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International Macroeconomic Fluctuations and the Current Account[1]

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Abstract

Intertemporal models of the current account generally assume that global shocks do not affect the current account. We use this assumption to identify global and country-specific shocks in a bivariate VAR. We test the quality of the identification using evidence from G7-data. In accordance with the theory, we observe a link between the global shock and a measure of the world real interest rate. We also find that long-term output growth is driven by global factors in most countries, that country-specific shocks are less persistent in smaller economies and generally less volatile than global shocks.

Keywords: Intertemporal approach to the current account, Cointegration, SVAR, International Business Cycles

JEL classification: F41, F43, C32

1 Introduction

Most intertemporal optimising models of the current account (Sachs (1981), Obstfeld (1986, 1995), Obstfeld and Rogoff (1996)) assume that global shocks should not have an impact on the current account. This assumption is best justified from Metzler's (1960) theory of the world real interest rate: the world is a closed economy and it is impossible that, in response to a global shock, all countries change their current account positions in the same direction. Therefore, the world real interest rate has to equate world savings and investment, leaving individual countries' current account positions unaffected.

Glick and Rogoff (1995) have estimated a structural econometric model in which they found, indeed, the response of the current account to global shocks to be insignificant. But whereas Glick and Rogoff construct country-specific and global Solow-residuals which then enter their econometric framework as generated regressors, the present paper follows the opposite approach: we use the assumption of the theory as a device to identify global and country-specific components in a bivariate vectorautoregression. We then test the quality of this theory-driven identification scheme using cross-country evidence. In spite of its simplicity, our method does very well in identifying country-specific and global shocks and we are able to identify global shocks with a measure of the real world rate of interest.

Our approach gives us a powerful empirical framework with which to fish for stylized facts in the world economy. Little stylized knowledge is available on the question in which way the major industrialized countries are prone to international shocks and how they adjust to them. Our empirical setup that focuses on the current account as the key variable of international macroeconomic transmission 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.

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:

[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.

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

Here, $\mathbf{e}_t = \left[e_t^c, e_t^w\right]'$ 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 generally driven by both global and country-specific shocks. If however, international capital mobility is sufficiently high, global shock should impinge on the world interest rate, leaving current accounts unaffected. In the structural MA representation (8), this amounts to requiring that, once, we condition on past information, i.e. in particular on past current account positions, global shocks should have no effect in the period they occur, i.e. $\mathbf{D}_0 = \mathbf{D}(\mathbf{0})$ should be lower triangular. In the next section, we exploit this feature for identification purposes.

3 Econometric Implementation

The structural form (8) cannot be estimated directly. Rather, denoting $\mathbf{X}_t' = \left[CA, Y_t\right]$, we estimate a reduced-form moving average representation

$$\mathbf{\Delta X}_t = \mathbf{C}(\mathbf{L})\varepsilon_t \tag{9}$$

We require the reduced-form residuals to be a linear combination of the structural shocks:

$$\varepsilon_t = \mathbf{S}\mathbf{e}_t \tag{10}$$

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 $\mathbf{\Omega}$ of the reduced-form residuals is given by

$$\mathbf{\Omega} = \mathbf{S}\mathbf{S}' \tag{11}$$

In our two-dimensional system, this condition imposes three restrictions on $\mathbf{S}$. To just identify $\mathbf{S}$, one further restriction is needed. In section two we required that $\mathbf{D}_0 = \mathbf{D}(\mathbf{0})$ be lower triangular, because global shocks are not supposed to have an impact on the current account in the period they occur. From this and $\mathbf{C}_0 = \mathbf{C}(\mathbf{0}) = \mathbf{I}_2$ we get

$$\varepsilon_t = \mathbf{D}_0\mathbf{e}_t = \mathbf{S}\mathbf{e}_t = \begin{bmatrix} s_{11} & 0 \\ s_{21} & s_{22} \end{bmatrix} \mathbf{e}_t \tag{12}$$

which identifies $\mathbf{S}$ as the lower Choleski factor of $\mathbf{\Omega}$.

For estimation, we approximate $\mathbf{C}(\mathbf{L})$ 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 cointegration. As Campbell and Shiller (1987) have shown, present-value relations like (7) give rise to cointegration. In the present context, the cointegrating relationship is a trivial one: as can be seen from (8), the current account is an $I(0)$ process whereas Y_t contains a unit-root. The system is therefore cointegrated with cointegrating vector $\beta' = \begin{bmatrix} 1, 0 \end{bmatrix}$. 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):

$$\mathbf{\Gamma}(\mathbf{L})\mathbf{\Delta X}_t = \alpha CA_{t-1} + \varepsilon_t \tag{13}$$

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

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 with respect to 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 $\mathbf{SS}'= \mathbf{\Omega}$. It is well known (see e.g. Johansen (1995)) that in a cointegrated model like (13) the innovations to the common trend are given by

$$\eta_t = \alpha'_\perp \varepsilon_t \tag{14}$$

where $\alpha_\perp$ is the orthogonal complement to α. In the present bivariate setup with one cointegrating relationship, $\alpha'_\perp = \left[-\alpha_2, \alpha_1\right]$. Then from $\varepsilon_t= \mathbf{S}\mathbf{e}_t$ and $\eta_t= \alpha'_\perp \varepsilon_t$ we get

$$\eta_t = (\alpha_1 s_{21} - \alpha_2 s_{11})\, e^c_t + (\alpha_1 s_{22} - \alpha_2 s_{12}) e^w_t \tag{15}$$

If S is the Choleski-factor of Ω then $s_{12} = 0$ and (15) specializes to

$$\eta_t = (\alpha_1 s_{21} - \alpha_2 s_{11})\, e^c_t + \alpha_1 s_{22} e^w_t$$

From this we get a simple closed-form expression for the long-run variance share of country-specific shocks in trend output growth:

$$\frac{(\alpha_1 s_{21} - \alpha_2 s_{11})^2}{(\alpha_1 s_{21} - \alpha_2 s_{11})^2 + \alpha_1^2 s_{22}^2} \tag{16}$$

Requiring that country-specific shocks be long-run neutral then amounts to requiring that the coefficient on e_t^c in (15) should equal zero. As long as we assume s_{21} and s_{11} to be non-stochastic, this amounts to a linear restriction on α. As shown e.g. in Johansen (1995), linear restrictions on the space spanned by α can be tested and these tests are asymptotically χ^2-distributed. In the present setting, the hypothesis can be formulated as follows:

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

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 α only. Still there is a simple way to test the hypothesis. Note that with $\mathbf{\Omega} = \{\omega_{ij}\}_{j,i=1,2}$, 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.

3.2 Assessing the quality of shock identification

The identifying assumption (12) is not testable in the framework of the cointegrated VAR (13) 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 start by looking at cross-country correlations of global and country-specific shocks, expecting 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

$$\mathbf{cov}(\mathbf{E}) = \mathbf{P \Lambda P}' \tag{17}$$

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. In the context of our problem, we would expect that the Bartlett test detects only one or very few principal components that explain the variation in the data once we apply it to global shocks. Conversely, we should expect that no principal components can be distinguished among country-specific shocks.

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 and we decided to condition the model for the UK on this variable.

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.

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 $\begin{bmatrix} 1, -0.08 \end{bmatrix}$.

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 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.

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, panel a) for the global shock and panel 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 obvious variable is US-output growth.

Figure 1 plots the first principal component and the US output growth rate whereas figure 2 presents the second principal component and changes in the ex post US real interest rate that we use to proxy world interest rates in this paper.

The close comovement between US output growth and the first principal component that is apparent from the visual impression of figure 1 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 2 suggests an important link between the two variables that might, however, only be reflected in their longer swings. Identifying this comovement requires the use of cointegration techniques. Gonzalo and Granger (1995) have suggested to examine the long-run properties of larger econometric systems by extracting common trends from low-dimensional VARs and analysing the comovement of the common trends in a separate VAR. We adopt their approach here: we cumulated the second principal component of the global shock which is nothing else than a linear combination of the common trends that we extracted from the seven country-models. It is given by

$$g_t = \iota_2' \mathbf{P}_w' \mathbf{E}_t^w \tag{18}$$

where ι_2 is just the second unit-vector and P_w is the loadings matrix defined by (17). We then specified a cointegrated VAR in the real interest rate and the cumulated second principal component of global shocks:

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

where $\mathbf{Z}_t' = \left[\sum_{i=0}^{t} g_i, r_t\right]$ 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. Two lags were sufficient to whiten residuals. Johansen's (1988) test suggested the presence of one cointegrating relationships. The estimated cointegrating vector was $\beta_Z' = \left[1, 0.62\right]$ and the hypothesis $H_0 : \beta_Z' = \left[1, 1\right]$ 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:

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

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

In table 5 we test the overidentifying restriction imposed by the Blanchard-Quah identification. 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 α. 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 with respect to output.

Table 6 gives the share of trend output variance that is explained by country-specific shocks, calculated according to (16). 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 long-run variance decomposition suggests that country-specific shocks generally contribute less to trend output volatility than do global shocks. Are countries hit harder by global shocks as compared to country-specific shocks or is it that global shocks are simply more volatile? The diagonal entries of $\mathbf{S}$ measure the variance of the structural shocks. Indeed, table 7 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 8 provides the results of tests for weak exogeneity, i.e. of the hypotheses $\alpha_i = 0,\ i = 1, 2$. These tests tell us which of the variables can be interpreted as the 'driving force' of the system. If, say, $\alpha_2 = 0$. then $\alpha'_\perp = \left[0, 1\right]$ and output is weakly exogenous in the long-run (see e.g. Johansen (1995)). From (15), the common trend would then only be fed by innovations in output itself, making any long-run impact of the current account on output impossible. 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.

In the US and German cases, it is the current account that is clearly found to be weakly exogenous. Under the Choleski-identification, this amounts to saying that global shocks have no long-run effect on output, as the coefficient on e_t^w in (15) will equal zero. Not that, conversely, in both the German and US cases, the Blanchard-Quah restriction was found to be strongly rejected when applied to country-specific shocks(table 5).

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 that is based on a simple Choleski decomposition. The identification scheme was assessed using cross-country evidence works surprisingly well. 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:

- Shocks orthogonal to the current account are more highly correlated across countries than are shocks to the current account themselves.

- There are two dominant principal components among shocks that are orthogonal to the current account 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.

- These results justify to call shocks to the current account 'country-specific' and those orthogonal to the current account 'global'. The assumption of intertemporal optimising models of the current account that global shocks do not affect the current account can therefore usefully be employed to identify small econometric models.

- 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.

- Changes in the US interest rate seem to trigger important current account reactions that are then found to be statistically exogenous with respect 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.

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6 Tables and Figures

Table 1: Johansen's tests for cointegration

	Trace test		MaxEV test	
H_0	$h = 0$	$h = 1$	$h = 0$	$h = 1$
US	**30.35**	2.64	**27.71**	2.63
Japan	*17.04*	4.05	*13*	4.04
Germany	**18.2**	2.05	**16.15**	2.05
France	13.79	0.64	*13.15*	0.63
Italy	**25.68**	0.05	**25.63**	0.047
UK	**21.25**	4.09	**17.16**	4.09
Canada	10.25	0.44	9.80	0.44
90% crit. val	15.58	6.69	12.78	6.69
95% crit. val.	17.84	8.08	14.6	8.08

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

Table 2: estimates of $\beta' = \left[1, \beta_2\right]$ and tests of $H_0 : \beta_2 = 0$

	US	Japan	Germany	France	Italy	UK	Canada
β_2	-0.003	0.01	-0.084	0.0004	-0.001	0.054	0.015
p-value	0.83	0.46	0.001	0.94	0.83	0.09	0.25

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

	US	Japan	Germany	France	Italy	UK	Canada
US	1						
Japan	-0.1932	1					
Germany	-0.2203	0.2888	1				
France	0.001465	0.2563	-0.1412	1			
Italy	-0.07919	0.2709	-0.06561	0.6595	1		
UK	0.09094	0.1825	-0.4724	0.1099	0.166	1	
Canada	0.1738	-0.2927	0.01252	-0.3498	-0.3039	0.03893	1
mean	-0.03	0.08	-0.09	0.08	0.10	0.019	-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

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

	US	Germany	Japan	France	Italy	UK	Canada
US	1						
Germany	0.4021	1					
Japan	0.2999	0.283	1				
France	0.3714	0.4497	0.3642	1			
Italy	-0.07883	-0.116	0.3682	0.3681	1		
UK	0.2934	0.3706	0.4597	0.4495	0.203	1	
Canada	0.7015	0.161	0.3364	0.4039	0.255	0.5147	1
mean	0.33	0.25	**0.35**	**0.40**	0.16	**0.38**	**0.39**
std-dev.	0.25	0.2098	0.06	0.040	0.21	0.11	0.19
p-value	0.12	0.13	0.001	0.0000	0.23	0.01	0.04

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

Table 4 a): Principal component analysis of global shocks

Principal Comp.	Variance explained	Latent roots	Bartlett Test
1	43.66	3.056	0.000
2	18.46	1.292	0.007
3	13.48	0.9434	0.02
4	9.463	0.6624	0.03
5	8.208	0.5745	0.02
6	4.612	0.3228	0.10
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	0.010
2	23.54	1.648	0.056
3	14.14	0.9901	0.24
4	12.02	0.8413	0.18
5	10.3	0.7211	0.17
6	5.095	0.3566	0.78
7	3.951	0.2766	NaN

Table 5: Tests of the Blanchard-Quah restriction

Test on α							
	US	Japan	Germany	France	Italy	UK	Canada
LR	13.44	9.06	15.2	0.92	15.41	1.06	0.48
p-val.	0.0002	0.002	0.0000	0.33	0.0000	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	0.15	0.000	0.19	0.43

LR is distributed as $\chi^2(1)$ and t-stat as $t(T-5)$ where $T = 32$ is the sample size

Table 6: Share of global shock 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 7: Relative variance of e^c and e^w: estimates of s_{11}/s_{22}.

US	Japan	Germany	France	Italy	UK	Canada	Average
0.3019	0.50	0.63	0.64	1.28	0.59	0.46	0.63

Table 8: 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

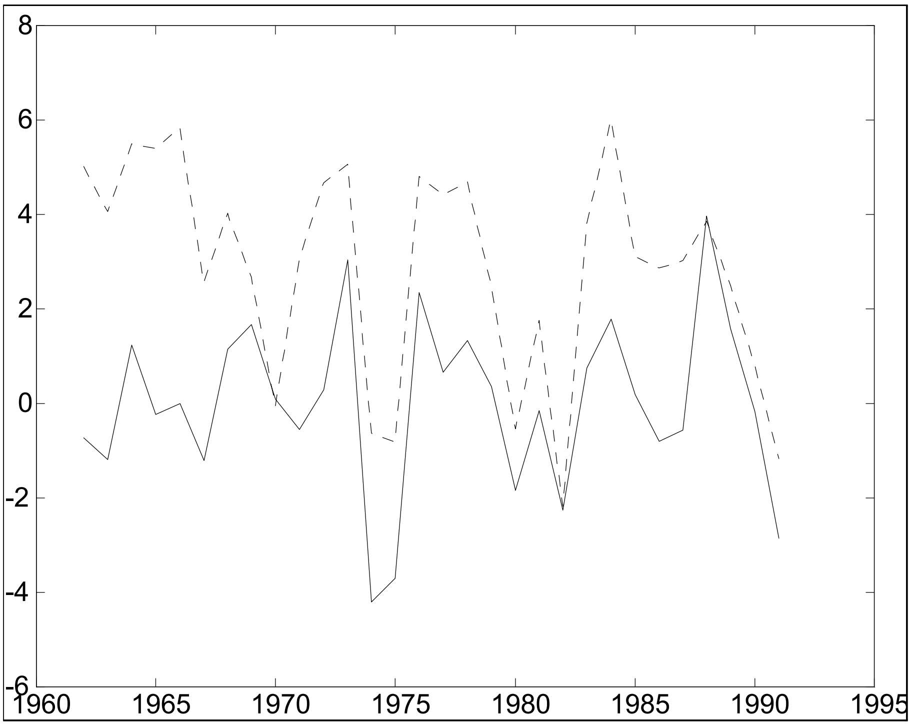

Figure 1: US GDP growth rates and the first principal component of global shocks

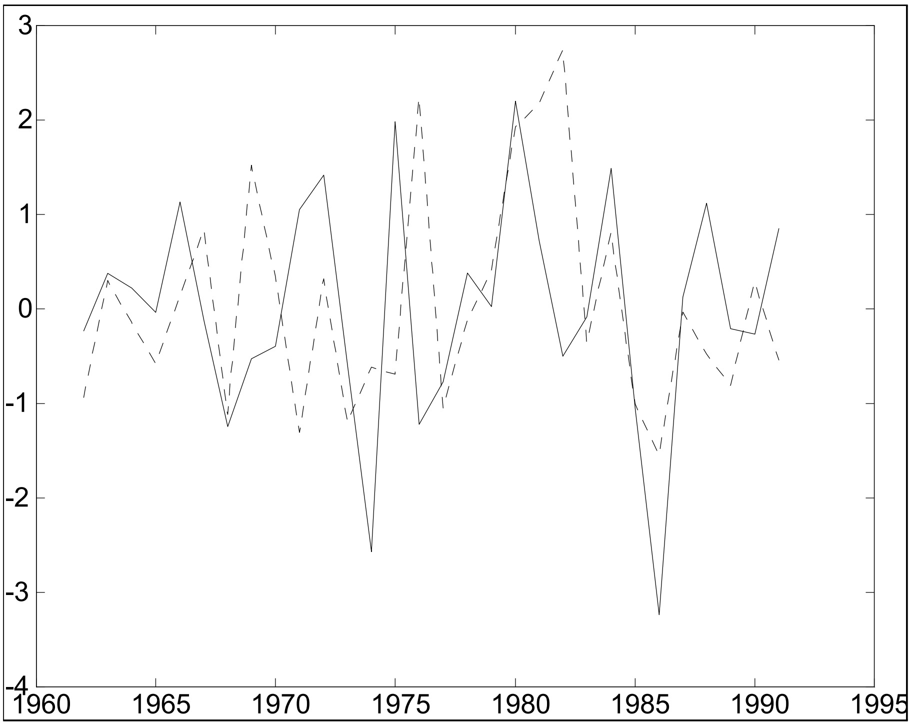

Figure 2: Changes in the U.S. real interest rate (dashed) and second principal component of global shocks