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MATHIAS HOFFMANN

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Abstract

We propose a simple intertemporal model of output and current account dynamics that we estimate using a cointegrated VAR approach. We suggest a method for identifying global and country-specific shocks from the VAR and test it, using cross-country evidence. Our results show that the identification scheme works well in practice, corroborating an important prediction of the intertemporal approach to the current account. We associate global shocks with movements in the US output growth rate. In accordance with the theory, we also observe a link between the global shock and a measure of the world real interest rate. This link is more pronounced in the long-run than in the short-run.

JEL CLASSIFICATION: F41, F43, C32

1 Introduction

Little stylized knowledge is available on the question in which way industrialized countries are prone to international shocks and how they adjust to them. In this paper, we propose a simple model centered around the current account as the key variable of macroeconomic transmission. Our setup offers a compact framework in which the following questions can be tackled:

- Can we validly identify global and country-specific shocks using a simple model of the world economy?
- How persistent are global and country-specific shocks?
- Can we associate global shocks with observable economic variables?
- What drives the development of long-run output in the seven biggest economies? Is it global shocks or country-specific shocks? Do shocks to the current account drive output or do output shocks determine the current account?

The theoretical framework of the paper is provided by the intertemporal approach to the current account initiated by Sachs (1981) and extended by Obstfeld (1986, 1995). Since the appearance of the landmark book by Obstfeld and Rogoff (1996), the intertemporal approach has also become a textbook paradigm. Our empirical implementation relies on a structural VAR approach that is embedded in a cointegrated model. We think that such a framework is a good vehicle with which to fish for stylized facts in international macro: it contains enough economics to avoid the risk of 'measurement without theory' but is at the same time simple and data-driven.

The paper's layout is as follows: section two presents a simple intertemporal optimisation model of the current account that highlights the econometric implications of the intertemporal approach and suggests how permanent and transitory components of output can be identified. In

Section 3, we suggest an identification scheme to identify country-specific and global shocks and discuss its econometric implementation. In Section 4, we present results; in particular, we discuss the quality of our identification scheme, using cross-country evidence. Section 5 concludes.

2 The intertemporal approach

In our empirical implementation, we will use expected utility, which is quadratic in consumption, in an intertemporal setting: i.e. the representative consumer maximizes

$$E_t \sum_{i=0}^{\infty} \left(\frac{1}{1+r} \right)^i \left[C_{t+i} - \frac{h}{2} C_{t+i}^2 \right] \tag{1}$$

subject to the intertemporal budget constraint

$$B_{t+1} = (1+r)B_t + Y_t - C_t \tag{2}$$

where Y_t is output, C_t is consumption and r represents the world real interest rate. B_t denotes the stock of net foreign assets which is required to be non-explosive:

$$\lim_{i \to \infty} B_{t+i}(1+r)^{-i} = 0 \tag{3}$$

The current account is defined as[1]

$$CA_t = \Delta B_{t+1} \tag{4}$$

In such a model agents behave as if all variables actually realize their expected values.

[1]In this model, a change in the net foreign asset position, B_t, will require an international flow of funds. The current account is more generally defined as the difference between savings and investment, $CA = S - I$ and of course that is the case here as well once we define $S_t = Y_t - C_t + rB_t$. The equality between CA_t and ΔB_{t+1}, will hold only under the assumption that no price changes affect the country's net foreign asset position. This would, e.g., happen whenever the real exchange rate changes.

This certainty-equivalence feature yields a simple forward looking solution for the consumption function:

$$C_t = \frac{r}{1+r}\left[(1+r)B_t + \sum_{s=0}^{\infty}\left(\frac{1}{1+r}\right)^s E_t Y_{t+s}\right]$$

Plugging this into the definition of the current account, we get

$$CA_t = Y_t - \frac{r}{1+r}\sum_{s=0}^{\infty}\left(\frac{1}{1+r}\right)^s E_t Y_{t+s} = Y_t - \tilde{Y}_t \tag{5}$$

where $\tilde{Y}_t$ denotes the permanent value of output.

Now let us specify a simple process for output:

$$Y_t = Y_{t-1} + \sum_{i=0}^{\infty} \mathbf{c}_i' \mathbf{e}_{t-i} \tag{6}$$

Here, $\mathbf{e}_t = [\ e_t^c,\ \ e_t^w\ ]'$ denotes the vector of country-specific and global shocks which are assumed to have unit variance and are serially and contemporaneously uncorrelated.

We can rewrite equation (5) to yield:

$$CA_t = -\sum_{s=1}^{\infty}\left(\frac{1}{1+r}\right)^s E_t \Delta Y_{t+s} \tag{7}$$

Then, from (6) we get

$$E_t \Delta Y_{t+s} = \sum_{i=0}^{\infty} \mathbf{c}_{i+s}' \mathbf{e}_{t-i}$$

Plugging this into (7) yields:

$$CA_t = -\sum_{s=1}^{\infty}\left(\frac{1}{1+r}\right)^s \sum_{i=0}^{\infty} \mathbf{c}_{i+s}' \mathbf{e}_{t-i} = -\sum_{i=0}^{\infty} \mathbf{d}_i' \mathbf{e}_{t-i}$$

where $\mathbf{d}_i' = \sum_{s=1}^{\infty}\left(\frac{1}{1+r}\right)^s \mathbf{c}_{i+s}'$.

The above setup gives us a simple joint representation of current account and output in differences:

$$\begin{bmatrix} \Delta CA_t \\ \Delta Y_t \end{bmatrix} = \begin{bmatrix} (\mathbf{1}-\mathbf{L})\mathbf{d}'(\mathbf{L}) \\ \mathbf{c}'(\mathbf{L}) \end{bmatrix} \mathbf{e}_t = \mathbf{D}(\mathbf{L})\mathbf{e}_t \tag{8}$$

Note that in this structural moving-average representation, the dynamics of the current account are driven by global and country-specific shocks. If however, international capital mobility is sufficiently high, all countries will react to a global shock in the same way - wanting to save more or less, depending on which way the shock goes. But not all can run current account deficits or surpluses at the same time. Rather, a global shock should then impinge on the world interest rate and equilibrate world saving and investment.

This reasoning has two implications:

- The current account should react more strongly to country-specific shocks than to global shocks.
- global shocks should be associated with changes in the world interest rates.

In the sequel of the paper, we will use the first of these two implications to identify country-specific and global shocks. The quality of this identification is then assessed using the second.

3 Econometric Implementation

In the structural model (8), both variables are stationary. In this paper, however, we are concerned with the long-run properties of output, i.e. with its permanent component. We will therefore consider a system in the level of output and the current account:

$$X_t' = [\ CA, \ \ Y_t\ ] \tag{9}$$

In such a system, output is $I(1)$ whereas the current account is stationary. This amounts to saying that the two variables share one common trend or in other words, there is a trivial cointegrating relationship with cointegrating vector $\beta' = \begin{bmatrix} 1, & 0 \end{bmatrix}$. This, becomes clearer once we express X_t in terms of a (structural) Beveridge-Nelson (1981)/Stock-Watson (1988) representation:

$$\mathbf{X}_t = \mathbf{D(1)} \sum_{i=0}^{t} \mathbf{e}_i + \mathbf{D}^*(\mathbf{L})\mathbf{e}_t \tag{10}$$

where $\mathbf{D}_i^* = -\sum_{l=i+1}^{\infty} \mathbf{D}_l$ and $\mathbf{D(1)} = \sum_{i=1}^{\infty} \mathbf{D}_i$.

Because CA is stationary, we have $\mathbf{d}'(\mathbf{1}) = 0$ and therefore

$$\mathbf{D(1)} = \begin{bmatrix} \mathbf{d}'(\mathbf{1}) \\ \mathbf{c}'(\mathbf{1}) \end{bmatrix} = \begin{bmatrix} 0 & 0 \\ c_{CA}(1) & c_Y(1) \end{bmatrix}$$

Hence, $\mathbf{D(1)}$ has reduced rank and the long-run dynamics of the system are driven by the stochastic trend $\mathbf{c}'(\mathbf{1}) \sum_{i=0}^{t} \mathbf{e}_i$.

The structural shocks are unobservable and therefore the moving average-representation of $\mathbf{\Delta X}_t$ or the BN-representation for $\mathbf{X}_t$ cannot be estimated directly. Rather, we assume that it is possible to estimate a reduced-form moving average

$$\mathbf{\Delta X}_t = \mathbf{C(L)}\boldsymbol{\varepsilon}_t \tag{11}$$

In which the only way the global and country-specific shocks get 'mixed up' is that they are a linear combination of the reduced-form residuals:

$$\boldsymbol{\varepsilon}_t = \mathbf{S}\mathbf{e}_t \tag{12}$$

As we assumed the global and country-specific shocks to be $i.i.d.$ and to have unit-variance as well as to be contemporaneously uncorrelated, the variance-covariance matrix Ω of the reduced-form residuals is given by

$$\boldsymbol{\Omega} = \mathbf{SS}' \tag{13}$$

In our two-dimensional system, this condition imposes three restrictions on $\mathbf{S}$. To just identify $\mathbf{S}$, one further restriction is needed.

Theory predicts that the current account should react only weakly to global shocks. We will exploit this property here to disentangle global from country-specific shocks. In so doing, we will impose the restriction that global shocks do not have an effect on the current account in the period they occur (they can however have a non-zero effect later). In fact, imposing this restriction amounts to a very simple identifying restriction: identification is achieved by means of a Choleski decomposition of the variance-covariance matrix of the reduced form residuals, $\mathbf{\Omega}$. To see this, note that the first component of $\boldsymbol{\varepsilon}_t$ is the reduced-form innovation to the current account. Requiring that only country-specific shocks drive this component, we get

$$\mathbf{S} = \begin{bmatrix} s_{11} & 0 \\ s_{21} & s_{22} \end{bmatrix} \tag{14}$$

But together with $\mathbf{\Omega} = \mathbf{SS}'$ this uniquely identifies $\mathbf{S}$ as the lower Choleski-factor of $\mathbf{\Omega}$.

Hence, we can map the structural MA-form into the reduced form:

$$\mathbf{C(L)S} = \mathbf{D(L)} \tag{15}$$

And as our interest will be particularly in long-run forces:

$$\mathbf{C(1)S} = \mathbf{D(1)}$$

We will now approximate $\mathbf{C(L)}$ by a VAR-representation. Note, however, that a finite-order VAR representation for $\mathbf{\Delta X}_t$ does not exist due to the presence of a common trend. It follows from Granger's representation theorem (Engle and Granger (1987)) that $\mathbf{\Delta X}_t$ can be represented in the form of a vector-error correction model (VECM):

$$\mathbf{\Gamma(L)\Delta X}_t = \boldsymbol{\alpha} CA_{t-1} + \boldsymbol{\varepsilon}_t \tag{16}$$

where $\mathbf{\Gamma(L)}$ is a 2×2 matrix-polynomial and $\boldsymbol{\alpha}' = \begin{bmatrix} \alpha_1 & \alpha_2 \end{bmatrix}$.

Once we have estimated this model, we can express the long-run structure of output as a function of the parameters of the VECM. In particular, as demonstrated in Johansen (1995), the matrix $\mathbf{C}(\mathbf{1})$ can be given a closed-form representation in terms of the parameters of the cointegrated VAR:

$$\mathbf{C}(\mathbf{1}) = \boldsymbol{\beta}_{\perp}(\boldsymbol{\alpha}'_{\perp}\boldsymbol{\Gamma}(\mathbf{1})\boldsymbol{\beta}_{\perp})^{-1}\boldsymbol{\alpha}'_{\perp}$$

Now note that the structure of this matrix is such that it maps the reduced-form disturbances $\boldsymbol{\varepsilon}_t$ into the span of $\boldsymbol{\alpha}_{\perp}$. The disturbances $\boldsymbol{\alpha}'_{\perp}\boldsymbol{\varepsilon}_t$ accumulate to the permanent component of $\mathbf{X}_t$ whereas transitory disturbances will be in the null space of $\mathbf{C}(\mathbf{1})$ We can therefore define the permanent disturbances as

$$\eta_t = \boldsymbol{\alpha}'_{\perp}\boldsymbol{\varepsilon}_t \tag{17}$$

and by requiring that permanent and transitory disturbances be orthogonal to each other, we get the transitory shocks as

$$\tau_t = \boldsymbol{\alpha}'\boldsymbol{\Omega}^{-1}\boldsymbol{\varepsilon}_t \tag{18}$$

Denoting

$$\boldsymbol{\theta}'_t = \left[\begin{array}{cc} \eta_t, & \tau_t \end{array}\right] \tag{19}$$

we then have $var(\theta) = diag\left\{var(\eta), var(\tau)\right\} = \left[\begin{array}{cc} \boldsymbol{\alpha}'_{\perp}\boldsymbol{\Omega}\boldsymbol{\alpha}_{\perp} & 0 \\ 0 & \boldsymbol{\alpha}'\boldsymbol{\Omega}^{-1}\boldsymbol{\alpha} \end{array}\right]$.

In the present bi-variate case with $\boldsymbol{\beta}' = \left[\begin{array}{cc} 1, & 0 \end{array}\right]$, we have $\boldsymbol{\beta}'_{\perp} = \left[\begin{array}{cc} 0, & 1 \end{array}\right]$. Furthermore, $\boldsymbol{\alpha}'_{\perp} = \left[\begin{array}{cc} -\alpha_2, & \alpha_1 \end{array}\right]$. Let also $\boldsymbol{\Gamma}(\mathbf{1}) = \left\{\gamma_{ij}\right\}_{i,j=1,2}$. Then it is easily verified that $\mathbf{C}(\mathbf{1})$ is of the form

$$\mathbf{C}(\mathbf{1}) = \left[\begin{array}{cc} 0 & 0 \\ c_{21}(1) & c_{22}(1) \end{array}\right] \tag{20}$$

where

$$c_{21}(1) = \frac{-\alpha_2}{-\alpha_2\gamma_{12} + \alpha_1\gamma_{22}} \text{ and } c_{22}(1) = \frac{\alpha_1}{-\alpha_2\gamma_{12} + \alpha_1\gamma_{22}} \tag{21}$$

3.1 The long-run effects of shocks

In a seminal paper, Blanchard and Quah (1989) identified demand and supply disturbances from a bivariate system, requiring that the former do not have a long-run effect on output. Their restriction postulates a form of long-run neutrality that - in various settings - is often suggested by economic theory. This is why the Blanchard-Quah identification scheme has proven very popular in applied work over the last decade (for applications of the Blanchard-Quah scheme see e.g. Bayoumi and Eichengreen (1992 a and b) and Bayoumi and Taylor (1995)).

Also in the context of this paper, the Blanchard-Quah identification seems an obvious candidate. Economic models will often require that country-specific shocks are long-run neutral w.r.t. output. For example in the Glick and Rogoff (1995) model, the empirical implementation will yield results that are at odds with the short-run dynamics of the intertemporal theory if in the theoretical model country-specific total factor productivity is required to follow a random walk.

In a recent study, Rogers and Nason (1998) use a structural VAR approach and employ various identification schemes. They find Choleski-type identifications to yield long-run dynamics that are inconsistent with long-run identification schemes in the spirit of Blanchard and Quah (1989) and vice versa. They do however, not single out one identification scheme that is superior to the others in its ability to identify global and country-specific shocks. This would require cross-model evidence which we will provide in this paper: the Choleski-identification scheme proposed in the previous section works well in identifying global and country-specific shocks. We will argue that it focuses on an immediate implication of the intertemporal approach (global shocks do not impinge on the current account) whereas the Blanchard-Quah scheme will ensue in some intertemporal models but not in others. After the model has been identified by the Choleski-scheme, it becomes possible to test the Blanchard-Quah scheme as an overidentifying restriction . We will now show that in the presence of a cointegrating relation it is particularly easy to test this overidentifying restriction.

Let for now the matrix $\mathbf{S} = \{s_{ij}\}_{i,j=1,2}$ define just any identification scheme such that $\mathbf{SS}' = \boldsymbol{\Omega}$.

Then from $\boldsymbol{\varepsilon}_t = \mathbf{S}\mathbf{e}_t$ and $\eta_t = \boldsymbol{\alpha}'_\perp \boldsymbol{\varepsilon}_t$ we get

$$\eta_t = (\alpha_1 s_{21} - \alpha_2 s_{11})\, e_t^c + (\alpha_1 s_{22} - \alpha_2 s_{12}) e_t^w \tag{22}$$

Requiring that country-specific shocks be long-run neutral then amounts to

$$\frac{s_{21}}{s_{11}} = \frac{\alpha_2}{\alpha_1}$$

This is a testable proposition (conditional of course, on the identifying assumptions that give us $\mathbf{S}$): α_2 and α_1 are parameters of the reduced form and as such their estimates are unaffected by the identification scheme chosen. As shown e.g. in Johansen (1995), linear restrictions on the space spanned by $\boldsymbol{\alpha}$ can be tested and these tests are asymptotically χ^2-distributed. In the present setting, the hypothesis can be formulated as follows:

$$\boldsymbol{\alpha} = \mathbf{H}\psi \text{ where } \mathbf{H} = \begin{bmatrix} s_{11}/s_{21} \\ 1 \end{bmatrix}$$

If furthermore, we want to take account of the estimation uncertainty in s_{21}/s_{11}, this will no longer be a linear hypothesis on $\boldsymbol{\alpha}$ only. Still there is a simple way to test the hypothesis. Note that with $\Omega = \{\omega_{ij}\}_{j,i=1,2}$, for the Choleski-factor we have

$$\mathbf{S} = \begin{bmatrix} \sqrt{\omega_{11}} & 0 \\ \omega_{21}/\sqrt{\omega_{11}} & \sqrt{\omega_{22} - \omega_{21}^2/\omega_{11}} \end{bmatrix}$$

and hence $s_{21}/s_{11} = \omega_{21}/\omega_{11}$. Then in the framework of the conditional model

$$\Delta Y_t = \frac{\omega_{21}}{\omega_{11}} \Delta CA_t + \left(\alpha_2 - \frac{\omega_{21}}{\omega_{11}}\alpha_1\right) CA_{t-1} + \textit{lagged dynamics}$$

testing the hypothesis we are interested in amounts to a $t-$test on whether the coefficient on CA_{t-1} is zero.

The Blanchard-Quah identification scheme links the period-zero impulse response of output and the current account, given by s_{21}/s_{11} to the relative long-run impulse response to (reduced-form) output and current account changes, given by α_2/α_1. This implies that the short-run dynamics of the system as given by the matrix **S** strongly influence the long-run dynamics and vice-versa. Under the Blanchard-Quah identification scheme, $\alpha_2 = 0$ implies $s_{21} = 0$ (note that in a cointegrated system $\boldsymbol{\alpha} = \mathbf{0}$ is not possible). Then, output is not only weakly exogenous in the long-run, but also, $\alpha_2 = s_{21} = 0$ implies that output is predetermined and also in the short-run unexpected output changes (which then coincide with global shocks) will drive the current account.

On the other hand, note that the Choleski-identification scheme we have suggested above will generically require the global shock to have some long-run impact on output: if **S** is the lower-Choleski-factor of Ω, $s_{12} = 0$ and $s_{22} > 0$. Hence, unless $\alpha_1 = 0$, i.e. we find the current account to be weakly exogenous, the Choleski-scheme will not be compatible with the Blanchard-Quah scheme w.r.t. to global shocks.

The preceding discussion puts us in a position to discuss the relative persistence of global and country-specific shocks. Recall the representation of the permanent shocks in (22) and note that the Choleski-identification scheme requires $s_{12} = 0$. Then

$$\eta_t = (\alpha_1 s_{21} - \alpha_2 s_{11})\, e_t^c + \alpha_1 s_{22} e_t^w \tag{23}$$

The coefficient on e_t^c, $\alpha_1 s_{21} - \alpha_2 s_{11}$, is a function of the output- and current-account response in period zero: s_{21} measures the period-zero output response to a country-specific shock whereas s_{11} measures the corresponding current-account response. These responses, in the long-run, get amplified by the coefficients $\alpha_1$and α_2. We can rewrite η_t as follows:

$$\eta_t = \alpha_1 \left[\left(\frac{s_{21}}{s_{11}} - \frac{\alpha_2}{\alpha_1} \right) s_{11} e_t^c + s_{22} e_t^w \right] \tag{24}$$

This equation tells us that the long-run impact of a one standard-deviation country-specific shock depends on the difference

$$\left(\frac{s_{21}}{s_{11}} - \frac{\alpha_2}{\alpha_1} \right)$$

The Blanchard-Quah identification scheme is compatible with the Choleski-scheme only if this difference is found to be zero.

The first ratio is the short-run impulse response of output relative to the current account. It tells us how a country-specific shock gets amplified in the period it occurs. The second term measures amplification as well, but now in the long-run: how much more strongly does output react to unexpected current account changes than to unexpected output changes?

Hence, we can interpret the difference between short-run and long-run adjustment as a measure of the relative contribution of country-specific shocks to the stochastic trend in output. Equivalently, we can understand it as a measure of the persistence of country-specific relative to global shocks. Because a measure of persistence should be positive, we here take the square of this difference and define:

$$\rho = \left(\frac{s_{21}}{s_{11}} - \frac{\alpha_2}{\alpha_1} \right)^2$$

Note also that this is a measure of persistence *net* of the relative variance of country-specific and global shocks: even if ρ is high, country-specific shocks may still explain a small share of long-run variance because they are less volatile than global shocks. In this sense, ρ tells us how much more persistent country-specific shocks are than global shocks - regardless of their respective volatilities. We address this issue in the next subsection.

3.2 What drives the common trend?

The share of long-run output variance explained by country-specific shocks is given by

$$\frac{(\alpha_1 s_{21} - \alpha_2 s_{11})^2}{(\alpha_1 s_{21} - \alpha_2 s_{11})^2 + \alpha_1^2 s_{22}^2}$$

which from the previous section can also be written as

$$\frac{\rho s_{11}^2}{\rho s_{11}^2 + s_{22}^2}$$

If $\alpha_1 = 0$, then the country-specific shock will explain all trend output growth variance and ρ goes to infinity. Shocks to the current account (which are assumed to be country-specific) accumulate to the stochastic trend in output and there will be no long-run feedback from output to the current account. We can think of the economy being driven by idiosyncratic shocks that are transmitted from the rest of the world.

If, however, $\alpha_2 = 0$, then the shocks to output drive the joint dynamics of the system and the current account is the variable that has to bear the adjustment burden in the long-run. Still, the share of trend output variance in this case will not be zero but is given by:

$$\frac{{s_{21}}^2}{{s_{21}}^2 + s_{22}^2}$$

The relative weight of country-specific shocks will depend on the relative period-zero impulse response of output to global and country-specific shocks. So, country-specific shocks will still have their role but now we should think of them as originating in the country, with the output reaction causally prior to the reaction of the current account.

Econometrically, tests of the hypothesis $\alpha_{1,2} = 0$ amount to tests of weak exogeneity in the sense of Engle, Hendry and Richard (1983): the dynamics of the remaining variable in the system can be correctly captured by conditioning on the weakly exogenous variable in the sense that no long-run feedback relations are neglected. We present tests of this hypothesis in the empirical section of the paper.

3.3 Assessing the quality of shock identification

The identification of global and country-specific shocks in this model rests on insights derived from the theory: not all countries of the world can run current account surpluses or deficits simultaneously. Hence, the world interest rate should adjust and the effect on current accounts should be small or even zero.

Even though this seems a plausible assumption, it is clearly not testable in the framework of the model as the Choleski-decomposition we impose is just-identifying. However, our analysis will proceed in the same way for all major seven industrialized countries. Those countries account for roughly 60 percent of world economic output. How global or country-specific the shocks we identified actually are can be assessed using cross-country information. We will discuss this issue here.

A logical starting point is certainly to look at cross-country correlations of global and country-specific shocks. Here, we would expect that on average, global shocks are more highly correlated across countries than country-specific ones. But how far should we push this idea? It seems unlikely that cross-country correlations of country-specific shocks are actually zero - shocks might after all be specific to a group of countries. Also, some upward movements in the current account in one country will correspond to downward movements in another country's current account. This reflects transmission of shocks and the fact that when we use the current account as an identification device for asymmetric/country-specific shocks, this means that the shock does not have to originate in this country. Rather, the country-specific shock is the outcome of a country's lending to and borrowing from many other countries, essentially an amalgam of many bilateral asymmetric shocks.

Likewise, global shocks should not be expected to be perfectly correlated. Rather, allowing for differences in internal transmission mechanisms, we should expect that the correlation is lower than unity.

An approach that takes account of the noise in the shock time series is principal component analysis. Let $\mathbf{E}_t^w = \{e_i^w\}_t^{i=1..7}$ be the vector of the stacked world-wide shocks and $\mathbf{E}_t^c$ be is the counterpart for the country-specific shocks. Then, the covariance matrix can be decomposed

$$\mathbf{cov}(\mathbf{E}) = \mathbf{P\Lambda P}' \tag{25}$$

where $\Lambda = diag(\lambda_1....\lambda_7)$ and $\lambda i \geqq \lambda_{i+1}$ $i = 1..6$. The principal components are given by $\mathbf{P}'\mathbf{E}_t$, where the first principal component explains the highest share of the variance, the second the second-highest etc.

In particular, it becomes possible to test how many principal components are sufficient to explain the variation in the data. A test for this kind of problem has been suggested by Bartlett (1954). The hypothesis of the Bartlett test is that the first k principal components explain the variance of the data whereas the last $p-k$ (where p is the dimension of the vector $\mathbf{E}$) are essentially indistinguishable. For the determinant of the dispersion matrix of normalized variables (i.e. like the shocks we are dealing with) is

$$\det(cov(E)) = \prod_{i=1}^{p} \lambda_i$$

Furthermore, it is

$$trace(E) = \sum_{i=1}^{p} \lambda_i = p$$

Hence, under the null

$$\det(cov(E)) = \lambda_1 \lambda_2 ... \lambda_k \left\{ \frac{p - \sum_{i=1}^{k} \lambda_i}{p-k} \right\}^{p-k}$$

The alternative is that there are $k+1$ significant principal components and the determinant of the dispersion matrix can then be written in an analogous way.

The ratio of the two determinants is given by

$$\left[\prod_{i=k+1}^{p} \lambda_i \right]^{-1} \left\{ \frac{\sum_{i=k+1}^{p} \lambda_i}{p-k} \right\}^{p-k} \tag{26}$$

When appropriately scaled with a factor involving sample size, the log of this expression can be given an approximate χ^2-distribution.

In the context of our problem, we would expect that such a test detects only one principal component that explains the variation in the data once we apply it to global shocks and a much larger number of significant principal components among the country-specific shocks.

Also, the theory suggests that the principal component driving the global shocks is associated with the world interest rate. We can test this implication by comparing $\mathbf{p}'\mathbf{E}_t^w$ with a measure of the world interest rate, where $\mathbf{p}'$ is the first row of $\mathbf{P}'$.

4 Empirical results

4.1 Estimation and model specifications

In this section, we report the results of the estimation of our model for the G7 countries. The data we used are annual real GDP from Gordon (1993), 1960-91 and current account / GDP ratios from Taylor (1996) and originally due to Obstfeld and Jones (1990). In order to make output volatilities comparable across countries, we transformed output into an index by dividing through by the first observation. We also divided the current account by the first observation of output, i. e. we considered $\mathbf{X}_t = \left[\ CA_t,\ \ Y_t\ \right]' / Y_0$. Standard information criteria suggested that the seven models should be specified with one or two lags. We decided for two lags throughout. The model was then estimated with an unrestricted constant term.

We also included a number of conditioning variables in some of the models: in testing for the number of cointegrating relationships, we could not reject the null of no cointegration in the case of the US and Canada. This, however, should not be too surprising as the theoretical model is designed for a small open economy in that is treats the world interest rate as fixed. The US interest-rate, however,seems to play an important global role. Indeed, it is likely that the U.S.current account contains a large 'speculative' component that is the outcome of international capital flows induced by changes in the interest rate differential vis-a-vis the rest of the world..

We therefore decided to include the German-U.S. interest rate differential as an exogenous regressor into the model for the US. Even though we found the UK current account to be stationary, it is likely to be driven to a large extent by changes in the price of oil. Movements in the oil price, however, are prime candidates for global shocks, so we decided to condition the model for the UK on the price of oil.

In table 1 we present the results of Johansen's tests for cointegration after the inclusion of conditioning variables. Generally, we reject the null of no cointegration more strongly than without those variables. For six countries we find one cointegrating relationship at the 5-percent level. In particular we now also find a highly significant cointegrating relationship in the U.S. case. Only for Canada we continue to accept the null. Still we decided to impose one cointegrating relationship in the estimation of all seven models.

Table 1: Johansen's tests for cointegration

	Trace test		MaxEV test	
H_0	$h=0$	$h=1$	$h=0$	$h=1$
US	**30.35**	2.639	**27.71**	2.639
Japan	*17.04*	4.045	*13*	4.045
Germany	**18.2**	2.052	**16.15**	2.052
France	13.79	0.6392	*13.15*	0.6392
Italy	**25.68**	0.04728	**25.63**	0.04728
UK	**21.25**	4.096	**17.16**	4.096
Canada	10.25	0.4452	9.804	0.4452
90% crit. val	15.58	6.69	12.78	6.69
95% crit. val.	17.84	8.803	14.6	8.083

5 (10) %-significant values are in bold (italics)

Once we impose a cointegrating relationship in the estimation, tests of the cointegrating space show that it is generally the current account that is stationary: for six countries is the hypothesis that $\beta' = \begin{bmatrix} 1, & 0 \end{bmatrix}$ is accepted at the 5-percent level. For Germany there seems to be a small but significant coefficient on output in the cointegrating vector. Our unrestricted estimate of β for Germany is $\begin{bmatrix} 1, & -0.08 \end{bmatrix}$.

Table 2: tests on the cointegrating space $\beta' = \begin{bmatrix} 1, & \beta_2 \end{bmatrix}$

	US	Japan	Germany	France	Italy	UK	Canada
β_2	-0.003	0.01	-0.08	0.0004	-0.001	0.05	0.01
p-value	0.83	0.46	0.001	0.94	0.83	0.09	0.25

Based on these pre-test results, we decided to proceed as follows: we imposed one cointegration relation in the estimation of all seven models.

However, in the estimation of the German model we left the cointegrating space unrestricted.

4.2 Global and country-specific shocks

We are now in a position to discuss the quality of the identification scheme we have proposed for global and country-specific shocks.

We start by exposing the correlation matrices of global and country-specific shocks and their average value across countries (this cross-sectional mean excludes the country itself, of course) in table 3. Here, we find first favourable evidence that our scheme works fairly well. Global shocks are on average more highly correlated than country-specific shocks. Also, the p-values of the global shock are much lower and the cross-sectional mean is significant at conventional levels in four out of seven cases, whereas for the country-specific shock it is never found to be significant.

Table 3 a): cross country correlation of country-specific shocks

	US	Japan	Germany	France	Italy	UK	Canada
US	1						
Japan	-0.19	1					
Germany	-0.22	0.28	1				
France	0.001	0.25	-0.14	1			
Italy	-0.07	0.27	-0.06	0.65	1		
UK	0.09	0.18	-0.47	0.10	0.16	1	
Canada	0.17	-0.29	0.01	-0.34	-0.30	0.03	1
mean	-0.03	0.08	-0.09	0.08	0.10	0.01	-0.12
std-dev.	0.15	0.25	0.25	0.34	0.33	0.24	0.22
p-value	0.40	0.37	0.35	0.40	0.38	0.47	0.30

values of cross-sectional means significant at 5 (10)% are in bold (italics)

Table 3 b): cross-country correlation of global shocks

	US	Germany	Japan	France	Italy	UK	Canada
US	1						
Germany	0.40	1					
Japan	0.29	0.28	1				
France	0.37	0.44	0.36	1			
Italy	-0.07	-0.11	0.36	0.36	1		
UK	0.29	0.37	0.45	0.44	0.20	1	
Canada	0.70	0.16	0.33	0.40	0.25	0.51	1
mean	0.33	0.25	**0.35**	**0.40**	0.16	**0.38**	**0.39**
std-dev.	0.25	0.20	0.06	0.040	0.21	0.11	0.19
p-value	0.12	0.13	0.001	0.00	0.23	0.01	0.04

values of cross-sectional means significant at 5 (10)% are in bold (italics)

We then proceeded to test whether principal component analysis makes any sense in our setting. If shocks are spherical or at least independent, then there is no point in finding a rotation such that one direction explains as much as possible of the variance. In other words: orthogonalizing the variates would not carry any benefit in this case as the variates *are* already orthogonal. Before proceeding to an analysis of principal components, we therefore performed a test of independence for both E^c and E^w.

The test clearly rejected the null of independence for both types of shocks ($p-$values of 0.01 and 0.00). In the case of country-specific shocks, this suggests that international transmission of these shocks plays an important role.

Table 4 gives the results of the principal component analysis. subtable a) for the global shock and subtable b) for the country-specific shocks. The first principal component of the global shock identified for the G7 explains 43 percent of the variance whereas for the country-specific shock it accounts for only 30 percent of the variance. This hints at a higher degree of 'commonality' among the global shocks.

In the fourth column of the same table we also provide the results of the Bartlett tests for dimensionality. At a conventional significance level of 5 percent, the tests suggests that country-specific shocks have one distinguishable principal components whereas the global shock displays five. This result seems somewhat at odds with our earlier finding that country-specific shocks have a lower cross-sectional correlation than global shocks. But note that once we lower the size of the test to 1 percent, then the principal components of the country-specific shock become indistinguishable whereas only two principal components survive for the global shock. Our results suggest that there is a reduced number of driving forces behind the global shocks. We will now try to identify these driving forces with observable economic variables. There are a few obvious candidates: as has been put forward in the introductory sections of this paper, theory suggests that changes in world interest rates are a prime candidate. Another variable is US-output growth.

Table 4 a): Principal component analysis of global shocks

Principal Comp.	Variance explained	Latent roots	Bartlett Test
1	43.66	3.056	2.981e-007
2	18.46	1.292	0.007342
3	13.48	0.9434	0.02079
4	9.463	0.6624	0.03481
5	8.208	0.5745	0.02402
6	4.612	0.3228	0.1096
7	2.12	0.1484	NaN

Table 4 b): Principal component analysis country-specific shocks

Principal Comp.	Variance explained	Latent Roots	Bartlett test
1	30.95	2.167	0.01094
2	23.54	1.648	0.05675
3	14.14	0.9901	0.2474
4	12.02	0.8413	0.1864
5	10.3	0.7211	0.1723
6	5.095	0.3566	0.7854
7	3.951	0.2766	NaN

The first and second principal components of the global shock are plotted in figure 1. Figure 2 gives the US output growth rate whereas figure 3 plots the US ex-post real interest rate.

Figure 4 plots the first principal component and the US output growth rate and figure 5 presents changes in the real interest rate and the second principal component.

The close comovement between US output growth and the first principal component that is apparent from the visual impression of figure 4 is confirmed by the correlation which is 0.68. There seems to be a link between the second principal component and the real interest rate but it does not show up very strongly in the correlation which is found to be 0.24. Also, this correlation is positive whereas from the theory we would expect that positive global shocks are associated with decreases in the real interest rate. Still, figure 5 suggests an important link between the two variables. We therefore proceeded to a more formal analysis of their joint time-series properties. Following the modelling approach suggested in Gonzalo and Granger (1995), we cumulated the second principal component of the global shock and the changes in the real interest rate. We then specified a cointegrated VAR in 2 lags:

$$\mathbf{\Gamma}_z(\mathbf{L})\mathbf{\Delta Z}_t = \boldsymbol{\alpha}_z\boldsymbol{\beta}_z'\mathbf{Z}_{t-1} + \mathbf{v}_t$$

where $\mathbf{Z}_t' = \left[\ \sum_{i=0}^{t} e_t^w, \quad r_t\ \right]$ and the covariance structure is given by

$$\mathbf{\Sigma} = \mathbf{var}(\mathbf{v}_t) = \{\sigma_{ij}\}_{i,j=1,2}$$

We included an unrestricted constant and a step dummy to account for the secular increase in interest rates in the early eighties. Johansen's (1988) test suggested the presence of one cointegrating relationships. The estimated cointegrating vector was $\beta_Z' = \left[\ 1, \quad 0.62\ \right]$ and the hypothesis $H_0 : \beta_Z' = \left[\ 1, \quad 1\ \right]$ was accepted with p-value 0.2. This suggests that in the long-run changes in the real interest rate are perfectly inversely correlated with global shocks.

Tests also suggested that the real interest rate represents the common stochastic trend in $\mathbf{Z}_t$, i.e. we found $\alpha_{2Z} = 0$ which suggests that we can write a conditional model of the global shock:

$$e_t^w = \frac{\sigma_{21}}{\sigma_{22}}\Delta r_t + \left(\alpha_{1Z} - \frac{\sigma_{21}}{\sigma_{22}}\alpha_{2Z}\right)\left(\sum_{i=0}^{t-1} e_i^w + r_{t-1}\right) + \textit{lagged dynamics}$$

Our estimate of σ_{21}/σ_{22} is -0.48, much higher in absolute terms than the correlation between Δr_t and e_t^w that we calculated earlier and that we found to be 0.24. Also, the correlation is now negative, in accordance with the theory.

The results suggest that the global shock is indeed negatively related to movements in the real interest rate. In the long-run the correlation seems perfect, whereas in the short-run it is somewhat less pronounced.

4.3 Persistence and the relative importance of global and country-specific shocks

Table 6 provides our estimates of persistence for country-specific shocks. The results are very interesting: for the four smallest economies, country-specific shocks are found to be much less persistent than global shocks,whereas for the G3, the U.S., Japan and Germany, we find them to be 6-15 times more persistent than global shocks. This result may be due to two reasons: the G3 economies are large vis-a-vis the other four economies and therefore may find it difficult to fully smooth country-specific shocks through international borrowing and lending. Country-specific shocks may therefore become very persistent relative to global shocks. On the other hand, our procedure may suffer from some mismeasurement. As our results have shown so far, it is more likely to work well with a small open economy and country-specific U.S.-shocks are correlated with global shocks.

Table 6: Relative persistence of e^c vs. e^w

	US	Japan	Germany	France	Italy	UK	Canada
ρ	9.155	6.335	15.42	0.1721	0.7241	0.01657	0.026

In table 7 we test the overidentifying restriction imposed by the Blanchard-Quah identification, i.e. that $\rho = 0$. The first row in the table pertains to the 'naive' test in which we assume s_{11}/s_{21} fixed and just test a linear restriction on $\boldsymbol{\alpha}$. The second row gives the test based on the regression of ΔY_t on ΔCA_t, CA_{t-1} and lagged values. The 'naive' test clearly rejects the hypothesis for the US, Japan, Germany and Italy. This picture is not changing a lot once we do the regression test. However, the US becomes a borderline case now with the hypothesis accepted at the 13-percent level. In particular for the UK and Canada the data support the Blanchard-Quah identification. If we disregard the case of Italy, a general pattern is suggested by the data: the smaller the economy, the more likely are country-specific shocks to be long-run neutral w. r. t. output.

Table 7: Tests of the Blanchard-Quah restriction

	US	Japan	Germany	France	Italy	UK	Canada
Test on $\boldsymbol{\alpha}$							
LR	13.44	9.06	15.2	0.92	15.41	1.06	0.48
p-val.	0.0002	0.002	0.00	0.33	0.00	0.30	0.48
Regression test on $(\alpha_2 - \frac{\sigma_{21}}{\sigma_{11}}\alpha_1)$							
t-val.	1.13	2.63	3.26	1.018	3.97	0.87	0.17
p-val.	0.13	0.006	0.001	0.1588	0.00	0.19	0.43

LR is distributed as $\chi^2(1)$ and t-stat as $t(27)$

In table 8, we present the results of forecast error decompositions of changes in output and the current account. The result is interesting to contrast with our estimates of persistence: country-specific shocks seem to fully explain changes in the current account. This corroborates an important prediction of the intertemporal theory which predicts that the current account response to global shocks should be negligible. It also lends additional support to the validity of our identification scheme: if we think of a smooth current account response to structural shocks then we should not have done the data too much harm by imposing a zero-restriction in period zero.

It is interesting to compare the output decomposition with our estimates of the persistence of country-specific shocks: in the short-run global shocks explain the bulk of the output variance but the share of the country-specific shock is never negligible.

In the long-run the share of the country-specific shock increases. in particular so in the case of the G3. This reflects the high persistence of country-specific shocks in these countries. But note that even at the 10-year forecast horizon, country-specific shocks never explain much more than 50 percent of changes in output whereas the shocks where found to be 6-15 times as persistent as global shocks.

Table 8a: Variance share of ΔCA explained by country-specific shock

FC-horizon	US	Japan	Germany	France	Italy	UK	Canada
1 year	1	1	1	1	1	1	1
2	0.95	0.99	0.97	0.98	0.96	0.73	0.99
5	0.95	0.99	0.97	0.97	0.95	0.76	0.99
10	0.95	0.99	0.97	0.97	0.95	0.76	0.99

Table 8b: Variance share of ΔY explained by country-specific shock

FC- horizon	US	Japan	Germany	France	Italy	UK	Canada
1 year	0.38	0.0000	0.00	0.03	0.27	0.31	0.12
2	0.50	0.04	0.01	0.14	0.61	0.31	0.11
5	0.52	0.31	0.48	0.15	0.67	0.38	0.13
10	0.53	0.34	0.57	0.15	0.67	0.38	0.13

Table 9 gives the share of trend output variance that is explained by country-specific shocks. In line with our earlier finding that country-specific shocks are very persistent in the G3 countries, the share of variance that can be ascribed to these shocks is between 20 and 30 percent for Japan and Germany and amounts to roughly 80 percent for the US. Among the smaller G7-economies, Italy is special in the sense that 40 percent of trend output variance is explained by the country-specific shock. For all other countries, the share of trend output variance explained by the country-specific shock is negligible.

Overall, the variance decompositions suggest that country-specific shocks are generally less volatile than global ones. The diagonal entries of **S** measure the variance of the structural shocks. Indeed, table 10 that gives the estimates of the ratio s_{11}/s_{22} shows that global shocks are generally one and a half (0.63^{-1})times as volatile as country-specific ones.

Table 9: Share of e^c in trend output variance

US	Japan	Germany	France	Italy	UK	Canada
0.80	0.20	0.29	0.14	0.41	0.00	0.01

Table 10: Relative variance of e^c and e^u- estimates of s_{11}/s_{22}.

US	Japan	Germany	France	Italy	UK	Canada	Average
0.3019	0.5039	0.6351	0.6456	1.288	0.5913	0.4634	0.6348

Table 11 provides the results of the tests for weak exogeneity, i.e. of the hypotheses $\alpha_i = 0,\ i = 1, 2$. It is interesting to note that with the exception of Italy we find that at the 5-percent level at least one variable is found to be weakly exogenous for all countries.

Table 11: Tests of weak exogeneity (p-values)

	US	Japan	Germany	France	Italy	UK	Canada
CA	0.00	0.01	0.13	0.00	0.00	0.00	0.00
Y	0.62	0.08	0.00	0.13	0.00	0.16	0.53

In the US and German cases, the current account is clearly found to be weakly exogenous. Note that, under the Choleski-identification, this amounts to saying that global shocks have no long-run effect on output. In both the German and US cases, the Blanchard-Quah restriction was found to be strongly rejected (table 7).

This is compatible with the picture that emerged earlier in which the U.S. output trend is purely domestically determined but acts as a generator for world-wide macroeconomic fluctuations. For Germany, the finding that the current account drives the common trend and the fact that a non-trivial cointegrating relationship prevails between output and the current account suggests that German trend output growth in the period 1960-91 has largely been driven by shocks to the export sector, a notion that is frequently referred to as 'export-led' growth. (see e.g. the study by Marin (1992))

5 Conclusion

In this paper, we have suggested using the reduced form of a simple intertemporal model of the current account to measure stylized facts in the international transmission of macroeconomic disturbances. We have proposed a simple identification scheme for global and country-specific shocks. The identification scheme was assessed using cross-country evidence and seems to work reasonably well: global shocks are more highly correlated across countries than are country-specific shocks. Also, there are two dominant principal components among global shocks. Whereas one of them can straightforwardly be associated with US-output growth, the second one displays some short-run and perfect long-run correlation with a measure of the ex-post US real interest rate.

We have then used the proposed framework to collect stylized facts about the external adjustment of the G7 economies. Our results can be summarized as follows:

- Country-specific shocks account for most of the current account variance. This finding corroborates an important prediction of the intertemporal approach to the current account which suggests that the current account should react to the country-specific shock only.
- Country-specific shocks are much more persistent than global ones in the G3 economies and much less than global ones in the smaller G7 countries. Generally, the smaller the country, the less persistent are country-specific shocks.
- Country-specific shocks are generally found to explain only a moderate share of trend output growth.
- On average, global shocks are one and a half times more volatile than country-specific ones.
- Global shocks have two dominant principal components: the more important one is found to be highly correlated with US output

growth. In accordance with the intertemporal approach to the current account, the second one is in the long-run perfectly negatively correlated with the real interest rate. In the short-run there seems to prevail a smaller negative correlation.

- Changes in the US interest rate seem to trigger important current account reactions that are then found to be statistically exogenous w.r.t. to output dynamics in this country.

- In Germany, there is a non-trivial cointegrating relationship between output and the current account. Also, the current account seems to drive the stochastic trend in output as it is found to be weakly exogenous. Evidence for the German case seems inconclusive. We propose to interpret our findings as evidence of Germany's output growth over the period being driven by export-shocks.

References

[1] Barro, E. and Sala-i-Martin, X (1990).'World interest rates' in O. Blanchard and S. Fischer, eds., *NBER macroeconomics annual* 1990.

[2] Bartlett, M. S. (1954), 'A note on the multiplying factors for various χ^2-approximations', *Journal of the Royal Statistical Society*, Ser. B. 16, 296.

[3] Bayoumi, Tamim and Taylor, Mark P., (1995), "Macroeconomic Shocks, the ERM and Tripolarity," *The Review of Economics and Statistics*, 321-331.

[4] Bayoumi, Tamin and Eichengreen, Barry, (1992a), 'Shocking Aspects of European Monetary Union,' CEPR Discussion-Paper no. 643.

[5] Bayoumi, Tamin and Eichengreen, Barry, (1992b), 'Macroeconomic Adjustment under Bretton Woods and the Post-Bretton-Woods Float: An Impulse Response Analysis,' CEPR Discussion Paper no. 729.

[6] Beveridge, Stephen and Nelson, Charles R., "A New Approach to Decomposition of Economic Time Series into Permanent and Transitory Components with Particular Attention to Measurement of the 'Business Cycle"', *Journal of Monetary Economics* 7, 151-173.

[7] Blanchard, Olivier and Quah, Danny, (1989), "The Dynamic Effects of Aggregate Demand and Supply Disturbances," *American Economic Review* 79, 655-73.

[8] Campbell, J. Y, and Shiller, R.J., (1987), ' Cointegration and tests of present value models,' *Journal of Political Economy* 95: 1062-1088.

[9] Engle, R.F., Hendry, D.F. and Richard, J.-F. (1983), 'Exogeneity', *Econometrica* 51: 277-304.

[10] Engle, Robert F. and Granger, Clive W. J. (1987), "Cointegration and Error-Correction: Representation, Estimation and Testing," *Econometrica* 55: 251-76.

[11] Gonzalo, Jesus and Granger, Clive, (1995), "Estimation of Common Long-Memory Components in Cointegrated Systems," *Journal of Business and Economics Statistics* 13: 27-35.

[12] Gordon, R.J. (1993), *Macroeconomics*, Addison-Wesley. Data set also to be found at *http://www.nber.org/pub/gordon.*

[13] Johansen, Søren, (1988), "Estimation Hypothesis Testing of Cointegration Vectors in Gaussian Vector Autoregressive Models". *Econometrica* 59: 1551-80.

[14] Johansen, Søren, (1995), *Likelihood-Based Inference in Cointegrated Vector Autoregressive Models*, Oxford University Press.

[15] Johansen, Søren, (1997), "Common Trends in the Cointegrated VAR Model," *European University Institute*, mimeo.

[16] Kendall, M. and Stuart, A. (1975), *The Advanced Theory of Statistics*, vol. 3. London.

[17] King, Robert, Plosser, Charles I., Stock, James H. and Watson, Mark W., (1991), "Stochastic Trends and Economic Fluctuations", *American Economic Review* 81: 819-40.

[18] Marin, D. (1992), 'Is the Export-Led Growth Hypothesis valid for industrialized countries?', *Review of Economics and Statistics* 74: 678-88.

[19] Nason, J. and Rogers, J. H. (1998), 'Investment and the Current-Account in the Short-Run and the Long Run', mimeo, University of British Columbia.

[20] Obstfeld, M., (1986), 'Capital Mobility in the World Economy: theory and measurement', *Carnegie Rochester Conference Series on Public Policy* 24: 55-104.

[21] Obstfeld, M., (1995), 'International Capital Mobility in the 1990s' in Kenen, P. (ed.) *Understanding Interdependence*, Princeton University Press, 1995.

[22] Obstfeld, M. and Rogoff, K. (1995), 'The intertemporal approach to the current account,' NBER working paper 4893.

[23] Obstfeld, M. and Rogoff, K., (1996), *Foundations of International Macroeconomics*, MIT Press, Cambridge.

[24] Sachs, J, (1981), 'The Current Account and Macroeconomics adjustment in the 1970s,' *Brookings Papers on Economic Activity* 1: 201-268.

[25] Stock, James H. and Watson, Mark W., (1988), "Testing for Common Trends", *Journal of the American Statistical Association* 83, 1097-1107.

6 Figures

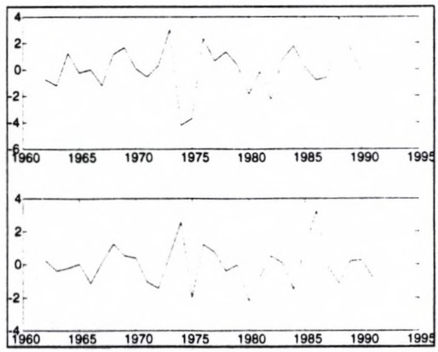

Figure 1: First (upper panel) and second principal components of the global shocks

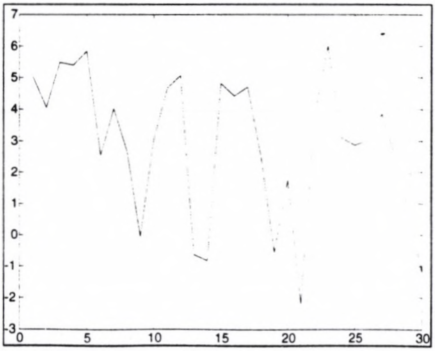

Figure 2: US GDP growth rates 1960-91.

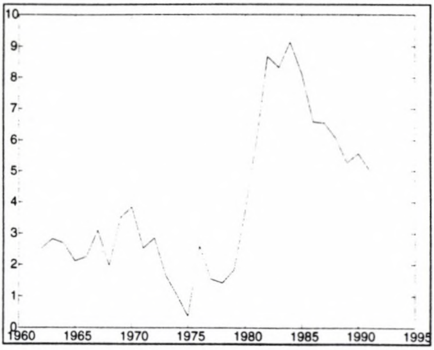

Figure 3: US real interest rate (ex-post, based on GDP-Defl.)

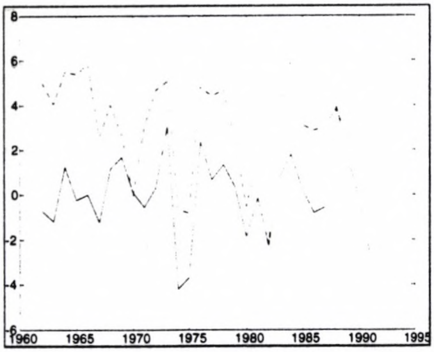

Figure 4: US GDP growth rates and the first principal component of global shocks

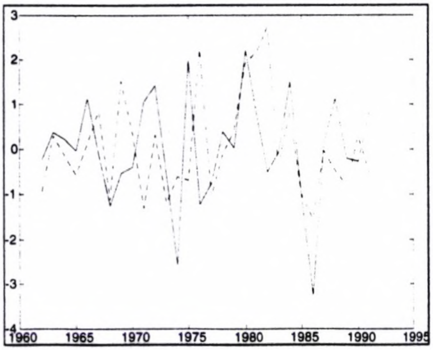

Figure 5: Changes in the real interest rate (dashed) and second principal component of global shocks

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