widehat{Var}(\beta_{SUR})$ *is unbiased; and (iv) if* $(\sigma_v^2; \sigma_e^2) = 0$ *and A4 hold, then,* $\widehat{Var}(\beta_{SUR}) = \widehat{Var}(\beta_{OLS})$ *is unbiased.*

Proof. See Appendix. ■

Note that as long as $tr(X(X'X)^{-1}X'J)$ increases, the exigency for Ψ decreases, in order to guarantee that $(B_{OLS} - B_{SUR}) \geqslant 0$. In other words, if the common shock is relatively significant in the disturbance term, then, the SUR technique might produce a less biased covariance matrix in respect to the OLS approach. Now, we present sufficient conditions for the Fuller & Battese (1974) estimator to be unbiased and consistent when applied to our setup. Initially, lets define $\widehat{\beta}_{FB}$ as the unfeasible Fuller & Battese (1974) estimator, and $\widehat{\widehat{\beta}}_{FB}$ as the feasible FB estimator, based on the estimated effects $\widehat{\sigma}_v^2; \widehat{\sigma}_e^2$ and $\widehat{\sigma}_\epsilon^2$.

Assumption A6: $X'\Omega^{-1}X$ is nonsingular, and $plim(\frac{X'\Omega^{-1}X}{T}) = Q_*$, when $T \to \infty$, with fixed N; where Q_* is a finite positive definite matrix;

Assumption A7: e_t and $\epsilon_{i,t}$ are independent and normally distributed;

Assumption A8: $plim[(\frac{X'\widehat{\Omega}^{-1}X}{T}) - (\frac{X'\Omega^{-1}X}{T})] = 0$ and $plim[(\frac{X'\widehat{\Omega}^{-1}\varepsilon}{\sqrt{T}}) - (\frac{X'\Omega^{-1}\varepsilon}{\sqrt{T}})] = 0$; where $\widehat{\Omega} = \widehat{\sigma}_\epsilon^2 I_{2NT} + \widehat{\sigma}_v^2(I_{2N} \otimes J_T) + \widehat{\sigma}_e^2(J_{2N} \otimes I_T)$ is the estimated Fuller & Battese (1974) covariance matrix.

Proposition 4 *Assume the Fuller & Battese (1974) two-way random error decomposition. Thus, it follows that: (i) if* $\sigma_v^2 = 0$, $(\sigma_\epsilon^2; \sigma_e^2) > 0$*, and A6 holds, then, (i) the FB estimator* $\widehat{\beta}_{FB}$ *is unbiased and consistent; (ii) if* $\sigma_v^2 = 0$, $(\sigma_\epsilon^2; \sigma_e^2) > 0$*, and A6-A7 hold, then, the FB estimator* $\widehat{\beta}_{FB}$ *is asymptotically normally distributed, i.e.,* $\widehat{\beta}_{FB} \sim N(\beta; (X'\Omega^{-1}X)^{-1})$*; and (iii) if A6, A7 and A8 hold, then, the feasible FB estimator* $\widehat{\widehat{\beta}}_{FB}$ *is asymptotically equivalent to* $\widehat{\beta}_{FB}$*.*

[14] Note that if $X = [1, ..., 1]'$, then, it follows that $(X'X)^{-1} = 1/2NT$ and $tr[XX'J] = tr[(J_{2NT})(J_{2N} \otimes I_T)] = tr[(J_{2N} \otimes J_T)(J_{2N} \otimes I_T)] = tr[(J_{2N}J_{2N}) \otimes (J_T I_T)] = tr(J_{2N}J_{2N})tr(J_T I_T) = Ttr(J_{2N}J_{2N}) = T(2N)(2N)$. Thus, $tr[(X'X)^{-1}X'JX - I] = tr[(X'X)^{-1}X'JX] - tr(I) = tr[X'JX]/2NT - 2NT = tr[XX'J]/2NT - 2NT = 2N(2NT)/2NT - 2NT = 2N(1-T) < 0$. In other words, if A1-A2 do not hold we cannot guarantee that $diag(\frac{\partial B}{\partial \sigma_e^2}) > 0$.

Proof. See Appendix. ■

Note that if one assumes that the residuals of each country are serially uncorrelated, then, $E(\varepsilon_{i,t}\varepsilon_{i,s} \mid X) = 0$; $\forall t \neq s$, which means that $\sigma_v^2 = 0$.[15] In addition, a fixed effect for v_i would be more appropriate in this framework, since we are focused on a specific set of N countries (see Baltagi, 2001 p.12). On the other hand, for the time component e_t we assume a random effect in order to avoid a significant loss of degrees of freedom due to the large T setup.

In order to verify the finite sample performance of the competing OLS, SUR and FB estimators, we next conduct a small Monte Carlo simulation.

2.5 Monte Carlo simulation

The econometric methodology described in section 2.2 suggests a Wald test to investigate a set of restrictions imposed to the coefficients of a VAR, used to forecast the current account. The Wald test verifies whether or not the optimal current account is statistically equal to the observed current account. Moreover, the Wald test could be conducted based on OLS, SUR and FB estimations for the coefficients of the VAR. Therefore, the goal of our experiment is to investigate the power and the size of the Wald test, comparing different techniques used for the estimation of the VAR (reproduced below for a generic country i):

$$\begin{bmatrix} \Delta Z_t^i \\ CA_t^i \end{bmatrix} = \begin{bmatrix} a^i(L) & b^i(L) \\ c^i(L) & d^i(L) \end{bmatrix} \times \begin{bmatrix} \Delta Z_{t-1}^i \\ CA_{t-1}^i \end{bmatrix} + \begin{bmatrix} \varepsilon_{1t}^i \\ \varepsilon_{2t}^i \end{bmatrix}. \tag{37}$$

One of the critical issues regarding Monte Carlo experiments is that of Data-Generating Processes (DGPs). In our experiment, we construct 100 DGPs, and for each DGP we generate 1,000 samples of the series $\begin{bmatrix} \Delta Z_t & CA_t \end{bmatrix}'$, by sampling random series of ε_t's. Moreover, each sample contains 1,000 observations, but, in order to reduce the impact of initial values we consider only the last $T = 100$ or 200 observations. Thus, the Monte Carlo simulation performs 100,000 replications of the experiment.[16]

Two important issues regarding the companion matrix of the generated series must be addressed at this point. The first one is related to the null hypothesis to be checked by the Wald test: In our simulation, we impose Ho to be true or false by just controlling the impact of the theoretical restrictions into the companion matrix. The magnitude of the theoretical restrictions is given by the gamma parameter, in which $\gamma = 1$ imposes Ho to be true, whereas $\gamma \neq 1$ leads to a false Ho (see appendix for details). The second issue is related to the stationarity of the VAR: In order to apply the econometric methodology, each sample of the experiment must be constructed to generate a covariance-stationary VAR. This way, we show (see also

[15] In addition, the "individual effect" translates in practice into an individual intercept; which is also expected to be zero in our setup, since all series are supposed to be weakly stationary and previously demeaned.

[16] A hybrid solution using E-Views, R and MatLab environments is adopted, since the proposed simulation is extremely computational intensive. We proceed as follows: an E-Views code initially generates the time series ΔZ_t and CA_t for each DGP. Then, it estimates the VAR coefficients based on OLS and SUR techniques and save all the replications and the Wald test results in the hard disk. Next, an R code computes the FB estimator and conducts the respective Wald test, also saving the results in a text file. Finally, a MatLab code reads all the results from the hard disk, computes the size of the Wald test based on the three estimators and also constructs the power results.

appendix for details) how to guarantee the stationarity of the VAR by properly "choosing" the eigenvalues of the companion matrix inside the unit circle, and then calculating the coefficients of the companion matrix that generate those eigenvalues.

Results

Concerning the size of the Wald test, we calculate the estimated significance level by simply observing the frequency of rejection of the null hypothesis in the 100,000 replications of the experiment under conditions where the null hypothesis is imposed to be true. Regarding the power of the test, we also compute the rejection frequencies, but under conditions where the null hypothesis is now imposed to be false.

Table 3 - Size of the Wald test

Model	OLS	SUR	FB
(a) N=2 ; T=100	0.0171	0.0653	0.0534
(b) N=2 ; T=200	0.0154	0.0559	0.0520
(c) N=5 ; T=200	0.0175	0.0620	0.0503

Note: The nominal size of the test is α =5%, in which

empirical size = (frequency of p-values below the α (nominal size) / (MC*DGP);

where (MC*DGP) = 100,000 = total number of replications.

Overall, the results suggest that FB-based test has an adequate size in all cases, whereas the OLS-test exhibits a serious finite sample bias, as already predicted by Proposition 2. In the same line, the SUR-test also seems to show a (small) non-zero bias due to the presence of global shocks, as previously discussed in Proposition 3.

The power investigation can be conducted by controlling the experiment under conditions where the null is imposed to be false, i.e., $\gamma \neq 1$. If the null hypothesis were only "slightly false" (e.g., $\gamma = 0.9$), one would expect power to be lower than if it were "grossly false" (e.g., $\gamma = 0.5$). The results of the power investigation corroborate this expectation and are presented in next tables. To adjust for size distortion, we also report a "size-corrected" power.[17] The results with size correction are quite similar, suggesting that the Wald test might exhibit similar power across the considered estimation procedures.

[17] "Size-corrected power" is just power using the critical values that would have yielded correct size under the null hypothesis.

Table 4 - Power of the Wald test

(a) N=2;T=100	without size-correction			size-corrected power		
γ	OLS	SUR	FB	OLS	SUR	FB
0.9	0.051	0.150	0.126	0.082	0.087	0.104
0.7	0.439	0.541	0.545	0.520	0.533	0.527
0.5	0.725	0.795	0.782	0.729	0.727	0.742
(b) N=2;T=200	without size-correction			size-corrected power		
γ	OLS	SUR	FB	OLS	SUR	FB
0.9	0.165	0.264	0.252	0.241	0.246	0.242
0.7	0.568	0.645	0.635	0.628	0.643	0.627
0.5	0.847	0.887	0.878	0.872	0.881	0.874
(c) N=5;T=200	without size-correction			size-corrected power		
γ	OLS	SUR	FB	OLS	SUR	FB
0.9	0.113	0.226	0.200	0.184	0.205	0.194
0.7	0.544	0.674	0.638	0.642	0.660	0.633
0.5	0.795	0.851	0.834	0.827	0.823	0.826

Notes: a) Power = (frequency of p-values below the α% nominal size) / (MC*DGP),

b) Gamma < 1 indicates a false Ho.

3 Empirical Results

3.1 Data

All data are from the national accounts of IFS – International Financial Statistics (IMF). The CA_t and ΔZ_t series for the G-7 countries are constructed from seasonally adjusted quarterly data (at annual rates), and are expressed in 2000 local currency.[18] In addition, all data are converted in per capita real terms, by dividing it by the implicit GDP deflator and the population. It is worth mentioning that the current account data are not directly obtained from the balance of payments data sets, since these series are not available for all of the countries for an extensive period of time, and it would lead to an arbitrary allocation of "net errors and omissions" in the current account.

Sample 1 (G-7): USA, Canada, Japan, United Kingdom, Germany, Italy, and France.
Period: 1980:q1–2007:q1 (set of 7 countries, 109 time periods, with a total amount of $2NT = 1,526$ observations).

Sample 2: USA, Canada, Japan, United Kingdom, and Germany.
Period: 1960:q1–2007:q1 (set of 5 countries, 189 periods, with a total amount of $2NT = 1,890$ observations).

[18] Based on IMF's World Economic Outlook (April, 2005 - statistical appendix), we adopt the following fixed conversion rates (after 31/12/1998) between the Euro and the currencies of Germany, France and Italy: 1 Euro = 1.95583 Deutsche mark = 6.55957 French francs = 1,936.27 Italian lire.

3.2 Granger Causality

One of the four implications of the theoretical model,[19] listed by Otto (1992), is that CA_t helps to forecast ΔZ_t. According to our results, one can verify that the null hypothesis (CA_t does not Granger-cause ΔZ_t) is rejected at 5% level for Canada and Japan.

Table 5 - Comparison of Results (Ho: CA_t does not Granger-cause ΔZ_t)

	Granger Causality (p-value)			
Country \ Author	**IG**	**O**	**G**	**A**
USA	0.16944	0.0001	0.0004	-
CAN	0.04617	0.21	0.40	-
JPN	0.02426	-	0.62	-
UK	0.47515	-	0.68	-
GER	0.83352	-	0.76	-
ITA	0.91321	-	-	-
FRA	0.59218	-	-	0.10

Notes: IG means Issler & Gaglianone (our results), O refers to Otto (92), G indicates Ghosh (95), A refers to Agénor et al. (1999).

Our results are quite in contrast to the literature, probably due to the different sample periods. For instance, Otto (1992) rejects Ho for the USA (at 1% level), but does not reject it for Canada. Ghosh (1995) also rejects Ho for the USA (at 1% level) and does not reject it for Canada, Japan, UK and Germany, and Agénor et al. (1999) present a p-value of 0.10 for France. The different results could possibly be explained by the broader range of our sample period, in comparison to the previous studies, which do not account for all global and idiosyncratic shocks occurred in the last decades: Our sample period covers quarterly data from 1960 until 2007, whereas Otto (1992) considers the period 1950-88, Ghosh(1995) studies the period 1960-88, and Agénor et al. (1999) covers 1970-96.

Thus, our results indicate that, with the exception of Canada and Japan, the current account does not help to forecast the net output of the G-7 countries, indicating that the agents possibly do not have any additional information to predict ΔZ_t, other than those contained in the past of their own series. Recall Proposition 1, which states that if the Granger causality implication is not verified, then, the optimal current account should not be generated, since it would lead to spurious results. In the present work, the Granger causality is only verified for Canada and Japan. Thus, for the other countries, the model should be rejected and the optimal current account should not be generated. For instance, the case of UK (sample 1) can be analyzed as an example of spurious result. The VAR(1) estimated for this country is given by

VAR coefficients for **UK**

$$\begin{bmatrix} \Delta Z_t \\ CA_t \end{bmatrix} = \begin{bmatrix} a_1 = -0.211322 \ \text{(-2.40)} & b_1 = -0.066236 \ \text{(-1.76)} \\ c_1 = -0.133631 \ \text{(-1.12)} & d_1 = 0.854978 \ \text{(16.99)} \end{bmatrix} \times \begin{bmatrix} \Delta Z_{t-1} \\ CA_{t-1} \end{bmatrix} + \begin{bmatrix} 5.477849 \ \text{(1.67)} \\ 2.626709 \ \text{(0.59)} \end{bmatrix} + \begin{bmatrix} \varepsilon_{1t} \\ \varepsilon_{2t} \end{bmatrix}$$

Note: t-statistics in parentheses

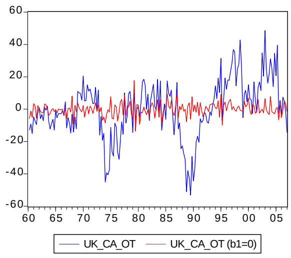
[19] The results of ADF unit root test and the cointegration analysis are presented in appendix.

where a_1 and d_1 are statistically significant, but this is not the case for b_1 and c_1. As described in table 6, the vector K (recall equation (22)) is extremely sensitive to variations in b_1 and could generate completely different CA_t^* series.[20] Assuming $b_1 = -0.066$ (instead of zero, since b_1 is not statistically significant), the model indicates that $\beta = 0.348$ (unlike the correct value of $\beta = 0$).

Table 6 - Vector $K = [\alpha; \beta]$ for UK

	Assume $b_1 = -0.066$	Assume $b_1 = 0$
α	0.134	0.172
β	0.348	0.000

Figure 1 - CA_t^* for UK



Note: The picture above exhibits two optimal current accounts for UK (sample 1), generated from different K vectors of table 6.

3.3 Correlation matrix

The residual correlation matrix obtained in the joint estimation of the VAR could be a starting point to justify the SUR technique, since the contemporaneous correlation across the G-7 countries should not be ignored.

Table 7 - Residual Correlation Matrix (sample 1)

	USA_DZ	USA_CA	CAN_DZ	CAN_CA	JPN_DZ	JPN_CA	FRA_DZ	FRA_CA	UK_DZ	UK_CA	GER_DZ	GER_CA	ITA_DZ	ITA_CA
USA_DZ	1.00	0.26												
USA_CA	0.26	1.00												
CAN_DZ	0.15	0.04	1.00	0.53										
CAN_CA	0.02	0.18	0.53	1.00										
JPN_DZ	0.13	0.15	-0.09	0.01	1.00	0.06								
JPN_CA	0.11	0.00	0.05	0.00	0.06	1.00								
FRA_DZ	0.07	0.04	-0.03	-0.11	0.25	-0.06	1.00	0.32						
FRA_CA	0.10	0.08	-0.13	-0.12	0.17	-0.02	0.32	1.00						
UK_DZ	0.02	-0.05	0.13	0.15	-0.12	0.00	-0.10	0.01	1.00	0.44				
UK_CA	-0.13	-0.08	0.06	0.08	-0.01	-0.06	0.04	0.05	0.44	1.00				
GER_DZ	0.10	0.16	-0.05	0.07	0.05	0.10	0.15	-0.04	-0.08	-0.06	1.00	0.20		
GER_CA	0.21	0.01	0.19	0.05	-0.04	0.04	0.11	0.09	0.17	0.12	0.20	1.00		
ITA_DZ	0.09	0.21	0.04	-0.06	-0.02	0.06	0.16	0.17	0.07	0.09	0.08	0.14	1.00	0.49
ITA_CA	0.02	0.28	0.07	0.03	0.13	-0.03	0.08	0.30	0.08	-0.08	0.11	0.07	0.49	1.00

Furthermore, a Likelihood Ratio (LR) test is able to provide a formal argument to adopt the SUR approach, instead of the OLS technique. Under the null hypothesis, the residuals covariance matrix (Ω) is a diagonal (band)[21] matrix, suggesting the OLS method. On the other hand, the alternative specification (H_1) supposes that Ω is a non-diagonal (band) matrix, recommending the SUR approach. This way, Ho imposes a set of restrictions on the residuals covariance matrix, since all elements out of the diagonal (band) are set to zero. In this case, according to Hamilton (1994), twice the log likelihood ratio for a Gaussian VAR is given by

$$2\left(L_1^* - L_0^*\right) = T\left(\ln\left|\widehat{\Omega}_0\right| - \ln\left|\widehat{\Omega}_1\right|\right), \tag{38}$$

[20] Optimal current account is generated by equations (16) and (17), and depends on the estimated coefficients of the VAR and the world interest rate (supposed 2% per year).

[21] The residuals covariance matrix is diagonal-band, in OLS estimation, because of the structure of the VAR, since each country has two equations (CA_t and ΔZ_t).

where L_0^* is the maximum value for the log likelihood under Ho (and L_1^* under the alternative hypothesis), T is the number of effective observations, $\left|\widehat{\Omega}_0\right|$ is the determinant of the residuals covariance matrix estimated by OLS, and $\left|\widehat{\Omega}_1\right|$ is the determinant of the same matrix estimated by SUR. Under the null hypothesis, the difference between L_1^* and L_0^* is statistically zero, and the LR statistic asymptotically follows a χ^2 distribution, with degrees of freedom equal to the number of restrictions imposed under Ho.

Table 8 - Results of the LR test

	Sample 1	Sample 2
T	1,520	1,870
$\left\|\widehat{\Omega}_0\right\|_{OLS}$	1.11734E+51	1.53125E+38
$\left\|\widehat{\Omega}_1\right\|_{SUR}$	4.47295E+50	1.23388E+38
ζ_{LR}	1,391.53	403.77
ζ_{LR}^*	1,385.13	402.91
5% critical value	$\chi^2_{(80)}$=101.88 ; $\chi^2_{(90)}$=113.15	$\chi^2_{(40)}$=55.76
1% critical value	$\chi^2_{(80)}$=112.33 ; $\chi^2_{(90)}$=124.12	$\chi^2_{(40)}$=63.69

Notes: a) T is the number of effective observations, and $\zeta_{LR} = T\left(\ln\left|\widehat{\Omega}_0\right| - \ln\left|\widehat{\Omega}_1\right|\right)$ is the LR statistic.

b) ζ_{LR}^* is a modification to the LR test to take into account small-sample bias, replacing T by $(T-k)$, where k is the number of parameters estimated per equation.

c) In sample 1, the degrees of freedom (dof)= 84, and in sample 2, dof=40.

Hence, the null hypothesis could be rejected in both samples, since the LR statistics are larger than the critical values. Therefore, the residuals covariance matrices are non-diagonal (band), and the SUR approach is better recommended than the OLS method.

3.4 Wald test

A formal comparison between $(CA_t^* - \mu_{CA}^*)$ and $(CA_t - \mu_{CA})$, to measure the fit of the model with the data, is provided by the restrictions (20) imposed to the coefficients of the VAR, through a Wald test, which asymptotically follows a χ^2 distribution (with the degrees of freedom equal to the number of restrictions). The acceptance of those restrictions in the Wald test means that both series of current account (optimal and observed) are statistically the same.[22]

We perform the Wald test for the G-7 countries based on two different types of time series. The first one is the usual time series suggested by the literature (e.g., Ghosh, 1995), in which the Wald test is conducted from a seasonally adjusted quarterly data (at annual rates), expressed in 2000 local currency, converted in per capita real terms, by dividing it by the implicit GDP deflator and the population. In order to compute common shocks among the considered countries, we also convert all series to 2000 U.S. dollars. In the second approach, however, data is not expressed in per capita terms and is converted to U.S. dollars by using a

[22] The Wald test can be implemented for several values of the world interest rate (r). However, the results are almost the same for values of r ranging from 1% to 6%. This way, we have adopted $r = 2\%$ following the literature.

proper exchange rate time series, due to the fact that the global shocks must be computed from a proper set of scaled and comparable time series. Since both methodologies lead to very similar results, we next present the results only for the later approach, based on different estimation procedures: OLS, SUR, FB.1 and FB.2 (see tables 13 and 14 in appendix for further details).[23]

The first estimator is merely OLS equation-by-equation, ignoring possible crossed effects among countries. The second approach, SUR, is just a feasible GLS estimator applied to the considered system of equations and, as already mentioned, might improve the efficiency when compared to the OLS method and, thus, the covariance matrix of a SUR estimation could result in a rejection of the model previously accepted in the OLS framework. The last estimators, FB.1 and FB.2, are based on a two-way random error decomposition procedure of Fuller & Battese (1974), in which the residual term is decomposed into individual effects, common shocks, and idiosyncratic terms, i.e., $\varepsilon_{i,t} = v_i + e_t + \epsilon_{i,t}$. The only difference between these two estimators is that FB.1 assumes a unique global shock e_t for the whole system of equations, whereas, FB.2 considers a common shock e_t^{CA} for the CA_t system of equations, and a different shock e_t^{DZ} for the ΔZ_t equations.

A comparison of these results with the empirical evidence found in the literature is presented in Table 9: Otto (1992) rejects the model for the USA and Canada, Ghosh (1995) rejects the model for Canada, Japan, the UK and Germany, but does not reject it for the USA, and Agénor et al. (1999) do not reject it for France.

Table 9 - Comparison with the literature

	Wald Test (p-value)						
Country	**OLS**	**SUR**	**FB.1**	**FB.2**	**Otto (92)**	**Ghosh (95)**	**Agénor et al.(99)**
USA (1)	0.1400	0.0207 (*)	0.034 (*)	0.0394 (*)	0.0041 (**)	1.19	-
CAN (1)	0.2402	0.0312 (*)	0.8692	0.8766	0.0020 (**)	95 (**)	-
JPN (1)	0.0291 (*)	0.0205 (*)	0.0032 (**)	0.0051 (**)	-	75 (*)	-
UK (1)	0.1796	0.1550	0.9970	0.9979	-	464 (**)	-
GER (1)	0.048 (*)	0.0097 (**)	0.6667	0.6898	-	90 (**)	-
ITA (1)	0.1425	0.0789	0.9771	0.9594	-	-	-
FRA (1)	0.0295 (*)	0.0000 (**)	0.7957	0.7525	-	-	0.314
USA (2)	0.0291 (*)	0.0069 (**)	0.0000 (**)	0.0000 (**)			
CAN (2)	0.1499	0.0123 (*)	0.8288	0.8361			
JPN (2)	0.0201 (*)	0.0160 (*)	0.0363 (*)	0.3019			
UK (2)	0.0139 (*)	0.0001 (**)	0.8727	0.6861			
GER (2)	0.0143 (*)	0.0078 (**)	0.7161	0.3260			

Notes: Ghosh (1995) presents chi-squared values; $USA_{(1)}$ indicates sample 1 and $USA_{(2)}$ means sample 2; (**) means rejection at 1% level, and (*) at 5% level.

First of all, note that the SUR estimator generally over rejects the model in comparison to OLS.[24] However, recall from Proposition 1 that a fair analysis of Table 9 should only consider countries in which

[23] EViews 5.1 was used to obtain the OLS and SUR estimators, whereas a code in R was developed for the FB estimators, which could alternatively be computed in SAS.

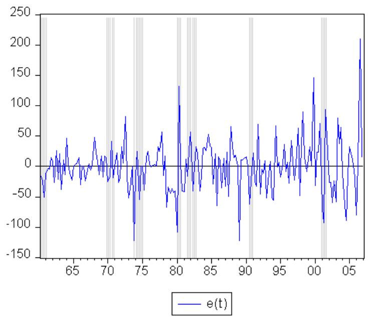
[24] An important remark is provided by Ghosh and Ostry (1995), which argue that the non-rejection of the model for a given country can occur because of the magnitude of the standard deviations in the coefficients of the VAR. High values for the standard errors could lead to a statistical equality between the optimal and observed current accounts, even if these series are graphically different.

the Granger causality can (indeed) be verified. Since from Table 5 it only occurs for Canada and Japan, and the later country failed at the Wald test, our empirical results suggest, contrary to the previous results found in the literature, that the PVM model of the current account cannot be rejected for Canada.

Secondly, note that due to the possible finite sample bias of OLS and SUR estimators, the results are quite different from the FB results, as already expected from Propositions 2-4. In fact, regarding the FB results, the Hausman (1978) m-statistic, that provides information about the appropriateness of the random effects specification, do not indicate a rejection[25] of the null hypothesis of zero correlation between regressors and effects.

More importantly, in both samples the FB approach suggest that $\sigma_v^2 = 0$, and that the global shock component (e_t) should not be ignored in the estimation process, since its relative importance in respect to the total residual is estimated as $\Psi = 0.42$ in sample 1, and $\Psi = 0.36$ in sample 2. Recall that $\Psi \equiv 2N\sigma_e^2/(\sigma_\epsilon^2 + T\sigma_v^2 + 2N\sigma_e^2)$ should be zero, in the case of no global shocks, and also recall Proposition 2(ii), which states that the finite sample bias of the OLS estimated covariance matrix increases as long as the common shocks become more present in the data. A global shock (e_t) time series is depicted in next figure and compared (for illustrative purposes) to U.S. recessions:

Figure 2 - Global shocks (sample 2) and U.S. recessions



Note: Gray bars represent the U.S. (NBER) recessions.

Note that some of the U.S. recessions indeed coincide with the negative peaks of e_t, including the most recent period in 2001, which is a natural result since the global shocks in the current account of the G-7 are expected to be (at least) partially driven by the world´s biggest economy movements.

4 Conclusions

The standard intertemporal optimization model of the current account is adopted to analyze the G-7 countries. In this framework, the perfect capital mobility allows the agents to smooth consumption via current account. The econometric approach of the model, developed by Campbell & Shiller (1987), consists of estimating an unrestricted VAR to verify the adherence of the theoretical framework onto the data. Furthermore,

[25] For instance, in sample 1, the FB.1 estimator, with 28 degrees of freedom, exhibit the m statistic equal to 11.19779 (p-value: 0.9980223); and for FB.2, with 14 degrees of freedom in each system, m=4.139241 (p-value: 0.9945687).

a Wald test is used to investigate a set of restrictions imposed to the VAR, used to forecast the current account, testing whether or not the optimal current account is equal to the observed series (null hypothesis).

In spite of all theoretical advances regarding the current account PVM models in recent years, we have opted to investigate some econometric techniques that could be used in the estimation process, in order to answer an important related question: Could an inappropriate econometric technique lead to wrong conclusions regarding the rejection of the current account PVM model? We focused on three estimation techniques (OLS, SUR and FB).

Firstly, a SUR technique could be recommended (instead of OLS) in order to properly consider contemporaneous correlations across the considered countries, that might be caused by global shocks such as the oil shocks of 70s or the financial crises of 90s. However, due to potential pitfalls associated with the SUR estimation, we also provide an application of the Fuller & Battese (1974) two-way random error decomposition, due to the existence of common shocks and the possible bias in the covariance matrix estimated with OLS and SUR techniques.

We investigate these estimators in a Monte Carlo experiment, and evaluate the power and size of the Wald test in the presence of global shocks, concluding that the FB-based test exhibits a good performance in the size investigation. On the other hand, the OLS-based test performs unsatisfactorily, and a small finite sample bias can also be detected for the SUR approach. The numerical simulations also indicate that there is no clear difference in the size-corrected power of the considered Wald test based on different estimators.

To summarize, the theoretical methodology, as well as the Monte Carlo simulations, suggests the adoption of the FB estimator (instead of OLS) in the presence of global shocks, small N and relatively large T. The SUR approach could be used in this case, but only when the common shock exhibits a low magnitude, since it directly leads to a biased estimated covariance matrix.

This paper also proposes a note on Granger causality and Wald tests: the Granger causality is addressed as a "sine qua non" condition for the entire validation of the model, since the construction of the optimal current account leads to spurious results when this condition is not verified. We further provide an empirical exercise by estimating the model for the G-7 countries. Indeed, the results substantially change with the application of the different estimation techniques.

More importantly, the FB framework indicates that the common shock in the G-7 countries can account for almost 40% of the total residuals, suggesting that the previous (OLS) estimations might be seriously biased, and the familiar inference procedures would no longer be appropriate. The error-decomposition procedure only indicates a rejection of the model for the USA and Japan, which is in sharp contrast to the previous literature. Putting all together, these findings suggest that the PVM cannot be rejected for Canada, and cast serious doubts to some results presented in the literature, based on OLS estimation, that ignore the presence of common shocks among countries.

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Appendix A. Note on Granger Causality and Wald tests

A VAR(2) model can be rewritten as a VAR(1) in the following way:

$$\begin{bmatrix} \Delta Z_t \\ \Delta Z_{t-1} \\ CA_t \\ CA_{t-1} \end{bmatrix} = \begin{bmatrix} a_1 & a_2 & b_1 & b_2 \\ 1 & 0 & 0 & 0 \\ c_1 & c_2 & d_1 & d_2 \\ 0 & 0 & 1 & 0 \end{bmatrix} \times \begin{bmatrix} \Delta Z_{t-1} \\ \Delta Z_{t-2} \\ CA_{t-1} \\ CA_{t-2} \end{bmatrix} + \begin{bmatrix} \varepsilon_{1t} \\ 0 \\ \varepsilon_{2t} \\ 0 \end{bmatrix}$$

or in a compact form $X_t = AX_{t-1} + \varepsilon_t$. Hence, the vector K is given by

$$K = \begin{bmatrix} \alpha_1 & \alpha_2 & \beta_1 & \beta_2 \end{bmatrix} = -h'(\frac{A}{1+r})(I_4 - \frac{A}{1+r})^{-1}$$

$$\therefore K = -\begin{bmatrix} 1 & 0 & 0 & 0 \end{bmatrix}(\frac{1}{1+r}\begin{bmatrix} a_1 & a_2 & b_1 & b_2 \\ 1 & 0 & 0 & 0 \\ c_1 & c_2 & d_1 & d_2 \\ 0 & 0 & 1 & 0 \end{bmatrix})(\begin{bmatrix} 1 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 \\ 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 1 \end{bmatrix} - \frac{1}{1+r}\begin{bmatrix} a_1 & a_2 & b_1 & b_2 \\ 1 & 0 & 0 & 0 \\ c_1 & c_2 & d_1 & d_2 \\ 0 & 0 & 1 & 0 \end{bmatrix})^{-1}$$

Thus, after some algebraic manipulations, the β_1 coefficient of the vector K is given by:

$$\begin{aligned} \beta_1 &= -\frac{a_1(1+r)(b1+b1r+b2)}{\phi} - \frac{a_2(b1+b1r+b2)}{\phi} \\ &\quad -\frac{b1(1+r)(1+2r+r^2-a1-a1r-a2)}{\phi} - \frac{b2(1+2r+r^2-a1-a1r-a2)}{\phi} \end{aligned}$$

$$\begin{aligned} \text{where } \phi &= 1 + 2a_1d_1r + 4r + a_1d_1r^2 - c_1b_2r - 2c_1b_1r + a_2d_1r - d_1 - d_2 - a_2 - a_1 - c_2b_1r + 6r^2 \\ &\quad -3d_1r - 2a_2r - a_2r^2 + a_2d_1 + a_2d_2 - c_2b_1 - c_2b_2 + a_1d_2r - c_1b_1r^2 + 4r^3 + r^4 \\ &\quad -3d_1r^2 - 2rd_2 - 3a_1r - 3a_1r^2 + a_1d_1 + a_1d_2 - d_1r^3 - r^2d_2 - a_1r^3 - c_1b_1 - c_1b_2 \end{aligned}$$

In this case, the optimal current account is given by

$$(CA_t^* - \mu_{CA^*}) = K(X_t - \mu) = \begin{bmatrix} \alpha_1 & \alpha_2 & \beta_1 & \beta_2 \end{bmatrix} \begin{bmatrix} \Delta Z_t - \mu_{\Delta Z} \\ \Delta Z_{t-1} - \mu_{\Delta Z} \\ CA_t - \mu_{CA} \\ CA_{t-1} - \mu_{CA} \end{bmatrix}$$

and the Wald test analyzes the joint restrictions: $\alpha_1 = \alpha_2 = \beta_2 = 0$ and $\beta_1 = 1$. Again, one should note that if the Granger causality is not verified (e.g., $b_1 = b_2 = 0$) then $\beta_1 = 0$. In this manner, the implication of Granger causality becomes a necessary condition for the validation of the VAR(2) model. The generalization of this result to a VAR(p) framework is straightforward, as presented in the proof of Proposition 1.

Appendix B. Proof of Propositions

Proof of Proposition 1. The VAR(p) can be rewritten as a VAR(1), in the following way

$$\begin{bmatrix} \Delta Z_t \\ \vdots \\ \vdots \\ \Delta Z_{t-p+1} \\ CA_t \\ \vdots \\ \vdots \\ CA_{t-p+1} \end{bmatrix} = \begin{bmatrix} a_1 & \cdots & \cdots & a_p & b_1 & \cdots & \cdots & b_p \\ 1 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & \ddots & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 1 & 0 & 0 & 0 & 0 & 0 \\ c_1 & \cdots & \cdots & c_p & d_1 & \cdots & \cdots & d_p \\ 0 & 0 & 0 & 0 & 1 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & \ddots & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 1 & 0 \end{bmatrix} \begin{bmatrix} \Delta Z_{t-1} \\ \vdots \\ \vdots \\ \Delta Z_{t-p} \\ CA_{t-1} \\ \vdots \\ \vdots \\ CA_{t-p} \end{bmatrix} + \begin{bmatrix} \varepsilon_{1t} \\ 0 \\ \vdots \\ 0 \\ \varepsilon_{2t} \\ 0 \\ \vdots \\ 0 \end{bmatrix}$$

or in the same compact form $X_t = AX_{t-1} + \varepsilon_t$. Supposing that the Granger causality is not provided by the data set, it follows that $b(L) = 0$, and the companion matrix A becomes:

$$A = \begin{bmatrix} A_{11} & 0 \\ A_{21} & A_{22} \end{bmatrix}$$

where $A_{11} = \begin{bmatrix} a_1 & \cdots & \cdots & a_p \\ 1 & 0 & 0 & 0 \\ 0 & \ddots & 0 & 0 \\ 0 & 0 & 1 & 0 \end{bmatrix}$; $A_{21} = \begin{bmatrix} c_1 & \cdots & \cdots & c_p \\ 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 \end{bmatrix}$; $A_{22} = \begin{bmatrix} d_1 & \cdots & \cdots & d_p \\ 1 & 0 & 0 & 0 \\ 0 & \ddots & 0 & 0 \\ 0 & 0 & 1 & 0 \end{bmatrix}$

In this case, vector K can be written as $K = -h'BC^{-1}$, where the matrices B and C are defined as $B = (\frac{A}{1+r})$ and $C = (I_{2p} - \frac{A}{1+r})$. It should be noted that matrices B and C are also partitioned matrices with a null upper-right block. According to Simon & Blume (1994, p.182), in this case, the inverse of the partitioned matrix C will also result in a null upper-right block matrix (see Theorem reproduced below).

Theorem 8.15 (Simon & Blume, 1994): Let C be a square matrix partitioned as $C = \begin{bmatrix} C_{11} & C_{12} \\ C_{21} & C_{22} \end{bmatrix}$ where C_{11} and C_{22} are square submatrices. If both C_{22} and $D \equiv C_{11} - C_{12}C_{22}^{-1}C_{21}$ are nonsingular, then C is nonsingular and

$$C^{-1} = \begin{bmatrix} D^{-1} & -D^{-1}C_{12}C_{22}^{-1} \\ -C_{22}^{-1}C_{21}D^{-1} & C_{22}^{-1}\left(I + C_{21}D^{-1}C_{12}C_{22}^{-1}\right) \end{bmatrix}$$

Thus, in our case, $C_{12} = 0$ and the term $-D^{-1}C_{12}C_{22}^{-1}$ becomes a null submatrix, suggesting that C^{-1} and the following product BC^{-1} are also NURB matrices. Finally, the vector $K = \begin{bmatrix} \alpha_1 & \ldots & \alpha_p & \beta_1 & \ldots & \beta_p \end{bmatrix}$ is given by selecting (through the vector $-h'$) the first line from the matrix (BC^{-1}), suggesting that all β_i $(i = 1, \ldots, p)$ coefficients are zero. ■

Proof of Proposition 2. (i) Consider the system of equations for any country i: $\begin{cases} \Delta Z_{i,t} = \mu_i^{\Delta Z} + a_i(L)\Delta Z_{i,t} + b_i(L)CA_{i,t} + \varepsilon_{i,t}^{\Delta Z} \\ CA_{i,t} = \mu_i^{CA} + c_i(L)\Delta Z_{i,t} + d_i(L)CA_{i,t} + \varepsilon_{i,t}^{CA} \end{cases}$; where $a_i(L)$, $b_i(L)$, $c_i(L)$ and $d_i(L)$ are polynomials of order p, and $\varepsilon_{i,t}^{\Delta Z} = (v_i^{\Delta Z} + e_t^{\Delta Z} + \epsilon_{i,t}^{\Delta Z})$. The equation for $\Delta Z_{i,t-1}$ can be expressed by: $(a_{i,2}\Delta Z_{i,t-2} + a_{i,3}\Delta Z_{i,t-3} + ... + a_{i,p+1}\Delta Z_{i,t-p+1}) + b_i(L)CA_{i,t-1} + \varepsilon_{i,t-1}^{\Delta Z}$. Note that $E(\varepsilon_{i,t}^{\Delta Z}\Delta Z_{i,t-1}) = E(v_i^{\Delta Z} + e_t^{\Delta Z} + \epsilon_{i,t}^{\Delta Z})(a_{i,2}\Delta Z_{i,t-2} + a_{i,3}\Delta Z_{i,t-3} + ... + a_{i,p+1}\Delta Z_{i,t-p+1} + b_i(L)CA_{i,t-1} + v_i^{\Delta Z} + e_{t-1}^{\Delta Z} + \epsilon_{i,t-1}^{\Delta Z}) = E(v_i^{\Delta Z}v_i^{\Delta Z}) = \sigma_{v,\Delta Z}^2 > 0 \therefore E(X'\varepsilon) \neq 0$ and $\widehat{\beta}_{ols}$ becomes inconsistent;

(ii) Recall that $Var(\beta_{OLS}) = \sigma^2(X'X)^{-1}$ and $\widehat{Var}(\beta_{OLS}) = s^2(X'X)^{-1} = (X'X)^{-1}[\frac{e'e}{2NT-k}]$; where e is the estimated OLS residual and $k = 4Np$. Thus, $\widehat{Var}(\beta_{OLS}) = \frac{(X'X)^{-1}}{2NT-k}[Y'(I - X(X'X)^{-1}X')Y] = \frac{(X'X)^{-1}}{2NT-k}tr[Y'MY]$; where $M \equiv I - X(X'X)^{-1}X'$ is an idempotent and symmetric matrix, in which $MX = 0$. This way, $tr[Y'MY] = tr[Y'M'MY] = tr[\varepsilon'M'M\varepsilon] = tr[M\varepsilon\varepsilon']$; and $\widehat{Var}(\beta_{OLS}) = \frac{(X'X)^{-1}}{2NT-k}tr[M\varepsilon\varepsilon']$. Therefore, $E[\widehat{Var}(\beta_{OLS}) \mid X] = E[\frac{(X'X)^{-1}}{2NT-k}tr(M\varepsilon\varepsilon') \mid X] = \frac{(X'X)^{-1}}{2NT-k}tr[M(E(\varepsilon\varepsilon' \mid X))] = (X'X)^{-1}\frac{tr[M\Omega]}{2NT-k} = (X'X)^{-1}\varphi$, where $\varphi \equiv \frac{tr[M\Omega]}{2NT-k}$. One can also rewrite last expression as $E[\widehat{Var}(\beta_{OLS}) \mid X] = (X'X)^{-1}\{\varphi X'X\}$ $(X'X)^{-1} = (X'X)^{-1}\{\varphi X'M_{12}X + \varphi X'M_{1.}X + \varphi X'M_{.2}X + \varphi X'M_{..}X)\}(X'X)^{-1}$, where the decomposition of X, based on the idempotent and symmetric matrices $\{M_{12}; M_{1.}; M_{.2}; M_{..}\}$, was applied in last equality. Now, assume that the considered bias in the variance estimator is zero, i.e., $E[\widehat{Var}(\beta_{OLS}) \mid X] = Var(\beta_{OLS})$. Note that this is true if and only if $\varphi = \gamma_1 = \gamma_2 = \gamma_3 = \gamma_4$. But this is a contradiction since, by assumption, $(\sigma_\epsilon^2; \sigma_e^2) > 0$ and $\sigma_v^2 = 0$, and thus $\gamma_1 = \gamma_2$ and $\gamma_3 = \gamma_4$, but $\gamma_1 \neq \gamma_3$;

(iii) It follows that $\frac{\partial Var(\beta_{OLS})}{\partial\sigma_e^2} = (X'X)^{-1}\{2N(X'(M_{.2}+M_{..})X)\}(X'X)^{-1} = (X'X)^{-1}\{2N(X'(\frac{J_{2N}\otimes I_T}{2N})X)\}$ $(X'X)^{-1} = (X'X)^{-1}X'JX\}(X'X)^{-1}$; where $J \equiv J_{2N}\otimes I_T$. Recall that $E[\widehat{Var}(\beta_{OLS}) \mid X] = (X'X)^{-1}\frac{tr[M\Omega]}{2NT-k}$ and, thus, $\frac{\partial E[\widehat{Var}(\beta_{OLS})|X]}{\partial\sigma_e^2} = \frac{(X'X)^{-1}}{2NT-k}(\frac{\partial tr[M\Omega]}{\partial\sigma_e^2}) = \frac{(X'X)^{-1}}{2NT-k}(\frac{\partial vec'(\Omega)}{\partial\sigma_e^2})vec(M') = \frac{(X'X)^{-1}}{2NT-k}(2N(vec'(M_{.2} + M_{..})))vec(M') = \frac{(X'X)^{-1}}{2NT-k}vec'(J)vec(M') = \frac{(X'X)^{-1}}{2NT-k}tr(MJ) = (X'X)^{-1}\frac{tr(MJ)}{tr(M)}$. Therefore, $\frac{\partial B_{OLS}}{\partial\sigma_e^2} = \frac{\partial Var(\beta_{OLS})}{\partial\sigma_e^2}$ $-\frac{\partial E[\widehat{Var}(\beta_{OLS})|X]}{\partial\sigma_e^2} = (X'X)^{-1}X'JX\}(X'X)^{-1} - (X'X)^{-1}\frac{tr(MJ)}{tr(M)} = (X'X)^{-1}\{X'JX - \frac{tr(MJ)}{tr(M)}X'X\}(X'X)^{-1}$ $\equiv B'AB$, where $A \equiv X'JX - \frac{tr(MJ)}{tr(M)}X'X$ and $B \equiv (X'X)^{-1}$. Now, let $\alpha \equiv \frac{tr(MJ)}{tr(M)} = \frac{tr[(I-X(X'X)^{-1}X')J]}{2NT-k}$ $= \frac{tr[J-X(X'X)^{-1}X'J]}{2NT-k} = \frac{tr[J]-tr[X(X'X)^{-1}X'J]}{2NT-k} = \frac{tr[J_{2N}\otimes I_T]-tr[X(X'X)^{-1}X'J]}{2NT-k} = \frac{2NT-tr[X(X'X)^{-1}X'J]}{2NT-k}$. By assumption A2(ii), it follows that $T > 2p \therefore 2NT > 4Np = k \therefore 2NT - k > 0$. By A2(i), we have that $tr[(X'X)^{-1}X'JX] - tr(I_{2NT}) > 0 \therefore tr[X(X'X)^{-1}X'J] > 2NT$. Thus, by A2 it follows that the scalar $\alpha < 0$. Now, recall that $A \equiv X'JX - \frac{tr(MJ)}{tr(M)}X'X = X'JX - \alpha X'X = X'JX + \delta X'X$; where $\delta \equiv (-\alpha) > 0$. By A1 we have that $(X'X) > 0$ and $(X'JX) > 0$, thus, it follows that $\delta X'X > 0$ and $A > 0$. Therefore, since B is non-singular, and A is a positive definite matrix, then, $\frac{\partial B_{OLS}}{\partial\sigma_e^2} = B'AB > 0 \therefore diag(\frac{\partial B_{OLS}}{\partial\sigma_e^2}) > 0$;

(iv) If $\sigma_v^2 = \sigma_e^2 = 0$, then, $\Omega \equiv \sigma_\epsilon^2 M_{12} + (\sigma_\epsilon^2 + T\sigma_v^2)M_{1.} + (\sigma_\epsilon^2 + N\sigma_e^2)M_{.2} + (\sigma_\epsilon^2 + T\sigma_v^2 + N\sigma_e^2)M_{..} = \sigma_\epsilon^2(M_{12} + M_{1.} + M_{.2} + M_{..}) = \sigma_\epsilon^2 I_{2NT}$; and thus $tr(M\Omega) = \sigma_\epsilon^2 tr(M)$. Therefore, $E[\widehat{Var}(\beta_{OLS}) \mid X] = (X'X)^{-1}\frac{tr[M\Omega]}{2NT-k} =$

$\sigma^2_\epsilon(X'X)^{-1}\frac{tr[M]}{2NT-k} = \sigma^2_\epsilon(X'X)^{-1}\frac{tr[I_{2NT}-X(X'X)^{-1}X']}{2NT-k} = \sigma^2_\epsilon(X'X)^{-1}\frac{(2NT-k)}{2NT-k} = \sigma^2_\epsilon(X'X)^{-1} = Var(\beta_{OLS}) \therefore B_{OLS} = 0$. ■

Proof of Proposition 3. (i) Since the OLS estimator is inconsistent, the OLS residuals would lead to a feasible GLS estimator (and thus SUR) to be inconsistent as well, since the model contains variables that are correlated with the residuals;

(ii) From the Fuller & Battese (1974) two-way random error decomposition, it follows that $\Omega = \gamma_1 M_{12} + \gamma_2 M_{1.} + \gamma_3 M_{.2} + \gamma_4 M_{..}$. By applying the definition of the previous M_{ii} matrices, it follows that $\Omega = \gamma_1(I_{2NT} - \frac{I_{2N}\otimes J_T}{T} - \frac{J_{2N}\otimes I_T}{2N} + \frac{J_{2NT}}{2NT}) + \gamma_2(\frac{I_{2N}\otimes J_T}{T} - \frac{J_{2NT}}{2NT}) + \gamma_3(\frac{J_{2N}\otimes I_T}{2N} - \frac{J_{2NT}}{2NT}) + \gamma_4(\frac{J_{2NT}}{2NT}) = [\frac{\gamma_3-\gamma_1}{2N}J_{2N} + \gamma_1 I_{2N}]\otimes I_T + [\frac{\gamma_2-\gamma_1}{T}I_{2N} + \frac{\gamma_1-\gamma_2-\gamma_3+\gamma_4}{2NT}J_{2N}]\otimes J_T$. By assumption, it follows that $(\sigma^2_\epsilon;\sigma^2_e) > 0; \sigma^2_v = 0$; and thus $\gamma_1 = \gamma_2$ and $\gamma_3 = \gamma_4$, but $\gamma_1 \neq \gamma_3$. Therefore, $\Omega = [\gamma_1 I_{2N} + \frac{\gamma_3-\gamma_1}{2N}J_{2N}]\otimes I_T \equiv \Phi_1 \otimes I_T$. On the other hand, in the SUR approach, it follows that $\widehat{\Omega} = \widehat{\Sigma}\otimes I_T$, in which $\widehat{\Sigma}$ is the estimated $(2NT \times 2NT)$ matrix, with ij-th element given by $\widehat{\sigma}_{ij} = \frac{e_i' e_j}{T}$, and e_i is a $(T\times 1)$ vector containing the residuals of the i-th equation estimated by OLS. By assumption A3, it follows that $[\sigma^2_\epsilon I_{2N} + \sigma^2_e J_{2N} - \widehat{\Sigma}] > 0 \therefore [\gamma_1 I_{2N} + \frac{\gamma_3-\gamma_1}{2N}J_{2N} - \widehat{\Sigma}] > 0 \therefore (\Phi_1 - \widehat{\Sigma}) > 0 \therefore (\widehat{\Sigma}^{-1} - \Phi_1^{-1}) > 0 \therefore X'\{(\widehat{\Sigma}^{-1} - \Phi_1^{-1})\otimes I_T)\}X > 0 \therefore X'\{(\widehat{\Sigma}^{-1}\otimes I_T) - (\Phi_1^{-1}\otimes I_T)\}X > 0 \therefore X'\{(\widehat{\Sigma}\otimes I_T)^{-1} - (\Phi_1 \otimes I_T)^{-1}\}X > 0 \therefore \Big(X'(\widehat{\Sigma}\otimes I_T)^{-1}X\Big) - \big(X'(\Phi_1\otimes I_T)^{-1}X\big) > 0 \therefore \big(X'(\Phi_1\otimes I_T)^{-1}X\big)^{-1} - \Big(X'(\widehat{\Sigma}\otimes I_T)^{-1}X\Big)^{-1} > 0 \therefore \big(X'\Omega^{-1}X\big)^{-1} - \widehat{Var}(\beta_{SUR}) > 0 \therefore Var(\beta_{SUR}) - E[\widehat{Var}(\beta_{SUR}) \mid X] > 0 \therefore B_{SUR} > 0$.

Now, we must show that $(B_{OLS} - B_{SUR}) \geqslant 0$. First, define $C \equiv [Var(\beta_{ols}) - Var(\beta_{sur})]$ and $D \equiv [E(\widehat{Var}(\beta_{sur}) \mid X) - E(\widehat{Var}(\beta_{ols}) \mid X)$. In other words, we must prove that $(C+D) \geqslant 0$, i.e., $C \geqslant 0$ and $D \geqslant 0$. The first part is just the well-known result that the GLS estimator will in most cases be more efficient, and will never be less efficient than the OLS estimator. Regarding the second part, define $\lambda = \frac{2NT-k}{tr[M\Omega]}$. Thus, it follows that $D \equiv [E(\widehat{Var}(\beta_{sur}) \mid X) - E(\widehat{Var}(\beta_{ols}) \mid X) \geqslant 0 \Leftrightarrow \Big(X'(\widehat{\Sigma}\otimes I_T)^{-1}X\Big)^{-1} - (X'\lambda I X)^{-1} \geqslant 0 \Leftrightarrow (X'\lambda I X) - \Big(X'(\widehat{\Sigma}\otimes I_T)^{-1}X\Big) \geqslant 0 \Leftrightarrow X'[\lambda I - (\widehat{\Sigma}\otimes I_T)^{-1}]X \geqslant 0 \Leftrightarrow \lambda I - (\widehat{\Sigma}\otimes I_T)^{-1} \geqslant 0 \Leftrightarrow (\widehat{\Sigma}\otimes I_T) - \frac{1}{\lambda}(I_{2N}\otimes I_T) \geqslant 0 \Leftrightarrow (\widehat{\Sigma} - \frac{1}{\lambda}I_{2N})\otimes I_T \geqslant 0 \Leftrightarrow (\widehat{\Sigma} - \frac{1}{\lambda}I_{2N}) \geqslant 0$. By assumption A5, we have that $\widehat{\Sigma} \geqslant 0$ and $\Psi \geqslant \frac{2N(2NT-k)}{(tr((X'X)^{-1}X'JX)-2NT)+2N(2NT-k)}$. Now, if we define $\pi \equiv (tr(X(X'X)^{-1}X'J)-2NT)/(2NT-k)$, it follows that $(\pi+2N) = \frac{(tr(X(X'X)^{-1}X'J)-2NT)+2N(2NT-k)}{(2NT-k)} \therefore \frac{2N}{(\pi+2N)} = \frac{2N(2NT-k)}{(tr(X(X'X)^{-1}X'J)-2NT)+2N(2NT-k)}$ and, thus, $\Psi \geqslant \frac{2N}{(\pi+2N)}$. This way, $2N(\sigma^2_\epsilon + 2N\sigma^2_e) \leqslant 2N\sigma^2_e(\pi+2N) \therefore (\sigma^2_\epsilon + 2N\sigma^2_e) \leqslant \sigma^2_e(\pi+2N) \therefore \sigma^2_\epsilon + 2N\sigma^2_e \leqslant \pi\sigma^2_e + 2N\sigma^2_e \therefore \sigma^2_\epsilon \leqslant \pi\sigma^2_e \therefore \sigma^2_\epsilon(2NT-k) \leqslant \sigma^2_e(tr(X(X'X)^{-1}X'J) - 2NT) \therefore \sigma^2_\epsilon(2NT-k) + \sigma^2_e(2NT - tr(X(X'X)^{-1}X'J)) \leqslant 0 \therefore 2NT\sigma^2_\epsilon + \sigma^2_e 2NT - \sigma^2_\epsilon k - \sigma^2_e tr(X(X'X)^{-1}X'J) \leqslant 0 \therefore 2NT\gamma_1 + (\frac{\gamma_3-\gamma_1}{2N})tr(J) - \gamma_1 tr(X(X'X)^{-1}X') - (\frac{\gamma_3-\gamma_1}{2N})tr(X(X'X)^{-1}X'J) \leqslant 0 \therefore tr\{\gamma_1 I_{2NT} + \frac{\gamma_3-\gamma_1}{2N}J - X(X'X)^{-1}X'(\gamma_1 I_{2NT}) - X(X'X)^{-1}X'(\frac{\gamma_3-\gamma_1}{2N})J\} \leqslant 0 \therefore tr\{\gamma_1 I_{2N}\otimes I_T + \frac{\gamma_3-\gamma_1}{2N}J_{2N}\otimes I_T - X(X'X)^{-1}X'([\gamma_1 I_{2N}\otimes I_T] + [\frac{\gamma_3-\gamma_1}{2N}J_{2N}\otimes I_T)\} \leqslant 0 \therefore tr\{(\gamma_1 I_{2N} + \frac{\gamma_3-\gamma_1}{2N}J_{2N})\otimes I_T - X(X'X)^{-1}X'([\gamma_1 I_{2N} + \frac{\gamma_3-\gamma_1}{2N}J_{2N}]\otimes I_T)\} \leqslant 0 \therefore tr\{(I - X(X'X)^{-1}X')([\gamma_1 I_{2N} + \frac{\gamma_3-\gamma_1}{2N}J_{2N}]\otimes I_T)\} \leqslant 0 \therefore tr[M\Omega] \leqslant 0$. Therefore, by A2(ii), $T > 2p \therefore (2NT-k) > 0$, and, thus, $\lambda = \frac{2NT-k}{tr[M\Omega]} \leqslant 0 \therefore (\widehat{\Sigma} - \frac{1}{\lambda}I_{2N}) \geqslant 0 \therefore (C+D) \geqslant 0$;

(iii) By assumption A4, it follows that $\sigma^2_\epsilon I_{2N} + \sigma^2_e J_{2N} = \widehat{\Sigma}$ or $\Phi_1 = \widehat{\Sigma}$; where $\Phi_1 \equiv \sigma^2_\epsilon I_{2N} + \sigma^2_e J_{2N}$. This way, $(\Phi_1 \otimes I_T) = (\widehat{\Sigma}\otimes I_T) \therefore X'(\Phi_1\otimes I_T)^{-1}X = X'(\widehat{\Sigma}\otimes I_T)^{-1}X \therefore E[(X'(\Phi_1\otimes I_T)^{-1}X)^{-1} \mid X] = E[(X'(\widehat{\Sigma}\otimes I_T)^{-1}X)^{-1} \mid X] \therefore (X'(\Phi_1\otimes I_T)^{-1}X)^{-1} = E[(X'(\widehat{\Sigma}\otimes I_T)^{-1}X)^{-1} \mid X] \therefore Var(\beta_{SUR}) =$

$E[\widehat{Var}(\beta_{SUR}) \mid X] \therefore B_{SUR} = 0$;

(iv) If $\sigma_v^2 = \sigma_e^2 = 0$, then, $\gamma_1 = \gamma_2 = \gamma_3 = \gamma_4 = \sigma_\epsilon^2$; and the covariance matrix is simplified to $\Omega = \sigma_\epsilon^2(I_{2NT})$, which is a diagonal matrix. By assumption A4, it follows that $\widehat{\Sigma} = \sigma_\epsilon^2 I_{2N}$, which is a diagonal matrix. According to Wooldridge (2002, Theorem 7.5), if $\widehat{\Omega} = \widehat{\Sigma} \otimes I_T$ is a diagonal matrix, then OLS equation by equation is identical to Feasible GLS. In addition, it follows that $(\sigma_\epsilon^2 I_{2N} \otimes I_T) = (\widehat{\Sigma} \otimes I_T)$ $\therefore X'(\sigma_\epsilon^2 I_{2N} \otimes I_T)^{-1}X = X'(\widehat{\Sigma} \otimes I_T)^{-1}X \therefore E[(X'(\sigma_\epsilon^2(I_{2NT}))^{-1}X)^{-1} \mid X] = E[(X'(\widehat{\Sigma} \otimes I_T)^{-1}X)^{-1} \mid X]$ $\therefore (X'\Omega^{-1}X)^{-1} = E[(X'(\widehat{\Sigma} \otimes I_T)^{-1}X)^{-1} \mid X] \therefore Var(\beta_{SUR}) = E[\widehat{Var}(\beta_{SUR}) \mid X] \therefore B_{SUR} = 0$. ■

Proof of Proposition 4. (ia) Following Greene (2003, p. 207), since $\Omega > 0$ and $\Omega' = \Omega$, then, it can be factored into $\Omega = C\Lambda C'$, where the columns of C are the eigenvectors of Ω, and Λ is a diagonal matrix with the eigenvalues of Ω. Let $\Lambda^{1/2}$ be a diagonal matrix with i-th diagonal element $\sqrt{\lambda_i}$, and let $T = C\Lambda^{1/2}$. Therefore, $\Omega = TT'$; and if we define $P' = C\Lambda^{-1/2}$, then, $\Omega^{-1} = P'P$. This way, pre-multiplying the considered model $Y = X\beta + \varepsilon$ by P, it follows that $PY = PX\beta + P\varepsilon$ or $Y_* = X_*\beta + \varepsilon_*$, in which $E(\varepsilon_*\varepsilon_*' \mid X) = P\Omega P'$. By considering (at this point) that Ω is known, it follows that Y_* and X_* are observed data. The FB estimator of the transformed model is given by $\widehat{\beta}_{FB} = (X'\Omega^{-1}X)^{-1}X'\Omega^{-1}Y = (X_*'X_*)^{-1}X_*'Y_* = (X_*'X_*)^{-1}X_*'(X_*\beta + \varepsilon_*) = \beta + (X_*'X_*)^{-1}X_*'\varepsilon_*$. Thus, $E(\widehat{\beta}_{FB} \mid X_*) = \beta + E[(X_*'X_*)^{-1}X_*'\varepsilon_* \mid X_*] = \beta$ if $E(\varepsilon_* \mid X_*) = E(P\varepsilon \mid PX) = 0$. However, note that since P is a matrix of known constants, we in fact just required that $E(\varepsilon \mid X) = 0$. By applying the error decomposition $\varepsilon_{i,t} = v_i + e_t + \epsilon_{i,t}$, it follows that $E(\varepsilon_{i,t} \mid X) = E(v_i \mid X) + E(e_t \mid X) + E(\epsilon_{i,t} \mid X)$. Now, recall that vector X is composed of lagged values of $\Delta Z_{i,t}$ and $CA_{i,t}$, which could generate a correlation between $\varepsilon_{i,t}$ and these lagged variables through the v_i random effect. However, since $E(v_i) = 0$ and, by assumption $\sigma_v^2 = 0$, it follows that $E(\varepsilon \mid X) = 0$ indeed holds, since e_t and $\epsilon_{i,t}$ are assumed to be uncorrelated with the regressors;

(ib) By considering $T \to \infty$ with fixed N, note that $plim(\widehat{\beta}_{FB}) = \beta + plim((X_*'X_*)^{-1}X_*'\varepsilon_*) = \beta + plim((\frac{X'\Omega^{-1}X}{T})^{-1}\frac{X'\Omega^{-1}\varepsilon}{T}) = \beta + Q_*^{-1}plim(\frac{X'\Omega^{-1}\varepsilon}{T})$, since by assumption A6 we have that $plim[(1/T)X'\Omega^{-1}X]^{-1} = Q_*^{-1}$. Thus, in order to obtain consistency, and apply the product rule of White (1984, Lemma 4.6), we need to show that $plim(\frac{X'\Omega^{-1}\varepsilon}{T}) = 0$. Following Greene (2003, p. 67), let $\frac{X'\Omega^{-1}\varepsilon}{T} = \frac{X_*'\varepsilon_*}{T} = \frac{1}{T}\sum_{t=1}^{T} w_t \equiv \overline{w}$, which is a $k \times 1$ vector; where $w_t \equiv \{\phi(L)(\Delta Z_t^i + CA_t^i)\sum_{j=1}^{2NT} s_{i,j}\varepsilon_{j,t}\}$; $\phi(L)$ is a lag polynomial of order p; and $s_{i,j}$ is the ij-th element of Ω^{-1}, which is a $(2NT \times 2NT)$ matrix. This way, $plim(\widehat{\beta}_{FB}) = \beta + Q_*^{-1}plim(\overline{w})$. Now, note that $E[w_t] = E(E[w_t \mid X]) = E(E[\{\phi(L)(\Delta Z_t^i + CA_t^i)\sum_{j=1}^{2NT} s_{i,j}\varepsilon_{j,t}\} \mid X]) = E(\phi(L)(\Delta Z_t^i + CA_t^i)\sum_{j=1}^{2NT} s_{i,j}E[\varepsilon_{j,t} \mid X]) = 0$, since from the exogeneity assumptions, it follows that $E(\varepsilon_{j,t} \mid X) = 0; \forall j, t \therefore E[w_t] = 0$ and, thus, $E[\overline{w}] = 0$. On the other hand, note that $Var(\overline{w}) = E(Var(\overline{w} \mid X)) + Var(E(\overline{w} \mid X))$, where the second term is zero, since $E(\varepsilon_{i,t} \mid X) = 0$ and, thus, $E(\overline{w} \mid X) = \frac{1}{T}\sum_{t=1}^{T} E(w_t \mid X) = \frac{1}{T}\sum_{t=1}^{T} E(\{\phi(L)(\Delta Z_t^i + CA_t^i)\sum_{j=1}^{2NT} s_{i,j}\varepsilon_{j,t}\} \mid X) = \frac{1}{T}\sum_{t=1}^{T}\phi(L)(\Delta Z_t^i + CA_t^i)\sum_{j=1}^{2NT} s_{i,j}E(\varepsilon_{j,t} \mid X) = 0$. Now, note that $Var(\overline{w} \mid X) = E(\overline{w}\overline{w}' \mid X)$ $= E[(\frac{1}{T}\sum_{t=1}^{T} w_t)(\frac{1}{T}\sum_{t=1}^{T} w_t') \mid X] = E[(\frac{1}{T}X_*'\varepsilon_*)(\frac{1}{T}\varepsilon_*'X_*) \mid X] = \frac{X_*'}{T}E(\varepsilon_*\varepsilon_*' \mid X_*)\frac{X_*}{T} = \frac{X_*'}{T}E(P\varepsilon\varepsilon'P' \mid X)\frac{X_*}{T} = \frac{X_*'}{T}PE(\varepsilon\varepsilon' \mid X)P'\frac{X_*}{T} = \frac{X_*'}{T}P\Omega P'\frac{X_*}{T} = \frac{X'}{T}P'P\Omega P'P\frac{X}{T} = \frac{X'}{T}\Omega^{-1}\Omega\Omega^{-1}\frac{X}{T} = \frac{X'}{T}\Omega^{-1}\frac{X}{T}$. Thus, it follows that

$Var(\overline{w}) = E(Var(\overline{w} \mid X)) = E(\frac{X'}{T}\Omega^{-1}\frac{X}{T} \mid X) = \frac{1}{T}E(\frac{X'\Omega^{-1}X}{T} \mid X)$. The variance of $\overline{w}$ will collapse to zero if the conditional expectation in parentheses converges to a constant matrix, so that the leading scalar $(1/T)$ will dominate the product as T increases. Thus, by considering assumption A6 (and a WLLN), it follows that $plim[Var(\overline{w})] = plim(\frac{1}{T})Q_* = 0$. Therefore, since $E[\overline{w}] = 0$ and $plim[Var(\overline{w})] = 0 \therefore plim(\overline{w}) = 0$ and, thus, $plim(\widehat{\beta}_{FB}) = \beta$;

(ii) Given that $\sigma_v^2 = 0$ and assumption A7 holds, it follows that $\varepsilon_{i,t} = e_t + \epsilon_{i,t}$ is also normally distributed, i.e., $\varepsilon \sim N(0, \Omega)$, where $E(\varepsilon\varepsilon') = \Omega$ is finite and nonsingular. Therefore, by applying White (1984, Theorem 1.3), the result is straightforward: $\widehat{\beta}_{FB} \sim N(\beta; (X'\Omega^{-1}X)^{-1})$;

(iii) By assumptions A6-A7, and Theorem 1 of Fuller & Battese (1974), it follows that (when $T \to \infty$ and N is fixed), $\widehat{\sigma}_e^2$ and $\widehat{\sigma}_\epsilon^2$ are consistent estimators of σ_e^2 and σ_ϵ^2 respectively. Thus, $\widehat{\Omega} = \widehat{\sigma}_\epsilon^2 I_{2NT} + \widehat{\sigma}_e^2(J_{2N} \otimes I_T)$ is asymptotically equivalent to Ω. Based on assumption A8, we can apply the asymptotic equivalence Lemma 4.7 of White (1984), in which $\widehat{\widehat{\beta}}_{FB}$ is asymptotically equivalent to $\widehat{\beta}_{FB}$. See also Greene (2003, p. 210) for further details. ■

Appendix C. Further results of the empirical exercise

Table 10 - ADF Unit Root test

	USA	CAN	JPN	UK	GER	ITA	FRA
C_t	-0.99	-2.19	-1.13	-1.17	-2.43	-2.04	-1.32
ΔC_t	-5.44 (**)	-14.02 (**)	-16.65 (**)	-15.76 (**)	-15.25 (**)	-9.78 (**)	-16.18 (**)
GNI_t^*	-1.51	-1.23	-1.85	-1.34	-1.15	-0.52	-2.14
ΔGNI_t^*	-6.08 (**)	-16.74 (**)	-14.58 (**)	-20.10 (**)	-12.01 (**)	-5.58 (**)	-15.62 (**)
CA_t	-3.20 (*)	-2.58 (+)	-3.21 (*)	-2.82 (+)	-2.02	-1.87	-1.24
ΔZ_t	-14.29 (**)	-15.41 (**)	-14.68 (**)	-18.61 (**)	-13.35 (**)	-14.11 (**)	-17.32 (**)

Notes: a) $GNI_t^* \equiv Y_t + rB_t - I_t - G_t$.

b) (**) indicates rejection of the null hypothesis of unit root at 1% level; (*) at 5% level and (+) at 10% level

c) $CA_t = Y_t + rB_t - I_t - G_t - \theta C_t$

d) The USA, CAN, JPN, UK and GER series range from 1960:1-2007:1, whereas ITA and FRA range from 1980:1-2007:1.

Table 11 - Johansen´s Cointegration Test

The cointegration vector between $(Y_t+rB_t - I_t - G_t)$ and C_t is given by $(1, -\theta)$

		Ho:	p=0	Ho:	p≤1
Country	θ	trace stat .	$\lambda_{máx.}$ stat.	trace stat.	$\lambda_{máx.}$ stat.
USA	0.895 (0.018)	13.72	10.63	3.09	3.09
CAN	1.024 (0.049)	9.54	8.41	1.13	1.13
JPN	1.105 (0.016)	15.06	12.83	2.23	2.23
UK	0.921 (0.026)	13.59	10.89	2.70	2.70
GER	1.114 (0.077)	5.61	4.05	1.56	1.56
ITA	1.120 (0.058)	13.16	9.23	3.93 (*)	3.93 (*)
FRA	1.267 (0.149)	4.64	4.60	0.03	0.03

Notes: a) p is the number of cointegrating relations; (*) indicates rejection of Ho at 5% level.

b) In column θ the standard deviation is presented in parentheses.

c) The USA, CAN, JPN, UK and GER series range from 1960:1-2007:1, whereas ITA and FRA range from 1980:1-2007:1.

Table 12 - Comparison of θ with the literature

Country	IG	Ghosh (95)	Agénor et al.(99)
USA	0.895	0.994	-
CAN	1.024	0.96	-
JPN	1.105	1.04	-
UK	0.921	0.98	-
GER	1.114	1.08	-
ITA	1.120	-	-
FRA	1.267	-	0.982

Note: IG means Issler & Gaglianone (our results).

Table 13 - Wald test (p-value) - Per capita time series, in 2000 U.S. dollars

Country	OLS	SUR	FB.1	FB.2
$USA_{(1)}$	0.1369	0.0201 (*)	0.0351 (*)	0.0586
$CAN_{(1)}$	0.3238	0.1430	0.0620	0.0734
$JPN_{(1)}$	0.0042 (**)	0.0001 (**)	0.0002 (**)	0.0004 (**)
$UK_{(1)}$	0.1685	0.1049	0.7692	0.7088
$GER_{(1)}$	0.1363	0.2351	0.6118	0.6690
$ITA_{(1)}$	0.1596	0.0301 (*)	0.9850	0.9652
$FRA_{(1)}$	0.0078 (**)	0.0000 (**)	0.0149 (*)	0.0139 (*)
$USA_{(2)}$	0.0256 (*)	0.0039 (**)	0.0120 (*)	0.0158 (*)
$CAN_{(2)}$	0.1932	0.0531	0.0509	0.2274
$JPN_{(2)}$	0.0015 (**)	0.0001 (**)	0.0000 (**)	0.0019 (**)
$UK_{(2)}$	0.0078 (**)	0.0001 (**)	0.5571	0.2576
$GER_{(2)}$	0.0174 (*)	0.0096 (**)	0.4876	0.1167

Notes: a) Ho: $(CA_t^* - \mu_{CA}^*) = (CA_t - \mu_{CA})$

b) (**) means rejection at 1% level, (*) at 5% level;

c) $USA_{(1)}$ indicates sample 1 (1980:1-2007:1), and $USA_{(2)}$ means sample 2 (1960:1-2007:1);

d) FB.1 means Fuller & Battese (1974) error decomposition, with a unique global shock,
whereas FB.2 considers distinct global shocks for CA_t and ΔZ_t equations.

Table 14 - Wald test (p-value) - Time series in U.S. dollars (not per capita)

Country	OLS	SUR	FB.1	FB.2
$USA_{(1)}$	0.1400	0.0207 (*)	0.0340 (*)	0.0394 (*)
$CAN_{(1)}$	0.2402	0.0312 (*)	0.8692	0.8766
$JPN_{(1)}$	0.0291 (*)	0.0205 (*)	0.0032 (**)	0.0051 (**)
$UK_{(1)}$	0.1796	0.1550	0.9970	0.9979
$GER_{(1)}$	0.0480 (*)	0.0097 (**)	0.6667	0.6898
$ITA_{(1)}$	0.1425	0.0789	0.9771	0.9594
$FRA_{(1)}$	0.0295 (*)	0.0000 (**)	0.7957	0.7525
$USA_{(2)}$	0.0291 (*)	0.0069 (**)	0.0000 (**)	0.0000 (**)
$CAN_{(2)}$	0.1499	0.0123 (*)	0.8288	0.8361
$JPN_{(2)}$	0.0201 (*)	0.0160 (*)	0.0363 (*)	0.3019
$UK_{(2)}$	0.0139 (*)	0.0001 (**)	0.8727	0.6861
$GER_{(2)}$	0.0143 (*)	0.0078 (**)	0.7161	0.3260

Notes: a) Ho: $(CA_t^* - \mu_{CA}^*) = (CA_t - \mu_{CA})$

b) (**) means rejection at 1% level, (*) at 5% level;

c) $USA_{(1)}$ indicates sample 1 (1980:1-2007:1), and $USA_{(2)}$ means sample 2 (1960:1-2007:1);

d) FB.1 means Fuller & Battese (1974) error decomposition, with a unique global shock,
whereas FB.2 considers distinct global shocks for CA_t and ΔZ_t equations.

Appendix D. Some details of the Monte Carlo simulation

Generating a covariance-stationary VAR

An initial idea to design the Monte Carlo experiment could consist on constructing the companion matrix, sorting it values from uniform distributions, in order to satisfy the restrictions imposed by the null hypothesis, and then verifying whether or not the eigenvalues of the companion matrix all lie inside the unit circle. However, this strategy could lead to a wide spectrum of search for adequate values for the companion matrix. This way, we propose an analytical solution to generate a covariance-stationary VAR, based initially on the choice of the eigenvalues, and then on the generation of the respective companion matrix. According to Hamilton (1994, page 259), if the eigenvalues of the companion matrix (F) all lie inside the unit circle, then the VAR turns out to be covariance-stationary. The eigenvalue vector (λ) of the companion matrix (F) for a VAR(p) is obtained from the following equation:

$$|F - \lambda I| = 0, \tag{39}$$

where I is the identity matrix. Two important properties of the eigenvalue vector, presented in Simon&Blume (1994, page 599), are reproduced below:

$$trace(F) = \sum_{i=1}^{n} \lambda_i \quad ; \quad \det(F) = \prod_{i=1}^{n} \lambda_i, \tag{40}$$

where n is the number of eigenvalues (equal to $2p$). The companion matrix for a VAR(1) is given by

$$F = \begin{bmatrix} a & b \\ c & d \end{bmatrix}. \tag{41}$$

The restrictions imposed to a VAR(1) to consider the optimal current account equal to the observed one (i.e., a true null hypothesis) are given by $c = a$; and $d = b + (1 + r)$. This way, in order to impose a false null hypothesis into the model we multiply the restrictions above by a gamma factor, resulting in the following companion matrix

$$F = \begin{bmatrix} a & b \\ \gamma a & \gamma(b + (1 + r)) \end{bmatrix}. \tag{42}$$

It should be mentioned that setting the gamma factor equal to unity we consider a true Ho, but imposing gamma less than unity we generate a false null hypothesis. This way, the eigenvalues of the companion matrix are given by

$$\det(F) = \gamma a(b + 1 + r) - \gamma ab = \lambda_1 \lambda_2 \quad \therefore \quad a = \frac{\lambda_1 \lambda_2}{\gamma(1 + r)} \tag{43}$$

$$trace(F) = a + \gamma(b + (1 + r)) = \lambda_1 + \lambda_2 \quad \therefore \quad b = \frac{\lambda_1 + \lambda_2 - a - \gamma(1 + r)}{\gamma}. \tag{44}$$

Therefore, to construct a covariance-stationary VAR(1) we sort from a uniform distribution (-1;1) the values of λ_1 and λ_2 and calculate the parameters a and b from the equations above. In a VAR(2) case, the companion matrix (considering the restrictions implied by the null hypothesis) is given by

$$F = \begin{bmatrix} a_1 & a_2 & b_1 & b_2 \\ 1 & 0 & 0 & 0 \\ \gamma a_1 & \gamma a_2 & \gamma(b_1 + (1+r)) & \gamma b_2 \\ 0 & 0 & 1 & 0 \end{bmatrix}. \tag{45}$$

The eigenvalue vector is obtained from

$$|F - \lambda I| = \begin{vmatrix} a_1 - \lambda & a_2 & b_1 & b_2 \\ 1 & -\lambda & 0 & 0 \\ \gamma a_1 & \gamma a_2 & \gamma(b_1 + (1+r)) - \lambda & \gamma b_2 \\ 0 & 0 & 1 & -\lambda \end{vmatrix} = 0. \tag{46}$$

Since $\det(F) = 0$, at least one eigenvalue is null ($\lambda_1 = 0$). Thus, from the last equation, it follows that

$$|F - \lambda I| = \lambda^4 - \lambda^3(a_1 + \gamma b_1 + \gamma(1+r)) + \lambda^2(\gamma a_1(1+r) - a_2 - \gamma b_2) + \lambda\gamma a_2(1+r) = 0 \tag{47}$$

$$\therefore a_2 = \frac{\lambda(\gamma a_1(1+r) - \gamma b_2 - \lambda(a_1 + \gamma b_1 + \gamma(1+r)) + \lambda^2)}{\lambda - \gamma(1+r)}. \tag{48}$$

The equation above must be valid for all values of λ_i. In particular, one can construct a system of 2 equations, with the expression above, setting $\lambda = \lambda_2$ and $\lambda = \lambda_3$. This way, we can explicit b_2 as a function of a_1, b_1, r, λ_2 and λ_3, as it follows:

$$\begin{aligned} b_2 = {} & ((-\lambda_2\gamma^2 + \gamma^2 a_1 - \lambda_3\gamma^2)r^2 + (\lambda_2^2\gamma + 2\lambda_2\gamma\lambda_3 - \lambda_2\gamma^2 b_1 - \lambda_3\gamma^2 b_1 - a_1\lambda_2\gamma - 2\lambda_2\gamma^2 + 2\gamma^2 a_1 \\ & -a_1\lambda_3\gamma + \lambda_3^2\gamma - 2\lambda_3\gamma^2)r + 2\lambda_2\gamma\lambda_3 - \lambda_2^2\lambda_3 + \lambda_2\gamma b_1\lambda_3 + a_1\lambda_2\lambda_3 - \lambda_3^2\lambda_2 + \lambda_2^2\gamma - \lambda_2\gamma^2 b_1 \\ & -\lambda_3\gamma^2 b_1 - a_1\lambda_2\gamma + \lambda_3^2\gamma - \lambda_3\gamma^2 - a_1\lambda_3\gamma + \gamma^2 a_1 - \lambda_2\gamma^2)/((1+r)\gamma^2). \end{aligned} \tag{49}$$

On the other hand, the trace of the companion matrix is given by

$$trace(F) = a_1 + \gamma(b_1 + (1+r)) = \lambda_2 + \lambda_3 + \lambda_4 \tag{50}$$

$$\therefore b_1 = \frac{(\lambda_2 + \lambda_3 + \lambda_4) - a_1 - \gamma(1+r)}{\gamma}. \tag{51}$$

Therefore, to construct a covariance-stationary VAR(2) we set $\lambda_1 = 0$ and sort (independently) from uniform distributions (-1;1) the values of a_1, λ_2, λ_3 and λ_4. Then, we choose a gamma factor in order to simulate the false (or true) null hypothesis. Thus, we obtain b_1 by the last expression and calculate b_2 and a_2 from the previous equations. This way, we construct the companion matrix with all eigenvalues inside the unit circle. A VAR(p) can be constructed in a similar way by following the presented methodology.

Besides the proper construction of the companion matrix, another important issue of the Monte Carlo simulation is the generation of the residuals for the VAR. Since the current accounts for different countries are nowadays expected to be globally linked, we construct the residuals of the VAR based on two components: an idiosyncratic shock, and a global shock, which is assumed to be common among the considered countries. The construction of these residuals is detailed in next section.

Constructing Global Shocks

A key issue regarding the Monte Carlo experiment is the generation of the residuals for the VAR. Inspired by Glick & Rogoff (1995), which study the current account response to different productivity shocks in the G-7 with a structural model including global and country-specific shocks, and based on the 'common sense' that the current account fluctuations have become more closely linked across countries in the last decades, we decompose the residuals of the VAR into an idiosyncratic shock and a global (common) component among countries. Thus, the VAR(p) for a country i can be written as:

$$\begin{bmatrix} \Delta Z_t^i \\ CA_t^i \end{bmatrix} = \begin{bmatrix} a^i(L) & b^i(L) \\ c^i(L) & d^i(L) \end{bmatrix} \times \begin{bmatrix} \Delta Z_{t-1}^i \\ CA_{t-1}^i \end{bmatrix} + \begin{bmatrix} \varepsilon_{1t}^i \\ \varepsilon_{2t}^i \end{bmatrix}, \tag{52}$$

where the residuals ε_t^i can be decomposed into an idiosyncratic shock (ν_t^i) and a global (common) component among countries (ξ_t), as it follows:

$$\begin{bmatrix} \varepsilon_{1t}^i \\ \varepsilon_{2t}^i \end{bmatrix} = \begin{bmatrix} \nu_{1t}^i \\ \nu_{2t}^i \end{bmatrix} + scale * \begin{bmatrix} \xi_{1t} \\ \xi_{2t} \end{bmatrix}, \tag{53}$$

or in a reduced form: $\varepsilon_t^i = \nu_t^i + scale * \xi_t$. The parameter *scale* is used to measure the importance of the global shock into the residuals of a country i. The numerical procedure adopted to construct the residuals in the Monte Carlo experiment first drops (independently) from a normal standard distribution the ν_{1t}^i and ν_{2t}^i series of shocks, resulting on a I_2 covariance matrix:

$$\Sigma^i = Var(\nu_t^i) = \begin{bmatrix} 1 & 0 \\ 0 & 1 \end{bmatrix}. \tag{54}$$

However, this covariance matrix does not represent a bivariate shock, and we must transform the covariance matrix Σ^i into a covariance matrix Ω^i, in a framework of a bivariate normal distribution, with parameters as close as possible to the data of the countries.

$$\Omega^i = \begin{bmatrix} \sigma_{1i}^2 & r^i \\ r^i & \sigma_{2i}^2 \end{bmatrix} \tag{55}$$

Thus, we must find a symmetric matrix X to make the following transformation: $\Omega = X\Sigma X'$. In our case, Σ is a diagonal and symmetric matrix, as it follows:

$$\therefore \Omega = X\Sigma^{1/2}\Sigma^{1/2}X' = \Sigma^{1/2}XX\Sigma^{1/2} \therefore XX = \Sigma^{-1/2}\Omega\Sigma^{-1/2}. \tag{56}$$

To obtain X we must calculate the square root of the matrix (XX). Adopting the eigenvalue decomposition, according to Simon&Blume (1994),[26] one could rewrite the (XX) matrix as a function of the eigenvectors (V) and the eigenvalue matrix (D), as it follows: $XX = VDV^{-1} = VD^{1/2}D^{1/2}V^{-1} = (VD^{1/2}V^{-1})(VD^{1/2}V^{-1})$, where D is a diagonal matrix filled with the eigenvalues of (XX). This way, the matrix X can be obtained

[26] For further details see Simon&Blume (1994), pages 590-595, and page 866 of Ruud (2000).

by $(VD^{1/2}V^{-1})$. Finally, after calculating X, one could use the constructed vector of residuals ν_t^i to obtain the $\widetilde{\nu}_t^i$ residuals (following a bivariate normal distribution with covariance matrix Ω^i):

$$\underset{\text{Tx2}}{\widetilde{\nu}_t} = \underset{\text{Tx2}}{\nu_t} \underset{\text{2x2}}{X} \tag{57}$$

where $Var(\widetilde{\nu}_t^i) = \Omega^i$, $Var(\nu_t^i) = \Sigma^i$, and T is the number of observations. The next step is to construct the global shocks, common to all countries. The procedure adopted in the Monte Carlo experiment drops (independently) from a normal standard distribution the ξ_{1t} and ξ_{2t} series of shocks, resulting in a covariance matrix equal to an I_2 matrix. However, this covariance matrix also does not represent a bivariate shock, and must be transformed, in the same way presented above (adopting the eigenvalue decomposition), into a covariance matrix Γ, representing a bivariate normal distribution:

$$\Gamma = Var(\xi_t) = \begin{bmatrix} 1 & w \\ w & 1 \end{bmatrix}, \tag{58}$$

where the parameter w represents the covariance between the global shocks on ΔZ_t^i and CA_t^i equations.

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