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Is the international propagation of financial shocks non-linear? Evidence from the ERM

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Abstract

This paper tests for the presence of non-linearities in the propagation of devaluation expectations among the countries that were members of the Exchange Rate Mechanism of the EMS. We show that whenever it is possible to estimate a model for financial interdependence, a full-information technique to detect such non-linearities is more efficient than the limited-information estimator proposed, in a similar context, by Rigobon (2000). This happens, in particular, when the periods of market turbulence are relatively short. Our evidence suggests that non-linearities in the propagation of devaluation expectations were a general phenomenon in the ERM. Normally the non-linearity amounts to a stronger effect in the same direction, but sometimes, as in the Dutch case, it implies a significant effect in the opposite direction: evidence of flight-to-quality.

Keywords: Contagion; EMS

JEL classification: E3; F3; P4

1. Introduction

The experience of recent financial crises suggests (see for instance Baig and

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Goldfajn, 1998) that the interdependence connecting financial markets in normal times tends to break down after unusually large shocks — in other words that the international propagation of financial shocks may be non-linear. Often, such non-linearity results in a sharp increase in the response of one market to a shock originated elsewhere: this, for instance, was the experience of the recent crises that originated in Mexico, Russia and South-East Asia in the 1990s. As we show in the present paper, however, the non-linearity may also produce a response whose sign is the opposite of what we observe in normal times, thus suggesting a flight-to-quality phenomenon.

The paper investigates this issue in the context of the European exchange rate experience. We study devaluation expectations among the currencies that belonged to the Exchange Rate Mechanism (ERM) of the European Monetary System. Our aim is to test for the presence of non-linearities in the way devaluation expectations spread across Europe during ERM crises.

To measure devaluation expectations we analyse the spreads between 3-month German rates and 3-month interest rates in other European countries. Our choice of a 3-month horizon is justified by two observations: we need an horizon long enough, so that interest rate spreads reflect exchange rate expectations, rather than money market intervention by the central banks; at the same time, the horizon should not be too long, otherwise spreads would average exchange rate expectations over long periods of time, and would thus fail to capture the expectation of an exchange rate crisis precisely enough.

We use a sample running from January 1988 to August 1992. This was a period characterized by the absence of realignments (the last ERM realignment before the September 1992 devaluations occurred during 1987): we can thus assume that the monetary policy regime was constant throughout the period — namely determined by the exchange rate constraint. This assumption allows us to exclude that shifts in interdependence could be generated by a shift in the monetary regime in one of the countries considered — as a result, for instance, of the transition from fixed to flexible exchange rates.

We use weekly data (the spreads on German rates observed on the Wednesday of each week) for six ERM members (France, Italy, Spain, Belgium, Holland and Denmark) plus Sweden. We add Sweden because, though not formally an ERM member, the Swedish Krone shadowed the Deutschemark throughout our sample, until it was eventually devalued just after the break-up of the ERM. We instead exclude the UK. The pound joined the ERM in the middle of our sample (in the Fall of 1990): this change in monetary regime leaves too few observations to allow us to estimate a model of interdependence that includes the UK.

A number of recent papers (e.g. Baig and Goldfajn, 1998; Forbes and Rigobon, 1999; Rigobon, 2000) refer to non-linearities in the international propagation of financial shocks using the term ‘contagion’. While engaging, this definition could be deceptive, since it suggests that the non-linearity always shows up as an increase in the response of one market to a shock originated elsewhere — an assumption which, as we show, is not generally true, and overlooks the possibility of 'flight-to-quality' during episodes of market turbulence. Consider, for example, a situation in which markets classify countries in two classes, strong and weak: the arrival of some bad news could increase the interdependence within the two classes, while reducing the interdependence between them. Contagion is not the appropriate term to describe a situation like this, which, as we shall see, is not uncommon in our sample.

We proceed in three steps. First, we identify the channels through which shocks are normally propagated across markets: this can be done estimating a model of financial interdependence. The recent literature on interdependence and contagion (Rigobon, 1999) has stressed the importance of modelling interdependence to avoid a spurious detection of contagion: the same argument applies to the detection of non-linear interdependence. Next, we identify 'crises', namely episodes of market turbulence during which non-linearities may arise. Finally we test the hypothesis that during crises the normal channels of interdependence are modified.

Implementing step one is often difficult, particularly when interdependence extends over many countries, and thus requires the estimation of large models. In a paper whose purpose is very similar to ours, Rigobon (2000) solves this problem using a limited information technique based on an instrument which is constructed splitting the sample into high- and low-volatility observations. Here we show that the full-information estimation of a model for interdependence avoids the problems that arise when the size of the sub-sample of high-volatility observations is small, thus delivering a more powerful test.

2. Estimating financial interdependence

In this section we describe the three-step procedure outlined above and we apply it to the propagation of devaluation expectations among ERM members. We start from the estimation of a statistical model (a reduced-form VAR) that describes the joint process generating the interest rate spreads. We specify this model allowing for the constraints imposed on each country by membership in the ERM.

Consider, for simplicity, an ERM consisting of only three countries: country 1, which represents Germany, the core of the ERM, and countries 2 and 3, two other members of the system. Let R_1, R_2 and R_3 be their short term interest rates, and $s_{21} = R_2 - R_1$ and $s_{31} = R_3 - R_1$, the spreads, which reflect expectations of exchange rate depreciation. The conditional distribution of s_{21} and s_{31} is described by the following reduced form:

$$\begin{pmatrix} s_{21,t} \\ s_{31,t} \end{pmatrix} = \begin{pmatrix} \pi_{11} & \pi_{12} \\ \pi_{21} & \pi_{22} \end{pmatrix} \begin{pmatrix} s_{21,t-1} \\ s_{31,t-1} \end{pmatrix} + \begin{pmatrix} u_{1,t} \\ u_{2,t} \end{pmatrix} \tag{1}$$

$$\begin{pmatrix} u_{1t} \\ u_{2t} \end{pmatrix} \Big| I_{t-1} \sim \left[\begin{pmatrix} 0 \\ 0 \end{pmatrix}, \begin{pmatrix} \sigma^2_{1t} & \sigma_{12,t} \\ \sigma_{12,t} & \sigma^2_2 \end{pmatrix} \right]$$

Note that the residuals in (1) are heteroskedastic and non-normal: this allows for the possibility that the sample includes periods of financial turbulence, to be detected by standard tests. We thus re-specify (1) as

$$\begin{pmatrix} s_{21,t} \\ s_{31,t} \end{pmatrix} = \begin{pmatrix} \pi_{11} & \pi_{12} \\ \pi_{21} & \pi_{22} \end{pmatrix} \begin{pmatrix} s_{21,t-1} \\ s_{31,t-1} \end{pmatrix} + B^{-1} \begin{pmatrix} \epsilon_{1,t} \\ \epsilon_{2,t} \end{pmatrix} \tag{2}$$

$$\begin{pmatrix} \epsilon_{1,t} \\ \epsilon_{2,t} \end{pmatrix} = \left(I + \begin{pmatrix} a_{11} & a_{12} \\ a_{21} & a_{22} \end{pmatrix} \begin{pmatrix} d_{1,t} & 0 \\ 0 & d_{2,t} \end{pmatrix} \right) \begin{pmatrix} \epsilon^l_{1,t} \\ \epsilon^l_{2,t} \end{pmatrix}$$

$$\begin{pmatrix} \epsilon^l_{1t} \\ \epsilon^l_{2t} \end{pmatrix} \Big| I_{t-1} \sim N \left[\begin{pmatrix} 0 \\ 0 \end{pmatrix}, \begin{pmatrix} \sigma^2_1 & \sigma_{12} \\ \sigma_{12} & \sigma^2_2 \end{pmatrix} \right]$$

where $\epsilon_{1,t}$, $\epsilon_{2,t}$ are the structural shocks, B is the matrix defining the contemporaneous feedbacks between $s_{21,t}$ and $s_{31,t}$, and thus their interdependence. The vector of dummies

$$d = \begin{pmatrix} d_{1,t} \\ d_{2,t} \end{pmatrix}$$

filters heteroskedasticity and non-normality out of the residuals by identifying episodes of market turbulence. The vector is partitioned in two blocks according to whether the event generating the turmoil occurred in country 1 or in country 2. $\epsilon^l_{1,t}$, $\epsilon^l_{2,t}$ are the structural shocks in low-volatility periods: they are normally distributed and homoskedastic. The coefficients in the A matrix allow for non-linearities: they describe how the propagation of financial shocks across countries is modified during periods of turmoil. The diagonal blocks of A define the extent to which the normal-time structural shocks get amplified within countries; the diagonal blocks allow for non-linearities in the propagation of such shocks across countries. A natural test for the absence of non-linearities is then a test of the following null hypothesis:

$$H_0: a_{ij} = 0, \text{ for each } i \neq j$$

Note that such a test could not be implemented in the reduced form (2) since it requires the identification of the parameters in the B matrix, i.e. of the channels of interdependence. In other words, the finding that shocks in country i propagate to country j in the reduced form does not necessarily imply the presence of a nonlinearity.[1] Consistently with this observation, we shall consider an identification scheme imposes no a-priori restrictions on the simultaneous feedback among the variables.

To give a simple example of how the test can be implemented, consider the case in which our structural model is just-identified by the restriction that, in each equation, the own lagged dependent variable is sufficient to capture the structural dynamics:

$$\begin{pmatrix} 1 & -\beta_{12} \\ -\beta_{21} & 1 \end{pmatrix}\begin{pmatrix} s_{21,t} \\ s_{31,t} \end{pmatrix} = \begin{pmatrix} \gamma_{11} & 0 \\ 0 & \gamma_{22} \end{pmatrix}\begin{pmatrix} s_{21,t-1} \\ s_{31,t-1} \end{pmatrix} + \left(I + \begin{pmatrix} a_{11} & a_{12} \\ a_{21} & a_{22} \end{pmatrix}\begin{pmatrix} d_{1,t} & 0 \\ 0 & d_{2,t} \end{pmatrix}\right)\begin{pmatrix} \epsilon^{l}_{1,t} \\ \epsilon^{l}_{2,t} \end{pmatrix} \tag{3}$$

$$\begin{pmatrix} \epsilon^{l}_{1,t} \\ \epsilon^{l}_{2,t} \end{pmatrix} \Bigg| I_{t-1} \sim N\left[\begin{pmatrix} 0 \\ 0 \end{pmatrix}, \begin{pmatrix} \sigma^2_{\epsilon 1} & 0 \\ 0 & \sigma^2_{\epsilon 2} \end{pmatrix} \right]$$

In this example the parameters β_{12} and β_{21} determine interdependence, while the vector of parameters a_{12} and a_{21} allow for non-linearities. The identification of (3) is achieved by setting $\gamma_{12} = \gamma_{21} = 0$, i.e. by restricting interdependence to occur only simultaneously: in other words we impose that markets react instantaneously.[2] This, as mentioned above, is the identifying assumption we shall use in our empirical work: it imposes no restrictions on the simultaneous feedback among financial variables; it instead restricts the shape of the response of each spread to structural shocks.[3]

It is instructive to compare our full-information approach to the limited information approach originally proposed by Rigobon (2000) who estimates interdependence using instrumental variables. The technique hinges upon splitting the sample into high and low volatility periods: an instrument is then constructed whose validity is guaranteed under the null of linearity. The test for non-linearity is then simply a test of the validity of the instruments.

To illustrate this procedure within our example, consider the simple case in which $a_{11} = a_{21} = 0$. In this case the dependent variable is split into high (s^h) and low (s^l) volatility observations in the following way:

[1] This point has been made by Rigobon (1999), who shows that simple correlations are the wrong indicator to detect 'contagion'.

[2] Note that this assumption does not exclude the possibility that spreads are serially correlated: the existence of interdependence in the structural model generates serial correlation in the joint conditional distribution (i.e. the reduced form) of interest rate spreads.

[3] Identification based on restrictions on the shape of impulse response functions has recently attracted interest in the structural VAR literature; see, for example, Uhlig (1999).

$$\begin{pmatrix} 1 & -\beta_{12} \\ -\beta_{21} & 1 \end{pmatrix}\begin{pmatrix} s^h_{21,t} \\ s^h_{31,t} \end{pmatrix} = \begin{pmatrix} \gamma_{11} & 0 \\ 0 & \gamma_{22} \end{pmatrix}\begin{pmatrix} s^h_{21,t-1} \\ s^h_{31,t-1} \end{pmatrix} + \left(I + \begin{pmatrix} 0 & a_{12} \\ 0 & a_{22} \end{pmatrix}\begin{pmatrix} d_{1,t} & 0 \\ 0 & d_{2,t} \end{pmatrix}\right)\begin{pmatrix} \epsilon^l_{1,t} \\ \epsilon^l_{2,t} \end{pmatrix} \tag{4}$$

$$\begin{pmatrix} 1 & -\beta_{12} \\ -\beta_{21} & 1 \end{pmatrix}\begin{pmatrix} s^l_{21,t} \\ s^l_{31,t} \end{pmatrix} = \begin{pmatrix} \gamma_{11} & 0 \\ 0 & \gamma_{22} \end{pmatrix}\begin{pmatrix} s^l_{21,t-1} \\ s^l_{31,t-1} \end{pmatrix} + \begin{pmatrix} \epsilon^l_{1,t} \\ \epsilon^l_{2,t} \end{pmatrix} \tag{5}$$

Consider now the following instrument[4] for s_{31}:

$$w_t = \begin{pmatrix} \dfrac{s^h_{31,t}}{T^h} \\ -\dfrac{s^l_{31,t}}{T^l} \end{pmatrix},$$

Using w_t as an instrument for s_{31} leads to the following just-identified instrumental variables estimator of the interdependence parameter β_{12}:

$$\beta_{12} = (w' s_{31})^{-1} w' s_{21}$$

The two usual conditions for validity and consistency of the IV estimator are checked by looking at the probability limits of $(w' s_{31})$ and $(w' \epsilon_1^*)$, where $\epsilon_1^* = a_{12} d_{2,t} \epsilon_2^l + \epsilon_1^l$

$$\begin{aligned} p\lim(w' s_{31}) &= p\lim \frac{1}{T^h} s^h_{31,t} s^h_{31,t} - p\lim \frac{1}{T^l} s^{l\prime}_{31,t} s^l_{31,t} \\ &= \left(-\frac{\beta_{21} a_{12}}{-1+\beta_{12}\beta_{21}} d_{2t} - \frac{a_{22} d_{2t}}{-1+\beta_{12}\beta_{21}} \right)^2 \sigma^2_{\epsilon 2} \end{aligned}$$

$$p\lim(w' \epsilon_1^*) = \left(\frac{\beta_{21} a_{12}}{1-\beta_{21}\beta_{12}} d^2_{2t} - \frac{a_{12} a_{22}}{1-\beta_{21}\beta_{12}} d^2_{2t} \right) \sigma^2_{\epsilon 2}$$

The validity of w_t as an instrument is guaranteed under $H_0 (a_{12} = 0)$, while its efficiency depends on the degree of heteroskedasticity between low and high volatility observations.

Within this framework, the presence of non-linearities can be tested applying a Hausman (1978) test for the validity of the instruments. The beauty of this approach depends on the fact that, in the presence of heteroskedasticity it does not require variables other than $s_{31,t}$ to implement the IV estimator. Avoiding the estimation of a structural model of interdependence has the obvious benefit of

[4]The statistical evidence discussed in this section is available from the authors upon request.

imposing milder identifying restrictions than those necessary to implement our full-information procedure. In particular, the limited information approach has the advantage of identifying the system even when the just identifying restrictions that we propose are not valid ($\gamma_{12} \neq 0$, $\gamma_{12} \neq 0$). This, of course, comes at the cost of less power: the loss of efficiency could be non-negligible in cases where the number of observations for one of the two alternative regimes is low. Think of the limiting case in which the high-volatility sub-sample consists of just one observation: asymptotic results in the entire sample, of dimension T, are still applicable, but obviously not in the sub-sample of high volatility observations. In such a situation our methodology, based on a full information estimation on the whole sample, with the inclusion of dummies for high-volatility periods, is instead still feasible. Obviously when T^h and T^l are sufficiently long and our just identifying restrictions are valid, the limited and full information approaches both produce consistent estimators and, therefore, the same results. (An example of this case is discussed in a companion paper (Bonfiglioli and Favero, 2000) where limited and full information approaches are applied to the analysis of interdependence and contagion between the US and the European stock markets.)

3. Looking for non-linearities in the propagation of devaluation expectations in the ERM

In this section we implement the technique outlined above to study the propagation of devaluation expectations across European money markets during the ERM. Are the normal channels of financial interdependence enough to explain the way such expectations were transmitted from one market to another, or is there evidence of non-linearities? As explained in Section 2, we proceed in three steps: we first identify the episodes of market turbulence analysing the residuals obtained from a reduced form VAR; next we estimate a structural model which describes interdependence among interest rate spreads; finally we test for non-linearities.

3.1. Detecting market turmoil in the ERM

We start our empirical investigation by specifying a reduced form VAR for the joint distribution of European interest rate spreads. The source for the data (weekly observations on 3-month Euro rates) is Datastream. We consider the spreads on German rates for seven countries: France, Italy, Spain, Holland, Belgium, Denmark and Sweden. (The reason for including Sweden, which at the time was not an ERM member, while excluding the UK, were explained in the Introduction.)

The estimation of a first-order VAR for interest rates spreads, as described in (1), produces a number of large residuals — defined as residuals with an absolute value three times larger than the estimated standard deviation. We have thus included in the VAR twenty-six dummies to eliminate a corresponding number of outliers, as described in (2). Estimates of the reduced-from coefficients are shown in Table 1, where we report the coefficients on the twenty-six dummies separately. The residuals obtained from a VAR that includes the dummies show no apparent evidence of correlation, nor of heteroskedasticity (this is confirmed by the tests reported in Table 1), although there remains some (moderate) non-normality.

With the exception of Holland, all spreads show a very high degree of persistence. Moreover, with no exception, the coefficient on the lagged dependent variable is the only significant coefficient in the lag structure. When we study the equilibrium properties of the data applying the Johansen (1995) procedure, we find evidence in favour for the stationarity of the spread only for Holland[5] — a result which suggests, except for Holland, a low credibility of the exchange rate commitment of the ERM members considered in our sample.

Each one of the observations identified by the twenty-six dummies can be traced to a piece of news relevant for financial markets in the ERM: we describe such news in Table 1 below the corresponding observation/s. Based on this information, shocks are defined 'local' or 'common' depending on whether they hit a single ERM-member, or more countries at the same time. The dummies corresponding to common shocks are by definition significant in more than one country. Looking at Table 1 we see, however, that often, even when a shock is identified as local, the coefficient on the corresponding dummy is significant not only in the country where the shock originates, but in other countries as well. For example, the dummy that identifies the Dutch shock of May 3, 1989 is significant not only in Holland, but also in Spain. The finding that the dummy corresponding to a local shock is significant in more than one country does not necessarily imply the presence of a non-linearity, since it could simply be the effect of normal interdependence.

To identify non-linearities we thus estimate a structural model to control for interdependence.

3.2. Modelling interdependence and testing for non-linearities

To model interdependence we estimate a structural simultaneous model (see Hendry (1995)) for the determination of interest rate spreads.

As already discussed with reference to (1), we achieve identification by

[5]According to both the trace and the maximum eigenvalue tests, the rank of the long-run matrix is one, and the restrictions that only the Dutch spread belongs to the equilibrium relationship is not rejected. However, some care in interpreting these results must be exercised in the light of the presence of dummies, and of the results reported in Johansen (1999).

Table 1
A reduced form model of European interest rate spreads

Sample: November 2, 1988–September 9, 1992.
Weekly data observed on the Wednesday of each week
Estimation by OLS. Standard errors in brackets.

dep. var.	constant	s_{t-1}^{NL}	s_{t-1}^{FR}	s_{t-1}^{IT}	s_{t-1}^{ES}	s_{t-1}^{DK}	s_{t-1}^{SW}	s_{t-1}^{BG}
s_t^{NL}	0.008	0.63	−0.02	0.001	0.006	−0.04	−0.02	0.10
	(0.043)	(0.06)	(0.03)	(0.009)	(0.01)	(0.02)	(0.008)	(0.04)
s_t^{FR}	−0.02	−0.17	0.90	0.013	0.015	0.028	−0.02	0.06
	(0.06)	(0.09)	(0.04)	(0.013)	(0.015)	(0.023)	(0.012)	(0.05)
s_t^{IT}	0.03	−0.21	−0.06	1.03	−0.04	0.07	0.01	−0.06
	(0.11)	(0.16)	(0.07)	(0.03)	(0.03)	(0.04)	(0.02)	(0.10)
s_t^{ES}	0.09	−0.2	0.2	0.1	0.96	−0.0004	−0.0004	0.09
	(0.07)	(0.10)	(0.05)	(0.2)	(0.02)	(0.03)	(0.01)	(0.06)
s_t^{DK}	0.03	−0.10	−0.018	−0.002	0.02	0.97	−0.03	0.04
	(0.08)	(0.12)	(0.05)	(0.02)	(0.02)	(0.03)	(0.02)	(0.07)
s_t^{SW}	−0.07	0.06	−0.03	−0.06	0.004	0.10	0.95	−0.17
	(0.12)	(0.2)	(0.08)	(0.03)	(0.03)	(0.05)	(0.03)	(0.10)

Testing for vector autocorrelation of the residuals (lags 1 to 7): $F(343,984)=0.97$ [0.61]
Testing for vector heteroskedasticity of the residuals: $F(952,3041)=0.86$ [0.99]

Standard deviations and correlation matrix of residuals

	σ	s_t^{NL}	s_t^{FR}	s_t^{IT}	s_t^{ES}	s_t^{DK}	s_t^{SW}	s_t^{BG}
s_t^{NL}	0.10	1.00	0.36	0.30	0.29	0.44	0.30	0.48
s_t^{FR}	0.15	0.36	1.00	0.26	0.39	0.33	0.20	0.49
s_t^{IT}	0.27	0.30	0.26	1.00	0.23	0.26	0.18	0.24
s_t^{ES}	0.17	0.29	0.39	0.23	1.00	0.36	0.28	0.44
s_t^{DK}	0.19	0.44	0.33	0.26	0.36	1.00	0.28	0.34
s_t^{SW}	0.31	0.30	0.20	0.18	0.28	0.28	1.00	0.30
s_t^{BG}	0.13	0.48	0.49	0.24	0.44	0.34	0.30	1.00

Dummies in the reduced form

Dummies	dep. var.						
	s_t^{NL}	s_t^{FR}	s_t^{IT}	s_t^{ES}	s_t^{DK}	s_t^{SW}	s_t^{BG}
21/12/88	0.32**	0.67**	−0.33	0.81**	0.59**	0.51**	0.54**
	(0.10)	(0.15)	(0.28)	(0.18)	(0.20)	(0.25)	(0.14)
common shock	Bundesbank raises policy rates						
08/03/89	0.07	0.024	0.99**	0.53**	0.29	0.50**	0.11
	(0.10)	(0.15)	(0.28)	(0.18)	(0.20)	(0.25)	(0.14)
local shock	Bank of Italy raises rates after bad trade deficit data						
03/05/89	0.29**	−0.0212	−0.09	0.39**	−0.03	0.02	−0.03
	(0.10)	(0.15)	(0.28)	(0.18)	(0.20)	(0.25)	(0.14)
local shock	Dutch government resigns						
17/05/89	−0.10	−0.20	−0.14	−0.72**	−0.33	−0.56**	−0.22
	(0.10)	(0.15)	(0.28)	(0.18)	(0.20)	(0.25)	(0.14)
local shock	Spain announces sharp cuts in public spending						
05/07/89	0.023	0.34**	−0.10	0.64**	0.34	−0.05	−0.16
	(0.10)	(0.15)	(0.28)	(0.18)	(0.20)	(0.25)	(0.14)
common shock	US dollar collapses						
11/10/89	0.18	0.093	−0.02	0.10	0.47**	0.03	0.61**
	(0.10)	(0.15)	(0.28)	(0.18)	(0.20)	(0.25)	(0.14)

Table 1. Continued

Sample: November 2, 1988–September 9, 1992.
Weekly data observed on the Wednesday of each week
Estimation by OLS. Standard errors in brackets.

dep. var.	constant	s^{NL}_{t-1}	s^{FR}_{t-1}	s^{IT}_{t-1}	s^{ES}_{t-1}	s^{DK}_{t-1}	s^{SW}_{t-1}	s^{BG}_{t-1}
common shock	Bundesbank raises interest rates							
18/10/89	−0.17	0.035	0.07	−0.23	2.60**	−0.42**	−0.04	
	(0.10)	(0.15)	(0.28)	(0.18)	(0.20)	(0.25)	(0.14)	
local shock	Danish Krona hits the bottom of the ERM band							
25/10/89	−0.14	0.012	0.96**	0.12	−1.62**	−0.24	0.03	
	(0.10)	(0.15)	(0.28)	(0.18)	(0.20)	(0.25)	(0.14)	
local shock	Bundesbank intervenes to prop up the Krona							
01/11/89	−0.12	−0.087	−1.48**	−0.26	0.08	−0.37	−0.28	
	(0.10)	(0.15)	(0.28)	(0.18)	(0.20)	(0.25)	(0.14)	
common shock	Bundesbank injects liquidity in the system							
21/03/90	−0.065	−0.018	0.08	0.14	−0.03	2.19**	0.09	
	(0.10)	(0.15)	(0.28)	(0.18)	(0.20)	(0.25)	(0.14)	
local shock	Sveriges Riksbank raises rates to stem capital outflows							
14/11/90	−0.16	−0.092	0.71**	−0.04	−0.27	1.86**	−0.19	
	(0.10)	(0.15)	(0.28)	(0.18)	(0.20)	(0.25)	(0.14)	
21/11/90	0.33**	−0.08	0.81**	0.15	0.14	1.21**	0.15	
	(0.10)	(0.15)	(0.28)	(0.18)	(0.20)	(0.25)	(0.14)	
05/12/90	−0.19**	−0.07	−0.44	−0.15	−0.40**	−0.94**	0.33**	
	(0.10)	(0.15)	(0.28)	(0.18)	(0.20)	(0.25)	(0.14)	
12/12/90	0.46**	0.02	0.27	0.15	0.44**	0.69**	0.23	
	(0.10)	(0.15)	(0.28)	(0.18)	(0.20)	(0.25)	(0.14)	
local shock	Swedish recession, exchange rate pressure and interest rates hike							
27/03/91	0.10	−0.04	−0.19	−0.58**	0.09	0.45**	0.01	
	(0.10)	(0.15)	(0.28)	(0.18)	(0.20)	(0.25)	(0.14)	
local shock	Bank of Spain intervenes buying French Francs							
11/12/91	0.065	0.07	−0.08	0.08	0.023	1.5**	0.13	
	(0.10)	(0.15)	(0.28)	(0.18)	(0.20)	(0.25)	(0.14)	
18/12/91	0.10	−0.08	0.01	0.03	0.12	0.66**	0.03	
	(0.10)	(0.15)	(0.28)	(0.18)	(0.20)	(0.25)	(0.14)	
25/12/91	0.14	0.30**	−0.11	−0.06	0.026	0.94**	0.10	
	(0.10)	(0.15)	(0.28)	(0.18)	(0.20)	(0.25)	(0.14)	
02/01/92	0.056	0.24	0.21	0.11	0.25	1.02**	0.01	
	(0.10)	(0.15)	(0.28)	(0.18)	(0.20)	(0.25)	(0.14)	
08/01/92	0.19**	−0.12	−0.36	0.13	−0.03	−0.81**	−0.12	
	(0.10)	(0.15)	(0.28)	(0.18)	(0.20)	(0.25)	(0.14)	
local shock	Swedish exchange rate crisis							
08/07/92	−0.17**	0.03	1.18**	0.05	0.06	−0.10	−0.10	
	(0.10)	(0.15)	(0.28)	(0.18)	(0.20)	(0.25)	(0.14)	
22/07/92	0.018	0.04	2.27**	0.25	0.25	0.28	0.07	
	(0.10)	(0.15)	(0.28)	(0.18)	(0.20)	(0.25)	(0.14)	
29/07/92	−0.028	−0.03	−1.58**	0.25	0.02	0.28	−0.07	
	(0.10)	(0.15)	(0.28)	(0.18)	(0.20)	(0.25)	(0.14)	
05/08/92	0.104	−0.09	−1.85**	−0.06	−0.04	0.28	0.01	
	(0.10)	(0.15)	(0.28)	(0.18)	(0.20)	(0.25)	(0.14)	
local shock	Italian political and fiscal crisis							
26/08/92	0.056	0.31**	0.49	0.23	0.07	1.62**	−0.05	
	(0.10)	(0.15)	(0.28)	(0.18)	(0.20)	(0.25)	(0.14)	
09/09/92	0.076	0.16	2.5**	0.05	0.031	9.35**	0.04	
	(0.10)	(0.15)	(0.28)	(0.18)	(0.20)	(0.25)	(0.14)	
common shock	ERM crisis							

restricting the lag structure of the model.[6] Our identifying assumption is the one that allows for the highest degree of interdependence. In fact it imposes no restrictions on the contemporaneous feedbacks. This is crucial if one wants to minimize the chances of spuriously detecting non-linearities (as discussed, for instance in Rigobon, 1999, 2000).

We then move from a just-identified structure to an over-identified model by restricting to zero all contemporaneous effects and all dummies that are not significantly different from zero. We estimate the structural model by Full Information Maximum Likelihood. Within this framework the validity of over-identifying restrictions can be tested by implementing the likelihood ratio test discussed in Hausman (1983). FIML estimates of our structural model of interdependence and tests for the validity of the over-identifying restrictions are reported in Table 2 (as for the reduced form we report separately, the coefficients on the dummies). The likelihood ratio test reported in Table 2 takes a value of 165, which under the null of the validity of the 169 over-identifying restrictions has a tail probability of 0.56.

The structural model displays very little interdependence. The only significant simultaneous links arise between Belgium and Holland, and between Denmark and France, Sweden, Belgium.

The relevant evidence to test for non-linearities is in the analysis of the significance of the dummies. Under the null of linearity the dummies associated with local shocks should be significant only in the country where the shock originates. The null is rejected in twelve out of twenty episodes of local shocks. For instance, interdependence is not enough to explain the transmission to Denmark, Italy, Holland and Belgium of the effects of the Swedish banking crisis of the early 1990s. The same is true for the transmission to Spain of the local Italian and Dutch shocks occurring, respectively, in March and May 1989.

Overall all countries in our sample display some evidence of non-linearity in the transmission of devaluation expectations. Such non-linearities imply a change in the transmission across countries of devaluation expectations, which normally amounts to a stronger effect in the same direction, but sometime implies a significant effect in the opposite direction. Consider for example the case of the Dutch spread: when adverse shocks hit weaker countries, such as Italy on July 8, 1992, or Sweden on January 8, 1992, the Dutch spread closes significantly. For instance, on July 8, 1992, the Italian spread widened by 135 basis points, while the Dutch spread narrowed by 17 basis points. We interpret this as evidence of ‘flight-to-quality’.

In the light of the significance of each individual coefficient, a joint test of the

[6]We allow for the existence of equilibrium relationships, but we do not impose any specific restriction on their parameters. In doing this we run the risk of a loss of efficiency in the estimation, but we rule out inconsistency due a possibly incorrect specification of the long-run structure of our statistical model (see Sims et al., 1990).

Table 2
A structural model of European interest rate spreads

Sample: November 2, 1988–September 9, 1992. Weekly data.
Estimation by OLS. See Table 1.

dep. var.	constant	lag. dep. var	s_t^{NL}	s_t^{FR}	s_t^{IT}	s_t^{ES}	s_t^{DK}	s_t^{SW}	s_t^{BG}
s_t^{NL}	−0.02	0.71							0.047
	(0.01)	(0.04)							(0.01)
s_t^{FR}	0.0005	0.93					0.038		
	(0.02)	(0.02)					(0.01)		
s_t^{IT}	0.056	0.98							
	(0.05)	(0.01)							
s_t^{ES}	0.03	0.98							
	(0.03)	(0.01)							
s_t^{DK}	0.03	0.97							
	(0.01)	(0.01)							
s_t^{SW}	0.08	0.93					0.06		
	(0.045)	(0.01)					(0.02)		
s_t^{BG}	−0.02	0.93					0.035		
	(0.01)	(0.02)					(0.01)		

LR test of over-identifying restrictions: $\chi^2(169) = 165.228$ [0.5676]

Standard deviations and correlation matrix of residuals:

	σ	s_t^{NL}	s_t^{FR}	s_t^{IT}	s_t^{ES}	s_t^{DK}	s_t^{SW}	s_t^{BG}
s_t^{NL}	0.10	1.00	0.37	0.24	0.29	0.47	0.23	0.46
s_t^{FR}	0.14	0.37	1.00	0.28	0.39	0.33	0.13	0.49
s_t^{IT}	0.27	0.24	0.28	1.00	0.21	0.26	0.08	0.24
s_t^{ES}	0.17	0.29	0.39	0.21	1.00	0.36	0.33	0.43
s_t^{DK}	0.20	0.47	0.33	0.26	0.36	1.00	0.29	0.33
s_t^{SW}	0.23	0.23	0.13	0.08	0.33	0.29	1.00	0.35
s_t^{BG}	0.13	0.46	0.49	0.24	0.43	0.33	0.35	1.00

Dummies in structural model of Table 2

Dummies	dep. var.						
	s_t^{NL}	s_t^{FR}	s_t^{IT}	s_t^{ES}	s_t^{DK}	s_t^{SW}	s_t^{BG}
21/12/88	0.31	0.74		0.88	0.65	0.45	0.57
	(0.10)	(0.14)		(0.17)	(0.18)	(0.24)	(0.12)
common							
08/03/89			0.73	0.34			
			(0.25)	(0.15)			
local, Italy							
03/05/89	0.30			0.37			
	(0.09)			(0.15)			
local, Holland							
17/05/89				−0.55			
				(0.15)			
local, Spain							
05/07/89		0.38		0.63			
		(0.12)		(0.15)			
common							
11/10/89					0.33		0.51
					(0.16)		(0.10)
common							
18/10/89					2.70	−0.37	
					(0.16)	(0.22)	
local, Denmark							
25/10/89			1.09		−1.62		
			(0.26)		(0.17)		

Table 2. Continued

Sample: November 2, 1988–September 9, 1992. Weekly data.
Estimation by OLS. See Table 1.

dep. var.	constant	lag. dep. var	s_t^{NL}	s_t^{FR}	s_t^{IT}	s_t^{ES}	s_t^{DK}	s_t^{SW}	s_t^{BG}
01/11/89			−1.13						
			(0.25)						
common									
21/03/90						2.19			
						(0.2)			
local, Sweden									
14/11/90			0.86			2.07			
			(0.25)			(0.2)			
local, Sweden									
21/11/90	0.2		0.82			1.00	0.44		
	(0.09)		(0.25)			(0.2)	(0.11)		
local, Sweden									
05/12/90	−0.25				−1.35	−0.8	0.39		
	(0.09)				(0.18)	(0.2)	(0.10)		
local, Sweden									
	12/12/90							0.31	0.45
	(0.08)					(0.28)			
local, Sweden									
27/03/91				−0.63		0.43			
				(0.15)		(0.21)			
local, Spain									
11/12/91						0.96			
						(0.21)			
local, Sweden									
18/12/91						0.55			
						(0.21)			
local, Sweden									
25/12/91		0.24				0.87			
		(0.12)				(0.21)			
local, Sweden									
02/01/92						0.91			
						(0.21)			
local, Sweden									
08/01/92	0.18					−0.88			
	(0.08)					(0.21)			
local, Sweden									
08/07/92	−0.17		1.35						
	(0.08)		(0.25)						
local, Italy									
22/07/92			2.35						
			(0.25)						
local, Italy									
29/07/92			−1.37						
			(0.26)						
local, Italy									
05/08/92			−1.68						
			(0.25)						
local, Italy									
26/08/92		0.25				1.6			
		(0.12)				(0.21)			
common									
09/09/92			2.7			9.04			
			(0.25)			(0.21)			
common									

null hypothesis of linearity, obtained by restricting to zero the effect of dummies for local shocks in all countries other those where the shock originates clearly rejects the null. The test, distributed as $\chi^2(15)$, takes a value of 129.53 (0.000).

3.3. Robustness

We check for robustness of our results along three dimensions. We allow for a more general lag structure; we test that our results on non-linearities hold even excluding the period of extreme turbulence which characterizes the Summer of 1992; we add the effective dollar exchange rate as an exogenous variable to our specification to test that our results are robust to the inclusion of one additional (indirect) channel of interdependence.

To check for robustness to the modification of the lag structure, we augmented our specification to include lags two, three and four of all spreads. Both likelihood ratio tests and traditional lag selection criteria point to our selected specification as the preferred one: our results are robust to the consideration of a model with higher dynamics.

We have also re-estimated our model excluding the observations following May 31, 1992: this implies the exclusion of the last six dummies reported in Tables 1 and 2. All coefficients on the remaining dummies and on the simultaneous feedbacks remain unaltered and the test for the validity of the over-identifying restrictions now distributed as a χ^2 with 137 degrees of freedom, takes a value of 134 with a tail probability of 0.54. Thus the evidence of non-linearities detected in the full sample is confirmed in the sub-sample which excludes the 1992 ERM crisis.

Finally, we augmented our specification to include the effective dollar exchange rate. If bilateral spreads within the ERM are affected by fluctuations in the dollar exchange rate (as documented in Giavazzi and Giovannini, 1991), then omitting this variable would lead us to underestimate interdependence and, therefore, to incorrectly emphasize the role of non-linearities. When including the contemporaneous and lagged effective dollar exchange rate (available at weekly frequency from the FRED database at the Federal Reserve Bank of St. Louis website) we were unable to reject the null that this variables do not significantly enter the structural form.[7]

4. Conclusions

This paper proposes a framework to test for non-linearities in the propagation of financial shocks across countries. We share with Forbes and Rigobon (1999) the

[7]In fact Rigobon (1999) considers two instruments, one using s_{31} and the other using s_{21}. We discuss only one of the two instruments as the argument is symmetrically extended to the second one.

view on the importance of modelling financial interdependence in order to detect non-linearities in the propagation of shocks, and we propose a full-information framework which might be more efficient than the limited information one proposed by Rigobon (2000), in particular when the high-volatility periods are relatively short. This benefit has to be weighted against the cost of more severe identifying restrictions. Our identifying assumption, however, is the one that allows for the highest degree of interdependence: in fact it imposes no restrictions on the contemporaneous feedbacks.

Studying the propagation of devaluation expectations among seven European countries over the period 1988–1992, we were able to reject the null of linearity. We identify a number of country-specific shocks, whose effects on other European markets were significantly non-linear. Our evidence suggests that such non-linearities were a general phenomenon within the ERM. We also find that such non-linearities sometimes imply a change in sign: a widening of the spread on German interest rates in country A associated with a closing of the spread in country B. This evidence suggests that the term 'contagion', often used in the literature, may not be the appropriate one to describe the propagation of financial shocks during periods of market turbulence since it overlooks the phenomenon of 'flight-to-quality' often observed during an international financial crisis.

Our findings are consistent with a large variety of models that describe alternative mechanisms which may lie behind such non-linearities: multiple equilibria due to expectations shifts, liquidity effects, herd behaviour, liquidity problems faced by foreign investors, and macroeconomic similarities, or dissimilarities among countries.

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