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THE TRANSMISSION MECHANISM IN A CHANGING WORLD

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SUMMARY

The paper aims to identify those factors that cause changes in the speed and strength of the international transmission of output shocks from the USA to specified European economies. These factors are identified through the use of generalized impulse response functions conditioned on histories defined by an abrupt transition VAR. The chosen transition variables comprise changes in exchange rates, financial prices, international capital flows, trade links and monetary policy instruments. Besides the identification of asymmetric responses, the proposed model is useful in analyzing the strong effect of the recent US recession on the European economies and changes in business cycle synchronization over time. Copyright © 2007 John Wiley & Sons, Ltd.

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1. INTRODUCTION

In this paper we propose a model that is able to capture recurrent changes in the transmission mechanism of external shocks. It relies upon a regime-switching mechanism given by variables that measure conditions in the domestic country that suffers the shock, possibly with respect to a benchmark country. These variables include measures of financial markets, financial structure, trade, international capital flows, the labour market and monetary policy. A method of identifying the variable that generates significant differences in the transmission mechanism of the same shock across regimes is proposed. Based on the estimates of the proposed model, we are able to describe the economic conditions that make the economy more susceptible to external shocks. Thus, the critical contribution of the model is to allow the identification of the economic conditions that switch the transmission mechanism, such that the same extraordinary shock can have different effects at different times.

The paper was originally motivated by the experience of the last downturn in the USA, which indicates three aspects to consider when modelling transmission of external shocks to Europe. First, despite some anticipations to the contrary, the European economy was strongly affected by the downturn in the USA. Second, this cyclical sympathy broke a pattern of desynchronization between the USA and the European countries that had held away for most of the previous two decades. Third, this increase in synchronization may be temporary and a result of common shocks affecting these economies. These points are usefully discussed in International Monetary Fund (2001), Doyle and Faust (2002) and Helbling and Bayoumi (2003). Between them, these observations point to

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a need better to understand what links the reactions to shocks in different economies, and how these links may recurrently have changed through time.

Our methodology is based on a VAR approach, more specifically on a nonlinear VAR, through which we are able to implement the regime-switching described above. Most existing work on the international transmission of shocks has either deployed non-VAR approaches, such as cross-correlation analysis and factor analysis, or has concentrated on linear VAR approaches. Among the former group we may cite Doyle and Faust (2002), Helbling and Bayoumi (2003), Bordo and Helbling (2004), Kose *et al*. (2003) and Monfort *et al*. (2004). These papers establish a number of features relevant for our enquiry. For example, Monfort *et al*. (2004) establish a presumption that cross-area spillovers from America to Europe have become stronger in their sample period (which runs to 2002), while the papers by Bordo and Helbling (2004) and by Helbling and Bayoumi (2003) establish that over long periods of time there has been a tendency for business cycle synchronization in downturns to increase; albeit, in the late 1980s and early 1990s there was a period when synchronization fell. Among the group of VAR papers Canova and Marriman (1998) remains a classic early statement of the issues involved in this type of work: interestingly, this study concludes by suggesting that future research—as in this paper—would benefit by introducing additional information variables into the VAR framework. More recent examples of linear-VAR based applications are to be found in Perez *et al*. (2003) and Peersman (2004). The first of these establishes that the magnitude of US shocks has increased relative to that of European ones in recent years, implying a relatively larger impact from moderate US shocks on the European economy than in the past. Peersman's study is interesting for us because it examines, as we do, the early millennium slowdown, but its emphasis, differently from ours, is on using restrictions to achieve identification of shocks. A recent paper that aims to explore the global nature of the dependencies of the world economy by Dees *et al*. (2007) has employed generalized impulse responses for measuring international transmission of shocks, as is also done in this paper.

In addition to the above, there is also an adjacent literature on the international transmission of shocks which has been spawned by the foreign exchange crises in South East Asia in 1997. That literature deals with the transmission of shocks to foreign exchange markets at high frequencies, and much of its concern is with 'contagion' and the differentiation between contagion and interdependence (e.g., see Forbes and Rigobon, 2001; Pesaran and Pick, 2004); by contrast our interest in this paper is with lower-frequency events and with interdependence. Nevertheless, a number of the channels of transmission identified in the contagion literature are relevant also for us.

The paper is structured as follows. In Section 2 we introduce the idea of capturing recurrent changes in shock transmission with a nonlinear VAR in the output growth of the USA, Germany and one of four other European countries (viz. France, Italy, Spain and the UK). The evolution of the parameters of the VAR depends on an observable 'transition' variable, and the parameters can vary either abruptly or smoothly. We discuss in turn model specification, estimation and selection of the transition variable and type of transition function. In Section 3 we use the preferred nonlinear specifications to compute the (generalized) responses to different types of shocks, and compare them with those from linear models. We focus on the existence of statistically significant asymmetries in the response to the same shock, depending on the regime occupied by the relevant transition variable and the frequency of regime change. The proposed model is evaluated in a out-of-sample exercise in Section 4. In Section 5 the nonlinear models are applied to see how far the European growth slowdown of 2001 can be explained as a result of a spillover from a US

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shock and to identify variables that explain increasing synchronization over time. In Section 6 we summarize the main results of the paper and suggest directions for additional research in this area.

2. MODELLING THE CHANGING TRANSMISSION MECHANISM

This section starts by presenting our observed transition (OT) nonlinear VARs, and is followed by the comparison of the proposed model with some nonlinear models recently proposed in the literature. It continues with details on model estimation (Section 2.3), selection of the transition variable (Section 2.4) and the choice of the shape of transition function (Section 2.5).

2.1. The OT-VAR

The observed transition vector autoregressive model (OT-VAR) is able to capture recurrent changes in the mechanism of transmission of an extraordinary external shock. The regime-switching is given by an observable variable that measures the economic conditions of the country that suffers the shock.

The specification of the OT-VAR is based on a trivariate VAR. The model includes the output growth of the US $y_{\mathrm{US},t}$, representing the rest of the world (or a large country), of Germany $y_{\mathrm{ger},t}$, representing the largest economy in Europe, and of another European country $y_{X,t}$, where X is either France, Italy, Spain or the UK. The economies of the USA and Germany are taken as the leading 'anchor' economies which may provide a focal point or attractor for the other European economies under analysis. There is some evidence for our period that most European economies can be thought of as moving from a US sphere to a German one, though the UK is a traditional exception (see, for example, Artis and Zhang, 1997).

The parsimonious description represented by our trivariate VAR necessarily omits other potentially important variables, such as output growth in Japan or measures of fiscal and monetary policy. Larger datasets could be analysed in a different framework, such as the structural factor model of Kapetanious and Marcellino (2003) or the global error-correcting model of Pesaran *et al.* (2004). Yet, enlarging the information set would make modelling non-linearity and time variation very difficult. Therefore we stick to our three variable models, which are sufficient to capture the most relevant interlinkages between the USA and Europe.

Given the dth lag of an observable transition variable z, $d > 1$, the covariance structure of the 3×1 vector y_t evolves according to the model

$$\begin{aligned} y_t = {} & [c_1 + A_{1,1} y_{t-1} + \cdots + A_{1,p} y_{t-p} + \varepsilon_{1,t}](1 - F(z_{t-d})) + \\ & [c_2 + A_{2,1} y_{t-1} + \cdots + A_{2,p} y_{t-p} + \varepsilon_{2,t}] F(z_{t-d}) + \varepsilon_{s,t} \end{aligned} \tag{1}$$

where $A_{s,j}$ is the 3×3 matrix of autoregressive coefficients of regime s ($s = 1, 2$) and lag j ($j = 1, \ldots, p$), c_s is a 3×1 vector of constants of regime s and $\varepsilon_{s,t} \sim N(0, \Sigma_s)$. The value of the transition function $F(z_{t-d})$ at time t is a scalar, implying that the coefficients of each equation change at the same point in time. This restriction ensures that we can obtain changes in the transmission mechanism of shocks from one country (equation) to another without capturing country-specific changes that do not affect the transmission of international shocks.

The transition function assumes values between 0 and 1. A possible specification is the logistic function $F(z_{t-d}) = 1/(1 + \exp(-\gamma(z_{t-d} - r)))$ (Teräsvirta, 1998). An advantage of the logistic

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function is that it nests the constant parameter model and the abrupt transition model depending on the values of the smoothing parameter γ. When $\gamma \longrightarrow 0$, the logistic function goes to 1/2; as a consequence, there is no parameter change. When $\gamma \longrightarrow \infty$, the function converges to an indicator function $F(z_{t-1}) = I(z_{t-d} > r)$, so there is an abrupt switching regime. Nonlinear models with regime switching given by an indicator function are generally called threshold models (Tong, 1990).[1]

The choice between an abrupt and a smooth transition model is mostly an empirical issue. Smooth transition models with logistic functions nest abrupt transition models, but the latter can be more precisely estimated when the restriction holds and the sample is short. We have experimented with both types of models, and have chosen to report results for abrupt transition only since in general they are better able to fit the data, perhaps due to the rather short sample available, 1970 : Q1-2001 : Q1. More details are presented in section 2.5.

An important assumption in our analysis is that the transition variable, together with the threshold and the delay, generates changes in the transmission mechanism of shocks but does not predict nor is predicted by the domestic business cycle. This hypothesis is well supported by the forecast results presented in Section 3.5. In addition, possible endogeneity of the transition variable is not an issue when the variable is entered, as here, in lagged form.

2.2. Comparing OT-VARs with Other Modelling Approaches

In contrast to structural VARs (Kwark, 1999) and models that decompose shocks into different components (Monfort *et al.*, 2004), the model proposed here characterizes differences in the effect of shocks based on the values of identified observed transition variables, so that shocks of the same size can have different effects depending on the regime occupied by these variables. We will refer to these differences as 'asymmetries'. The asymmetries in the transmission of shocks are captured by recurrent regimes that are determined by the value of an observed transition variable and an estimated transition function. This approach contrasts with the use of models to measure changes in the correlation across shocks assuming that the regimes are known (Favero and Giavazzi, 2002). Our method also differs from the adoption of models where trade variables determine weights to filter the impact of the shocks (Dassel, 2002; Abeysinghe and Forbes, 2001).

Within the class of regime-switching models, the switching, instead of being determined by an observable variable as in the OT-VARs, can be determined by an unobservable discrete variable (see Hamilton, 1989, and for example, Artis *et al.*, 2004, and Sims and Zha, 2006, for recent applications to Europe and the USA). Empirically, the practical relevance of switching in the parameters when their evolution is determined by an unobserved Markov process with constant probability of transition is limited; see, for example, Hamilton and Perez-Quiros (1996) or Sims and Zha (2006). It increases either when the transition probabilities are time-varying, as, for example, in Filardo (1994), Diebold *et al.* (1994), or when the evolution of the parameters follows random walk processes, as, for example, in Cogley and Sargent (2005), or when the time variation is driven by an observable continuous variable, as in this paper.

Smooth transitions with observable variables can be also obtained by including an error term (independent of the shocks) in an abrupt transition function (mixture models; see Lanne, 2006).

[1] A related model is the endogenous delay threshold specification of Koop and Potter (2003). Yet, this specification imposes restrictions that can be inappropriate in our context, where the interest focuses on the dynamic transmission of shocks.

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However, because in mixture models at each point in time the probability of being in a given regime depends also on an unobservable, it is harder to identify asymmetry in the responses to shocks given the domestic conditions measured by the transition variable.

Another important feature is that the OT-VARs are able to capture time-varying behaviour as well as recurrent regime behaviour (Carrasco, 2002). Depending on the combination of transition variable and transition function, the regimes may not be frequently recurrent, so that the model encompasses models with breaks. In addition, the model is also able to capture heteroscedasticity because we allow the variance of the error term to vary across regimes ($\varepsilon_{s,t} \sim N(0, \Sigma_s)$). This is relevant in view of the debate between Sims and Zha (2006) and Cogley and Sargent (2005) on the relative role of switches in the dynamics and in the variance.

2.3. Estimation

Estimation of OT-VARs assuming that the transition variable is known depends partially on the supposition about the shape of the transition function. Threshold VARs have been estimated by conditional least squares (Tsay, 1998) and by maximum likelihood (Hansen and Seo, 2002). Smooth transition VARs are normally estimated by nonlinear least squares (Van Dijk *et al.*, 2002), but there is at least one example in which maximum likelihood has been employed (Camacho and Perez-Quiros, 2002). Because the smooth transition (logistic) is a differentiable function, usual consistency properties are available for nonlinear least squares applied to models with smooth transition. In the case of VARs with abrupt transition, Tsay (1998) has proved consistency by conditional least squares. The simulations in Galvão (2006) show that both conditional least squares and maximum likelihood are adequate for estimating threshold VARs even in relatively small samples.

In this paper, we use the conditional maximum likelihood concentrated with respect to the parameters in the transition function to estimate OT-VARs with abrupt and smooth transitions. Conditional on the parameters of the transition function, the coefficients c_1, c_2, $A_{s,i}$ are obtained using least squares. The parameters that will be estimated in the smooth and abrupt transitions are, respectively, $\theta_{\mathrm{ST}} = (\gamma, r)$ and $\theta_{\mathrm{AT}} = (r, d)$. In both cases, the autoregressive order p is chosen by comparing the Schwarz information criterion of OT-VARs with autoregressive orders from 1 to **4** (p max). The delay in the smooth transition is chosen using the results of a linearity test for $d = 1, \ldots, 8$ (d max) before the estimation of the system (for details on this specification procedure, see Van Dijk *et al.*, 2002). The maximum likelihood estimators can be obtained by maximizing the following function:

$$\max -\frac{n}{2}\log(\det(\hat{\Sigma}(\theta)))$$

where $\hat{\Sigma}(\theta) = 1/T\sum_{i=1}^{T}\widehat{\varepsilon}_t(\theta)\widehat{\varepsilon}'_t(\theta)$. In the case of a smooth transition specification $\theta = \theta_{\mathrm{ST}}$, the parameters are obtained using numerical optimization. Initial values for the parameters are obtained by a grid search over possible values for the smooth parameter (γ) and the threshold (r). In the numerical optimization, the parameter space of r is constrained such that if γ is large we guarantee that we still have enough data information to estimate the autoregressive parameters.

A constraint that at least 10% of the observations are in each regime is employed to define the limits of the grid to obtain the maximum likelihood estimates for the OT-VAR with abrupt transition. The limits of the grid for the delay are also the same as those used in the smooth

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transition specification ($d\max = 8$). Specifically in the case of the abrupt transition OT-VARs, the maximization problem can be written as follows:

$$\hat{\theta}_{\mathrm{AT}} = (\hat{r}, \hat{d}) = \max_{\substack{r_l \leq r \leq r_u \\ 1 \leq d \leq d\max}} -\frac{n}{2} \log(\det(\hat{\Sigma}(\theta_{\mathrm{AT}})))$$

Based on $\hat{\theta}_{AT}$, one can divide the sample, so that the variance matrix of residuals conditional on the regime $\hat{\Sigma}_s(\hat{\theta}_{AT})$ is obtained. The fact that heteroscedasticity is not taken into account in the estimation does not affect the unbiasedness of the estimator with abrupt transition (see Galvão, 2006).

2.4. Selection of Transition Variables

Table I lists a set of candidate variables that were evaluated as possible transition variables in this study; a detailed data Appendix is available upon request.

Previous work has identified a number of variables that can influence or mediate the transmission of a shock in one country to another. In particular, monetary policy has been identified as capable of mediating to an important extent the transmission of external shocks to a domestic economy (e.g., Monfort *et al*., 2004); direct spillovers from external equity price shocks to domestic markets may also be important (e.g., Kose *et al*., 2004), while the contagion literature has identified

Table I. List of possible transition variables

1 Monetary policy instruments
a RST: short-term real interest rate
b ST: short-term interest rate
c DM2SA: Broad money M2
2 Financial variables
a DREER: real effective exchange rate
b D_ERE: exchange rate (US $)
c RLT: long-term real interest rates
d LT: long-term interest rate
e DSPI1(2): share prices
3 International financial flows
a BIN: net bank international deposits/GDP
b TNET: total net international flows/GDP
c BDBAL: external assets/liabilities of German banks
d USBCL: external assets/liabilities of US banks
e USBGR: Purchase of US bonds/total European purchase of US bonds
f USSTGR: Purchase of US corporates/total European purchase of US corporates
4 Trade
a TOT: terms of trade
b TRADE: trade with the chosen countries as percentage of total trade
5 Natural resources
a OLIMP: oil imports/GDP
b NETIMP: net mineral fuels/GDP
6 Labour market
a UNEM: unemployment rate

Note: The full description of the dataset including the construction of these variables is available from the authors on request.

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that exchange rate movements (specifically, devaluations) may provoke financial fragility and precipitate stronger effects of a foreign downturn as a result. Monetary variables and exchange rates also underlie the Exchange Market Pressure variables developed by Eichengreen *et al.* (1996) and often used in studies on the transmission of financial market crises. These considerations lead to a case for including monetary policy variables and asset prices (and their change) in the set of possible transition variables. The specific ones we chose to examine appear in groups 1 and 2 of Table I.[2]

Other papers (notably Imbs, 2004, but also Kose *et al.*, 2004, among others) have highlighted the role of capital movements and some alternative measures of these flows or indicators of vulnerability to flows (such as high values of external liability/asset ratios) are shown in the third group in Table I.

The terms of trade and, especially, measures of bilateral trade intensity are common fare in studies of international cyclical synchronization, stimulated not least by the seminal paper of Frankel and Rose (1998), even though there are now some doubts about their relevance when compared with the role of financial market linkages; see, for example, Van Rijikghem and Weder (2001). These measures are shown in group 4 of the table. Even more commonly cited in related studies (see Monfort *et al.*, 2004, for a recent example) is the role of oil prices and oil dependency, measures of which are shown in group 5 of the table.

Some studies of international business cycle synchronization have also employed measures of labour market rigidity (e.g., Artis, 2004), on the basis that such measures convey important information about the propagation mechanism attending the shock; here the actual level of unemployment is used, which may be strongly related to the natural rate and hence benefit from this argument—or this variable may simply stand in for a more general measure of the economy's output gap and hence that economy's ability to absorb a shock from outside.

One of the criteria for choice of variable in this study has to be the consideration that it should have displayed some variability over time in the sample period; otherwise it would not be capable of being identified as a factor which could underpin a change in the value of shock transmission.[3]

Though the list of variables in Table I is not exhaustive, it provides quite an extensive range, definitely larger than anything evaluated in the literature so far, to the best of our knowledge. The use of other indicators is made difficult by their non-availability for all countries with comparable definitions for a long enough period of time (most of our indicators are collected for the period 1970 : Q1–2001 : Q1).

Since all the variables in Table I can be given an interesting economic interpretation, we have to use statistical criteria for selecting those indicators most relevant to explain changes in shock transmission in our context.[4] Following common practice in the specification of observable transition models (e.g., Teräsvirta, 1998), the first step in the selection of the transition variable is the use of a variable addition test to verify whether adding $(y_{t-1}\ y_{t-2}\ \ldots\ y_{t-p})z_{t-d}$ significantly improves the fit of the linear VAR. Although this variable addition test has been proposed by Teräsvirta in the context of testing for linearity against the smooth transition alternative, it has power against the threshold alternative, as shown by Strikholm and Teräsvirta (2005) in

[2] The monetary policy variables are transformed before being employed as transition variables as the difference with respect to the German ones (in the case of Germany, the differences are with respect to the UK).

[3] Note that this is a different question from the one asked in some related studies (e.g., Artis, 2004) where the objective is to tease out the differences between countries in their reaction to a foreign shock—this is a cross-sectional question rather than a time series question such as the one studied here.

[4] All transition variables are standardized before selection and estimation.

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an experiment that employs this type of test to select the number of regimes in a threshold autoregressive model.

LR tests for the significance of the additional regressors are computed for each combination of trivariate VAR specification and transition variable in the starting set. Moreover, for each transition variable, the LR test is computed for $d = 1, \ldots, 8$. The transition variable z_t is retained if the null hypothesis of non-significance of the added regressors is rejected at the 10% level for at least one of the possible delays.

In a second step, for each of the transition variables retained from the previous step, OT-VARs are estimated as detailed in the previous subsection, and ranked on the basis of the Schwarz information criterion. Information criteria have been suggested to select threshold specifications by Tsay (1998), and criteria with heavier penalties can perform better (see Gonzalo and Pitarakis, 2002).

In order to reduce the computational time, in the following analyses we consider the transition variables associated with the best 50% of models. These are listed in the first column of Table II.

Three points are worth noting about the selected transition variables. First, at least one measure of monetary policy differential is retained for each country analysed. Second, the choice of the financial variables indicates two important markets for the transmission of shocks: currency and stocks. Some of these variables are close to those selected in the contagion literature, though measured at lower frequency. As noted earlier, however—and to use the felicitous distinction emphasized by Pesaran and Pick (2004)—this study is concerned with interdependence rather than contagion. Finally, looking at non-financial variables, there are similarities in some of the transition variables for Italy and Spain, in contrast to those for the UK and France. For the former countries, variables related to trade and labour market were chosen, while for the latter countries variables that measure the need for oil imports were selected.

2.5. Selection of the Shape of the Transition Function

In this subsection, we compare the fit of OT-VARs with smooth and abrupt transition. If OT-VARs with abrupt transition (OAT-VAR) fit the data dynamics as well as smooth transition models (OST-VAR), then it is better to use only OAT-VARs for two reasons. First, large values of the smoothing parameter γ affect the precision of the estimates of r and γ, since few observations around the threshold value are available given the short sample at our disposal (Teräsvirta, 1998). Second, since possible asymmetries in the transmission of shocks will be evaluated on the basis of generalized impulse response functions, the use of abrupt transition models is economically meaningful because the sample is clearly divided into two sets of histories, determined by the economic conditions represented by the transition variable.

To compare the fit of the competing specifications we employ a bounded Wald statistic. The latter was proposed by Altissimo and Corradi (2002) as a consistent decision rule to reject the null of no threshold effects, avoiding the problem of data-dependent p-values for sup Wald tests. Moreover, the statistic can also be employed as a consistent information criterion.

The single equation-bounded Wald statistics are written as

$$\mathrm{BW_{AT}} = \left[\frac{1}{2\ln(\ln(T))}\left[\sup_{\theta^{\mathrm{L}}_{\mathrm{AT}} \leq \theta_{\mathrm{AT}} \leq \theta^{\mathrm{L}}_{\mathrm{AT}}} T\left(\frac{\mathrm{SSR}(\hat{\theta}_{\mathrm{L}}) - \mathrm{SSR}(\theta_{\mathrm{AT}})}{\mathrm{SSR}(\theta_{\mathrm{AT}})}\right)\right]^{1/2}\right]$$

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Table II. Evidence of asymmetries of the generalized impulse responses from OAT-VARs

Affected country/ transition variable of the OAT-VAR	Asymmetric responses across regimes?	Infrequent regime switching?	US and BD shocks have opposite signs?	Estimated transition function	Regime with larger response to US shock	Regime with larger response to BD shock	Regime with larger response to a common shock
France							
RST	√			$I(z_{t-1} > 0.381)$	2nd (0.47) [1.7]	2nd (0.23) [2.22]	2nd (1.43) [1.60]
D_ERE							
DSPI							
DM2SA							
NETIMP							
Italy							
DREER	√			$I(z_{t-6} > -0.54)$	1st (1.42) [1.67]	2nd (0.92) [1.83]	1st (3.19) [1.45]
DSPI	√			$I(z_{t-7} > 0.006)$	2nd (0.91) [4.91]	1st (0.51) [1.08]	1st (2.72) [1.43]
BIN	√			$I(z_{t-4} > 0.899)$	2nd (1.48) [3.40]	2nd (1.55) [3.08]	2nd (3.50) [2.40]
TRADE	√	√		$I(z_{t-8} > 1.124)$	1st (0.38) [1.42]	1st (0.97) [1.18]	1st (1.78) [1.16]
ST							
UNEM	√	√		$I(z_{t-6} > 1.092)$	1st (1.54) [2.84]	1st (1.40) [2.73]	1st (3.40) [2.30]
Spain							
DREER	√			$I(z_{t-6} > 0.780)$	2nd (2.00) [3.44]	2nd (0.60) [2.15]	1st (1.49) [1.27]
DSPI	√			$I(z_{t-1} > -0.467)$	1st (2.44) [5.15]	2nd (0.79) [1.26]	2nd (2.25) [1.48]
TOT							
DM2SA	√			$I(z_{t-1} > -0.506)$	2nd (0.99) [1.8]	2nd (0.45) [1.36]	1st (3.08) [1.71]
BIN	√	√		$I(z_{t-4} > 0.060)$	2nd (1.23) [4.44]	2nd (0.67) [1.79]	2nd (2.28) [1.82]
TRADE	√	√		$I(z_{t-4} > 0.561)$	2nd (0.88) [8.26]	1st (0.49) [1.01]	2nd (1.99) [2.28]
UNEM							
UK							
RST							
DREER			√	$I(z_{t-2} > 0.650)$	2nd (0.80) [2.29]	1st (0.38) [2.14]	1st (1.44) [1.01]
USBCL			√	$I(z_{t-3} > 0.617)$	1st (1.05) [3.31]	2nd (0.44) [4.01]	1st (1.52) [18.40]
ST	√	√		$I(z_{t-4} > 0.651)$	1st (1.19) [2.13]	1st (1.40) [4.50]	1st (3.14) [1.98]
NETIMP			√	$I(z_{t-6} > 0.693)$	1st (1.04) [1.87]	2nd (0.75) [6.53]	1st (1.84) [1.11]
Germany							
RST							
D_ERE							
TNET	√			$I(z_{t-1} > -0.102)$	1st (1.09) [1.51]		1st (2.41) [1.67]
USGBR							
ST	√	√		$I(z_{t-4} > 0.522)$	2nd (1.44) [6.51]		2nd (2.91) [2.42]

Note: 'Asymmetric responses across regimes' means that the accumulated response after 4 quarters from the common shock is statistically different across regimes. 'Infrequent regime switching' means that the frequency of switching over the sample is smaller than 15%. 'US and BD shocks have opposite signs' means that the accumulated response after 4 quarters is positive or negative depending on the origin of the shock being the USA or Germany (BD). Values in parentheses in the last three columns are the accumulated response to the shock after 4 quarters, and the values in square brackets are the ratios of the responses between the regime with larger response and the regime with smaller response.

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Table III. Comparison of BW rules of OT-VARs with smooth and abrupt transitions

ot-var	France			Italy			Spain			UK			Germany		
	T	BW_{AT}	BW_{ST}	T	BW_{AT}	BW_{ST}	T	BW_{AT}	BW_{ST}	T	BW_{AT}	BW_{ST}	T	BW_{AT}	BW_{ST}
RST	116	**1.398**	1.188							117	1.416	**1.65**	117	**1.031**	0.805
DREER				116	**1.358**	1.179	116	0.303	**1.248**	108	**2.831**	2.829			
D_ERE	116	**0.593**	0.593										116	**1.668**	0.793
DSPI1	116	**1.264**	1.25	116	0.742	**1.018**	116	**1.393**	1.312						
TOT							116	0.717	**1.559**						
DM2SA	116	0.714	**0.94**				116	1.289	**1.381**						
BIN%				116	**1.772**	1.683	116	1.629	**1.663**						
TNET%													113	**1.259**	1.019
TRADE				116	0.663	**1.041**	116	1.29	**1.358**						
USBCL										84	5.461	**5.83**			
USSTGR													89	**2.16**	2.015
ST				113	**1.173**	1.062				117	**1.552**	1.332	117	**1.12**	0.88
NETIMP	116	**0.615**	0.613							117	1.7	1.347			
UNEM				116	**1.408**	0.933	107	**2.096**	2.096						

Note: Emboldened entries mean that the model fits better with the chosen transition variable than the competitor specification. T is the number of observations effectively employed in the estimation.

$$\mathrm{BW}_{\mathrm{ST}} = \left[\frac{1}{2\ln(\ln(T))} \left[\sup_{\theta_{\mathrm{ST}}^{\mathrm{L}} \leq \theta_{\mathrm{ST}} \leq \theta_{\mathrm{ST}}^{\mathrm{L}}} T \left(\frac{\mathrm{SSR}(\hat{\theta}_{\mathrm{L}}) - \mathrm{SSR}(\theta_{\mathrm{ST}})}{\mathrm{SSR}(\theta_{\mathrm{ST}})} \right) \right]^{1/2} \right]$$

where SSR(.) is the sum of squared residuals of one of the equations of the estimated VAR. The null of linearity is rejected in favour of abrupt transition if $BW_{AT} > 1$ and of smooth transition if $BW_{ST} > 1$. Similarly, abrupt transition is better than smooth transition if $BW_{AT} > BW_{ST}$.[5]

Table III presents the statistics using the SSR of the equation in the VAR of the country that suffers the shock. Qualitative results do not change if the BW_{AT} and the BW_{ST} are computed for other equations. Table III shows that for virtually all transition variables the null of linearity is rejected, and abrupt transition is better than smooth transition non-linearity for the majority of the models. Therefore, only OAT-VARs are applied to measure asymmetries in the transmission mechanism in the next section.

3. OAT-VAR AND THE INTERNATIONAL TRANSMISSION OF SHOCKS

We start this section with a discussion of alternative definitions of shocks. Then we indicate how to calculate generalized responses to these shocks in our context (Section 3.2). Finally, we compute the generalized responses using standard linear VARs (Section 3.3), and compare them with the shock responses from our OAT-VARs (Section 3.4). This allows us to identify asymmetries in the expected response to shocks conditional on the economic conditions of the country that suffers the shock.

[5] The parameters that maximize the Wald statistics are not exactly the same as the ones that maximize the likelihood as described in Section 2.3. However, the estimates that minimize the *trace* $(\hat{\Sigma}(\theta))$ do maximize the Wald statistic. So we re-estimate OAT-VARs and OST-VARs for the chosen transition variables using this sum of squared residuals as the objective function. Only small differences in these new estimates were found when compared with the maximum likehood ones.

3.1. Types of Shocks

We consider the transmission of three different types of shocks, v_i. Pure idiosyncratic shocks (PIS) have no contemporaneous effects on other countries. For example, when the source of the shock is the USA, a PIS is defined as the vector (0, 0, 1). This notation means that this quarter the USA grows, say, at a 4% quarterly rate rather than 3%, while the other countries are unaffected.

Idiosyncratic shocks with spillovers (SIS) originate in a single country but can have contemporaneous effects on the other countries, as measured by the covariance matrix of the VAR residuals. For example, in the case of a shock from the USA, we define the SIS as the vector $(\sigma_{13}/\sqrt{\sigma_{11}}\sqrt{\sigma_{33}}, \sigma_{23}/\sqrt{\sigma_{22}}\sqrt{\sigma_{33}}, 1)$, where σ_{ij} is the element in the ith row and jth column of the covariance matrix of the residuals in the relevant VAR, and the division by $\sqrt{\sigma_{33}}$ is used to make the size of the shocks comparable across countries. Notice that in practice this definition implies that the size of the spillover is equal to the cross-country correlation of the estimated VAR residuals.

Other definitions of SIS shocks are possible; for example, σ_{13}/σ_{33} would be equivalent to the coefficient in a linear regression of country one residual on the US residual, while the shock defined as

$$\left(\text{2nd row of} \begin{pmatrix} \sigma_{22} & \sigma_{23} \\ \sigma_{32} & \sigma_{33} \end{pmatrix}^{-1} \begin{pmatrix} \sigma_{12} \\ \sigma_{13} \end{pmatrix}, \frac{\sigma_{23}}{\sigma_{33}}, 1\right)$$

would correspond to a Choleski ordering USA–Germany–other country. The latter is basically the choice made by Favero and Giavazzi (2002) in their analysis of the transmission of financial shocks.

It is not clear on *a priori* grounds which definition is preferable, but all of them would be equal if the spillover effect was equal to zero, that is, if $\sigma_{13} = 0$ and $\sigma_{23} = 0$. Section 3.3 indicates that, for all countries, the responses to shocks originating in the USA based on linear VARs are not statistically different when using the definition of shock with (SIS) and without spillovers (PIS). As a consequence, the exact definition of a shock with spillover effects is not critical to the results of this paper.

Finally, a common shock (COS) is represented by a one-standard-deviation shock in each of the countries under analysis with no contemporaneous effects across countries. In particular, a COS is defined in vector notation as $(\sqrt{\sigma_{11}}/\sqrt{\sigma_{33}}, \sqrt{\sigma_{22}}/\sqrt{\sigma_{33}}, 1)$, where the shocks are again normalized by $\sqrt{\sigma_{33}}$.

3.2. Generalized Impulse Responses

The responses to shocks in nonlinear models can depend on the definition of the shock, the history of the system before the shock, and the shocks that are assumed to be hitting from $t+1$ to $t+n$, where n is the maximum horizon under consideration. The concept of generalized impulse responses (Koop *et al.*, 1996) allows us to construct the time profile of shocks conditional on a specific set of history and type of shock, and it assumes that 'normal shocks' (i.e., the average of past shocks) keep hitting the system over future horizons.

Previous papers using generalized responses (Koop *et al.*, 1996; Altissimo and Violante, 2001) did not compute confidence intervals around the estimated responses. In this paper we also propose a method to compute 95% confidence intervals by bootstrap. It is a highly computationally intensive procedure, but it is important for assessing the relevance of an asymmetric transmission mechanism.

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The transmission of the shocks v is computed as the difference between the expected value of y with and without the shocks. Therefore the generalized impulse response (GI) of the series in y_t to the shock v_i at horizon n conditional on the history Ω^s_{t-1} is defined as follows:

$$\mathrm{GI_Y}(n, v_i, \Omega^s_{t-1}) = E[y_{t+n}|v_i, \Omega^s_{t-1}] - E[y_{t+n}|\Omega^s_{t-1}] \tag{2}$$

Based on this concept of impulse response, the 'extraordinary' shock v_i (extraordinary because we are assuming that 'normal' or 'average' shocks hit the system from $t+1$ up to $t+n$) has different effects depending on the history $s = 1, 2$ at the moment that the shock hits the system. The two sets of histories are determined by the value of the transition function at $t-1$, but it is possible to hypothesize regime changes from t to $t+n$. In this set-up, there is a significant difference in the transmission of shocks conditional on the regime if it is possible to conclude that $\mathrm{GI_Y}(n, v_i, \Omega^1_{t-1}) \neq \mathrm{GI_Y}(n, v_i, \Omega^2_{t-1})$ based on the confidence intervals for the responses computed by bootstrap.

The generalized impulse responses (GI) are computed based on an estimated VAR or OAT-VAR (equation 1). The past values of the vector of endogenous variables y_t and of the transition variable z_t are written as $W_{t-1} = (y_{t-1}, \ldots, y_{t-p})$ and $Z_{t-1} = (z_{t-1}, \ldots, z_{t-d_{\max}})$ and are used to build a matrix of histories $\Omega_{t-1} = ((W_{t-1}, \ldots, W_{T-1})', (Z_{t-1}, \ldots, Z_{T-1})')$. This matrix of histories is partitioned to obtain GIs conditional on the regime when calculating GIs for the OAT-VARs, although the history does not affect responses in the case of VARs. In the earlier case, the first partition Ω^1_{t-1} has the rows of Ω_{t-1} such that $F(z_{t-d}) = 0$ and the second partition Ω^2_{t-1} has the rows of Ω_{t-1} such that $F(z_{t-d}) = 1$.

GIs are computed for five types of shock, so that the 5×3 matrix of shocks is $v = (v_1, \ldots, v_5)'$, where v_1 and v_2 are PIS and SIS with origin in the USA; v_3 and v_4 are the same type of shocks with origin in Germany; and v_5 is COS, standardized by the US values. We build GIs conditional on each of these five combinations of type and origin of shock.

The responses from the OAT-VAR for horizons larger than $t+d$ depend on predictions of z_t. Thus we use data simulated from an AR(p) of z_t to obtain a sequence of values of the transition variables $z_t, \ldots, z_{t+n}$. The autoregressive order of the AR for the full sample is obtained by minimization of the SIC. The residuals η_t of the AR(p) for z_t are saved to be used in the calculation of the conditional means of the GIs.

The algorithm employed to obtain $\mathrm{GI_Y}(n, v_i, \Omega^s_{t-1})$, which is the GI at horizon n conditional on a type of shock in v and one subset of histories of Ω_{t-1}, requires the following steps:

(1) Pick one row of shocks from the matrix v_i and pick one of the subsets Ω^s_{t-1} of the matrix Ω_{t-1}.
(2) Pick one of the rows of $\Omega^s_{i,t-1}$.
(3) Use these vectors to compute $y_t^{ss,m} = f(\Omega^s_{i,t-1}, \theta) + v_i$, where θ is a vector with all the estimated parameters of the model. This calculates the impact of the shock.
(4) Draw a subsample ε^* of size $n+1$ by bootstrapping from the VAR residuals ε. When calculating GIs for the OAT-VAR, draw also a subsample $\boldsymbol{\eta}^*$ of size $n+1$ by bootstrapping from $\boldsymbol{\eta}$, which are the residuals of the AR(p) of z_t.
(5) Use $\boldsymbol{\eta}^*$ to get a sequence $z_{t+0}, \ldots, z_{t+n}$ given the estimated AR(p) conditional on Z_{t-1}. Use ε^*, the sequence $z_{t+0}, \ldots, z_{t+n}$, and the estimated VAR to get $y_t^{ns,m}, \ldots, y_{t+n}^{ns,m}$. This calculates a sequence that describes the dynamics of the system when there is no shock.

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(6) Use $z_{t+0}, \ldots, z_{t+n}$, $y_t^{ss,m}$, the first N observations of ε^* and the estimated VAR to get $y_{t+1}^{ss,m}, \ldots, y_{t+n}^{ss,m}$. This computes the dynamic effect of the shock.

(7) Repeat steps 3–6 M times (800 in our plots). Thus, obtain $E[y_{t+n}|\Omega_{i,t-1}^s, v_i] = \frac{1}{M}\sum_{m=1}^{M} y_{t+n}^{ss,m}$ and $E[y_{t+n}|\Omega_{i,t-1}^s] = \frac{1}{M}\sum_{m=1}^{M} y_{t+n}^{ns,m}$. In this way, steps 3–6 are aimed at calculating the conditional expectations. Note that this could be calculated analytically for the linear VAR but not for the OAT-VAR. We use the same algorithm for both cases.

(8) Pick another row of $\Omega_{j,t-1}^s$ and repeat the procedure from 3 to 7 until all rows are considered.

(9) Average the conditional means over histories to get $E[y_{t+n}|\Omega_{t-1}^s, v_i]$ and $E[y_{t+n}|\Omega_{t-1}^s]$, so $\mathrm{GI_Y}(n, v_i, \Omega_{t-1}^s) = E[y_{t+n}|\Omega_{t-1}^s, v_i] - E[y_{t+n}|\Omega_{t-1}^s]$.

(10) Select another combination of shock in v_i and subset of histories Ω_{t-1}^s and repeat steps 2–9 until all possibilities are exhausted. This will generate a set of different GIs conditional on the shock and the set of histories.

The information on the regime-dependent covariance matrices of the OAT-VAR is employed in the computation of the GIs, so the subsamples of ε are taken also conditional on the history/regime.

Because of parameter uncertainty and finite sample size, inference on the GIs for each horizon is based on a 95% bootstrap confidence interval. The distribution of the GI values for each horizon, conditional on the same set of histories and type of shock, is built by simulating R samples of size T using the estimated parameters and bootstraps from the residuals. Then these samples are employed to re-estimate the model and to recalculate the GIs using the described algorithm. Because this procedure is heavily computer-intensive, we use $R = 200$ and $M = 400$.

3.3. Shock Transmission from Linear VARs

Using GDP growth data from 1970 : 1 to 2001 : 1, we estimate linear trivariate VARs to be used as benchmarks for the evaluation of the effects of external shocks in France, Italy, Spain, the UK and Germany. The autoregressive order of the VAR is selected by the Schwarz information criterion, as for the nonlinear VARs. The generalized impulse responses are calculated for all shocks using the computational procedure described in the previous subsection. The accumulated responses for horizons from $n = 0$ to $n = 8$ (quarters) are presented in Figure 1.

Three main findings emerge. First, for all countries a COS shock produces the largest effects at each horizon, followed by SIS and PIS. Yet the confidence intervals around the responses are large enough for the difference between SIS and PIS not to be statistically significant (results available upon request). This is in agreement with the small estimated covariances among the residuals of the three VAR equations, which makes the PIS and SIS shocks very similar.

Second, the accumulated responses to shocks are typically concave, with most of the effects taking place within one year. Ballabriga *et al.* (1999) obtain similar results for US shocks. Moreover, the pattern of response is rather similar across countries, though the COS shocks have larger effects in Spain and Italy at longer horizons. The rather large and slow response of Italy to US shocks is also present in the global model of Pesaran *et al.* (2004).

Third, German shocks have generally lower effects than US shocks.

Although these results are interesting and sensible from an economic point of view, we have to evaluate whether they are confirmed also when the transmission mechanism is allowed to change over time.

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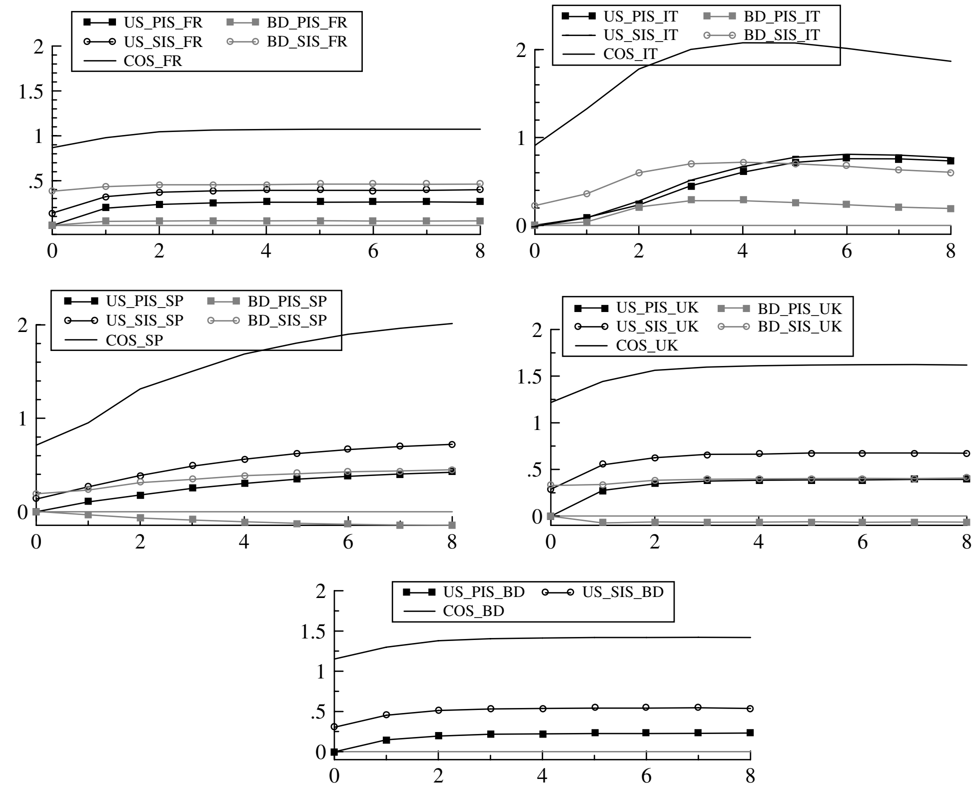

Figure 1. Accumulated generalized responses for France, Italy, Spain, the UK and Germany to PIS and SIS shocks from USA and Germany (BD) and to a common shock (COS) computed using trivariate VARs

3.4. Asymmetries in the Transmission Mechanism

Based on the 95% confidence intervals computed by bootstrap for the response accumulated after 4 quarters, we decided whether the responses are statistically different across regimes. Figures of the accumulated responses and confidence intervals are not presented owing to space restrictions but are available on request. Table II reports the OAT-VARs with the respective transition variables that present asymmetric responses. For these specifications, we also present the estimated transition function and the accumulated response after 4 quarters for the regime with the largest response to US (SIS), German (SIS) and common shocks. The values in square brackets show the ratio of the response between the regime with larger response and the regime with smaller response. Based on the values of the accumulated responses for each variable chosen in the two-step procedure for all shocks, we also indicate in Table II OAT-VARs that in a given regime present responses with opposite signs depending on the origin of the shocks.

The evidence of asymmetric response across histories is stronger for the OAT-VARs of Italy and Spain, which are countries that have experienced stronger economic changes in the last 30 years. Interestingly, we found for those countries transition variables that imply frequent regime changes connected to financial markets (exchange rates (DREER) and stock prices (DSPI)) and variables that imply infrequent regime changes related to increasing trade (TRADE) and financial

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interlinkages (BIN). A summary column showing the frequency of regime switching ('infrequent' regime switching is defined as a frequency of less than 15%) appears in Table II. It turns out that the volatility of the variables triggering regime change, perhaps not surprisingly, can be related to the frequency of the regime change itself: thus financial variables are generally highly volatile and associated with frequent regime change, while less volatile variables, such as trade, trigger less frequent regime change. That the transmission mechanism of an external shock should depend on the variables identified here does not seem surprising; the level of trade between two countries offers an obvious channel for the international propagation of a shock, while the strength of financial interlinkage similarly offers a channel for the magnification of a shock. The sensitivity of the transmission mechanism to changes in exchange rates and stock prices seems likely to be associated, at the broadest level, with the effects of asset price change on liquidity and financial wealth impacting the transmission mechanism via the consequent effects on both consumption and investment. A striking aspect of the UK responses is the evidence of an effect of different sign as between US and German shocks in certain circumstances.[6]

Focusing on the different countries, the evidence of asymmetries across histories for France is rather weak. Yet, observing the OAT-VARs for RST (difference between French and German real short-term interest rates), we can infer that the transmission mechanism of German shocks depends on the divergence of the French monetary policy with respect to Germany: when interest rates are too high, France is more susceptible to German shocks. Similarly, the ST transition variable for the UK shows an interesting monetary policy behaviour: when the interest rate is too low compared with that in Germany one year previously, the economy today is more fragile with strong responses to shocks. When observing the transition function of this variable over time (figure available on request), the UK is being more susceptible to external shocks at the end of the sample than at the end of the 1980s. Likewise, when using ST as transition variable for Germany (with the same endogenous variables as the UK VAR), the domestic economy is more fragile to external shocks when the interest rates are considerably higher than in the UK. In addition, as for Italy and Spain, international financial flows also appear important in giving rise to conditions in which the economy suffers strongly the effects of an external shock.

Therefore, the OAT-VARs identify financial markets to create recurrent regime changes that justify asymmetric responses to the same extraordinary shocks and also trade and financial integration (flows) to justify slower changes in the transmission mechanism. Asymmetries of responses depending on the country that originates the shock are also found for the UK. Because of the large differences of the responses between histories as presented in Table II, our results suggest that assuming the same response across histories would introduce a bias in the estimation of the responses.

With reference to the three findings that emerged from the linear VARs, they are basically at least qualitatively confirmed in this more general setting. In particular, it is still generally true that the COS shock generates the largest effects (except for OAT-VARs with DREER and DSPI for Spain); that the difference between SIS and PIS is in general not statistically different; that most of the effects take place within one year; and that the US shocks have stronger effects than the German ones (except for OAT-VARs with ST for UK and Italy).

[6] US shocks have positive effects on the UK while German shocks have negative effects when the real exchange rate is sufficiently depreciated, net imports of oil are small and the value of net claims of US banks on the UK is relatively low.

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4. EVALUATION OF OAT-VARS

Although we have presented in-sample measures supporting OAT-VARs in Section 2.5 and have used bootstrapped confidence intervals to identify asymmetric responses, a forecasting exercise is still a relevant tool to check the adequacy of the chosen specifications by measuring their accuracy compared with simpler specifications with the same type of information. This is even more relevant when our identification of asymmetric responses presented in Section 3.5 depends on many specification choices.

Using the OAT-VARs with the transition variables that suggest asymmetric responses, we consider 1- to 5-quarter ahead forecasts for the final five years in the sample, using rolling estimation with a 95 quarters window. To compare the relative performance of linear and nonlinear specifications, we compute the mean squared forecast error (MSFE). The use of rolling estimation allows us to compare the two loss functions using the Diebold and Mariano (1995) (DM) test with an asymptotic standard normal distribution even if the competing models are nested (see Giacomini and White, 2003). Because of the nonlinearities, iterative forecasts based on the conditional means ($E[y_{T+h}|I_T]$) are computed by bootstrap for $h > 1$ in the case of OAT-VARs.

Table IV reports the ratios of the MSFE of the OAT-VARs for predicting output growth for two benchmarks: the trivariate VAR and the same VAR with the transition variable as an additional endogenous variable. This latter benchmark is employed to consider also the information of the transition variable in the linear specification. Ratios smaller than one indicate a smaller loss from the OAT-VARs and an asterisk signals that the loss differential is statistically significant at the 5% level using the DM test with the small sample correction suggested by Harvey *et al.* (1997).

The performance of the OAT-VARs deteriorates with the forecast horizon when comparing with the VAR. When $h = 5$ only 4 out of 14 nonlinear models outperform the linear VARs and the gains are small, but at the shortest horizon 10 OAT-VARs beat the linear VARs, in a few cases with sizeable gains. The inclusion of the transition variable in the linear VAR in general worsens its forecasting performance. For example, now 12 out of 14 OAT-VARs outperform the linear VAR for $h = 5$. As in previous forecasting exercises comparing the forecasting performance of linear and nonlinear models with small samples (Terasvirta, 2005), these results do not indicate strong gains from nonlinearities.

For further checking on the adequacy of OAT-VARs, we evaluate their ability in predicting increases and decreases of output growth. We compute the test proposed by Pesaran and Timmermann (1992) using the OAT-VARs forecasts for two situations: $\Delta y_{X,t+h} > 0$ and $\Delta y_{X,t+h} < 0$. The results, presented in Table IV for selected forecast horizons, show that in most cases it is possible to reject the null of no predictive content of the directional forecasts from the OAT-VARs.

In summary, the forecasting evaluation conducted in this section confirms that OAT-VARs are adequate specifications to capture the effects of the domestic conditions (transition variables) in the transmission mechanism.

5. ADDITIONAL APPLICATIONS OF OAT-VARS

In this section we provide additional applications of the OAT-VARs. In particular, we use them to evaluate the consequences for European countries of the 2001 US recession and to identify the variables that explain increasing synchronization.

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Table IV. Evaluation of out-of-sample forecasts of y_X

	Ratio RMSFE										Pesaran-Timmermann Predictive Test					
	Benchmark: VAR					Benchmark: VAR including the transition variable					$h=1$		$h=3$		$h=5$	
	$h=1$	$h=2$	$h=3$	$h=4$	$h=5$	$h=1$	$h=2$	$h=3$	$h=4$	$h=5$	$-$	$+$	$-$	$+$	$-$	$+$
France																
RST	1.32*	1.42*	1.43*	1.45*	1.56	0.96	0.79	0.74	0.71*	0.60	0.006	0.001	0.004	0.004	0.033	0.033
Italy																
DREER	0.92	1.17*	1.15	1.30	1.24	1.00	0.86*	1.00	1.00	0.89	0.000	0.339	0.000	0.057	0.001	0.073
DSPI	1.01	1.09	0.98	1.14	1.20	1.12	1.01	1.13	1.03	0.86*	0.000	0.077	0.000	0.057	0.000	0.000
BIN	0.98	1.27*	1.68*	1.33	1.49	0.92	0.81*	0.62	0.82	0.60	0.000	0.006	0.000	0.341	0.000	0.042
TRADE	1.05	1.02	1.14	1.05	0.98	0.79	0.91*	0.97	1.14	1.04	0.000	0.013	0.000	0.100	0.000	0.124
UNEM	0.90	0.88	1.01	1.18*	1.06	0.93	1.15	0.92*	0.95	0.89*	0.000	0.044	0.001	0.432	0.000	0.011
Spain																
DREER	0.86	1.02	1.07	0.87	0.93	1.08	0.92*	0.95	0.94	0.90*	0.002	0.009	0.000	0.001	0.013	0.047
DSPI	0.91	0.97	0.90	0.97	0.96	0.83*	0.89*	1.04	0.85*	0.89*	0.013	0.013	0.012	0.003	0.021	0.021
DM2SA	0.85	0.92	0.94	0.90	1.00	0.95	1.07	0.93	0.96	0.98	0.013	0.013	0.117	0.015	0.002	0.000
BIN	0.84	0.93	1.02	1.06	1.13	1.23	0.86	0.99	0.94	0.83*	0.000	0.000	0.009	0.009	0.199	0.085
TRADE	0.96	0.87	0.97	1.00	1.01	0.65*	1.08	1.08	0.76*	0.86*	0.451	0.062	0.017	0.000	0.500	0.013
UK																
ST	0.78*	1.10	1.29	1.15*	0.97	1.00	0.93	0.86*	0.78*	0.86*	0.055	0.128	0.003	0.012	0.022	0.180
Germany																
TNET	1.07	1.09	1.06	1.17*	1.02	0.84	0.94	1.02	0.87	0.97	0.123	0.123	0.003	0.116	0.100	0.223
ST	0.91	1.06	1.10	1.13	1.09	1.33*	1.12	0.99	0.86*	1.04	0.005	0.017	0.004	0.015	0.002	0.036

Note: y_X is the output growth of indicated country X. The RMSFE is computed with $n = 20$ rolling forecasts (window size 95). The ratios presented imply that OAT-VARs are more accurate when ratios are smaller than 1. * means that the null of equal forecast accuracy is rejected at 10% using the Diebold and Mariano test with variance computed with Newey–West procedure and using critical values from the t distribution. The entries for the Pesaran–Timmermann predictive test are p-values computed for decreasing ($-$) and increasing ($+$) output growth. Rejection of the PT test suggests that the model is able to predict the direction of output growth.

5.1. OAT-VARs and the 2001 Downturn

The previous section presented evidence of regime-changing behaviour in the transmission mechanism of external shocks. Here we return to the motivation outlined in the introduction. We employ our model to answer the question whether—and to what extent—the European phase of slow growth from 2001 : 1 onwards can be explained primarily as a spillover from a US shock (generally associated with the bursting of the 'dotcom bubble' there) or has to be explained as a response to a 'common shock'. Recall that our model has been developed on the basis that it incorporates changes in response dependent on the value of given transition variables relative to their thresholds. If indeed the European phase of slow growth can be thought of largely as a spillover from the USA, then it is clear that our method offers a relatively refined way of capturing it.

We characterize the US downturn as generated by a −3% shock in the annual growth rate at 2000 : 4. Shocks of this size and sign had occurred in the past, and they are associated with dated US recessions. To measure the effects of this shock we compute the responses for this specific sized shock. Notice that, notwithstanding the overall nonlinearity of the OAT-VAR, the effects of the shock are proportional to its sign and size because the shock does not affect the regime-changing probabilities. The latter depend on the history at the time of the shock, which defines which mechanism will regulate the shock transmission.

We assume that if there had been no shock the economies would have grown during 2001 at the rates predicted by the IMF in October 2000 (*World Economic Outlook*, October 2000). For example, the predicted GDP growth rate for the USA was 3.2%, while the actual was 0% (as reported in (*World Economic Outlook*, October 2002). This also justifies our choice of a −3 percentage point shock for the USA. The last column of Table V reports the corresponding forecast errors for the European countries, which are substantially larger for France, Germany and Italy than for Spain and the UK.

We consider the transition variables that generate asymmetric responses across regimes to compare with responses from linear VARs. Notice that even if these variables are selected using data until 2001 : 1, their choice is not biased in favour of producing a good forecasting performance, but only by classifying asymmetric histories.

The responses are presented in Table V. The linear VARs indicate that about two-thirds of the US shock is transmitted to European countries, with the strongest effects in the UK and Germany (−2%) and the mildest in France (−1.2%). Although there is a large variety of predicted responses based on OAT-VARs, only three of them are statistically different from zero. As a consequence, in general they suggest that responses from an extraordinary US negative shock may be zero. Interestingly, large negative and significant responses are found for Italy, Spain and Germany. These responses are generated using the estimated regime for 2000:Q4 that, in all these three cases, is the one in which US shocks have large effects. So these economies were in the most vulnerable regime when they were hit by the US shock.

In summary, the results suggest that the 2001 downturn was generated by a US shock for some specific transition variables and countries, such as Germany and Italy, while, even if the existence of switching regimes increases the expected responses of a US shock, a common shock is needed to generate stronger responses for some other countries, e.g., France. Overall, an element of 'common shock' must have been present to the extent that European holdings of dotcom shares were significant, and Castrèn *et al*. (2003) set out an analytical model suitable for analysing this case.

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Table V. Effect of a −3 percentage point annual GDP growth US shock at 2000:4

Country	Transition variable	US shock	Actual effect
France	RST	−1.0	−3.2
	Linear	−1.2*	
Italy	DREER	−4.3	
	DSPI	−2.6*	
	BIN	−1.3	−2.5
	TRADE	−1.4	
	UNEM	−3.9	
	Linear	−1.9*	
Germany	TNET	−3.5*	
	ST	−4.5	−3.3
	Linear	−2.0*	
Spain	DREER	−1.7	−1.2
	DSPI	−7.0	
	DM2SA	−2.0*	
	BIN	−0.8	
	TRADE	−2.7	
	Linear	−1.7	
UK	ST	−3.6	−1.1
	Linear	−2.0*	

Note: Values in percentage points per year. The size of the US shock is −3, which is approximately the forecast error made by the IMF for US growth in 2001 (actual growth for 2001 in WEO, October 2002, minus forecast growth for 2001 in WEO, October 2000, exact forecast error is −3.2). The responses are obtained either from OAT-VARs with the indicated transition variables, or from a linear VAR (Linear). * means that the response is statistically different from zero. Actual effect indicates for each country the difference of actual growth for 2001 as in WEO, October 2002, and the forecast growth for 2001, as in WEO, October 2000.

5.2. OAT-VARs and Increasing Synchronization

An important motivation for this paper was to improve our understanding of synchronization between the economies of the USA and European countries. An interesting way to use our estimated OAT-VARs is to observe whether they provide a basis for predicting a definite change in synchronization over time. In our modelling, an increase in synchronization over time implies that if an extraordinary US shock hits a European economy, it will have larger effects than previously.

Table VI shows the percentage of time that a given European country is in the regime in which US shocks have a larger effect. This percentage is computed for two subsamples of 11 years and for all the OAT-VARs with specifications that create asymmetric responses.[7] On average, the likelihood of the European countries being in the regime of stronger transmission of US shocks is larger after 1990, which is evidence of increasing synchronization. For France, the UK and Germany, the key variable to explain this synchronization is monetary policy asymmetry that leaves these countries more vulnerable to shocks from outside Europe (OAT-VARs with ST and

[7] The results for the OAT-VAR with BIN (financial flows) as transition variable are not presented because there is no regime change after 1980 for this specification.

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Table VI. Percentage of time that country is in the regime in which US shocks have strong effects

Country	Transition variable	Reg	1979 : Q1–1989 : Q4	1990 : Q1–2000 : Q4
France	**RST**	1st	31.8	45.45
Italy	**DREER**	1st	65.9	72.73
	DSPI	2nd	56.8	61.36
	TRADE	1st	95.5	65.9
	UNEM	1st	90.9	81.82
Spain	DREER	2nd	22.7	13.64
	DSPI	1st	25	31.82
	DM2SA	2nd	90.9	63.64
	BIN	2nd	52.3	47.73
	TRADE	2nd	18.18	95.5
UK	**ST**	1st	43.2	84.09
Germany	TNET	1st	54.5	27.3
	ST	2nd	9.09	36.6
Average			50.52	55.97

Note: Emboldened transition variables indicate that the effect of an extraordinary US shock has increased over time.

RST). For Italy and Spain, booming stock markets explain this vulnerability (OAT-VARs with DSPI). Finally, growing trade integration with other European countries explains the largest part of the increasing synchronization of Spain.[8]

6. CONCLUSIONS

In this paper, we propose a model that allows for recurrent changes in the transmission mechanism of an extraordinary shock depending on an observable transition variable. Using the generalized impulse responses from this model, we identify changes in the transmission of external shocks to European countries. The transition variables that underlie the largest modifications in the transmission mechanism include exchange rates, financial prices, international capital flows, trade links and monetary policy instruments. These modifications in the transmission mechanism are responsible for increasing synchronization over time for some countries and specific transition variables.

Moreover, there are some gains in allowing for a changing transmission mechanism when analysing the strong effect of the recent US recession on the European economies. In particular, we find that for Germany and Italy most of the slowdown in 2001 can be explained by the US shock when its effects are evaluated within a nonlinear VAR whose parameter evolution is driven by financial variables such as international financial flows or share prices. For France instead a common shock is required to explain the 2001 recession, while the slowdown was rather limited in Spain and the UK. This diversity of effects of the US shock across countries suggests that it is

[8] It may be worth pointing out that this identification of synchronization is independent of the occurrence of recessions in the USA. This is supported by the fact that the economies spend more time in the regime that suggests stronger synchronization ($\simeq$55%) than the USA in recession ($\simeq$15%), where the latter periods are identified by reference to the NBER dating of recessions in the USA.

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dangerous—or at least premature—to consider Europe or the euro area as a whole when discussing synchronization and interrelationships between countries' business cycles. If the introduction of the euro will lead, eventually, to a greater conformity of reaction to external shocks, this had not taken place by 2001. The disappearance of the effects of monetary policy asymmetry with the adoption of the single currency did not also reduce the effects of differently timed stock market surges, for example.

Future research could exploit the modelling approach proposed in this paper to examine either changes in the monetary transmission mechanism or the impact of fiscal shocks on the economy.

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