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ABSTRACT

This paper applies a full-information technique to test for the presence of contagion across the money markets of ERM member countries. We show that whenever it is possible to estimate a model for interdependence, a test for contagion based on a full information technique is more powerful. We test for the presence of contagion after having identified episodes of country-specific shocks, whose effects on other European markets are significantly different from those predictable from the estimated channels of interdependence. Using data on three-months interest rate spreads on German rates for seven countries over the period 1988-1992, we are unable to reject the null of contagion. Our evidence suggest that contagion within the ERM was a general phenomenon, not limited to a subset of weaker countries, the exception in the sample being France. Our results are mute as to the question of what lies behind these episodes of contagion; they show, however, that it is not always true that one only detects contagion when one applies poor statistical techniques.

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1. Introduction

The observation that financial crises tend to be correlated across markets has prompted a large number of empirical papers aimed at testing whether interdependence (across markets and/or across countries) is enough to explain such correlations, or whether instead what we observe is the symptom of a different phenomenon, commonly called "contagion"–that is a change in the way shocks are transmitted across countries during crises periods.

Cross-market correlations typically increase during a crisis: this, however, should not come as a surprise and is certainly not enough to conclude that what we observe bears evidence of contagion. As shown by Forbes and Rigobon [3], standard measures of cross-market correlations are conditional on market movements over the estimation period: thus, during a period of turmoil, when the volatility of asset prices increases, such measures are upward biased. When one recognizes this point, there often appears to remain little evidence of contagion. This has led Forbes and Rigobon [4] to conclude that cross-market linkages do not change during crises: interdependence is enough to explain comovements across markets, even when these appear to be exceptionally high.

To test for the presence of contagion one should thus proceed in two steps. First, identify the channels through which shocks are normally propagated across markets, by estimating a model of interdependence; next, run a test of the hypothesis that such channels are modified during crises periods.

Implementing step one, however, is difficult when interdependence extends over many countries, or markets, and thus requires the estimation of large models. Rigobon [10] solves this problem using an innovative, limited information technique, based on instrumental variables. This paper shows that whenever the estimation of a model for interdependence is feasible, the use of a full information technique allows the construction of a more powerful test. It also avoids the problems generated by the size of the sample of high-volatility observations, which is typically small.

In the empirical section of this paper we test for the presence of contagion across the countries that were members of the European Exchange Rate Mechanism (ERM) during episodes of market turbulence. We focus on the money market, and in particular on 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. 3-month rates are a good compromise.

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 estimate, over the sample, a common model for interdependence across the money markets of ERM members–the first step in implementing our full information test.

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 Swedisk 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.

2. Estimating interdependence to detect contagion

There is obviously no contagion without a crisis. Thus, in order to detect contagion, we first need to identify a set of shocks that could have been transmitted across countries in an unusual way. In order to identify such shocks, we estimate, over our sample, a model of interdependence among European interest rate spreads. The statistical model we adopt is a reduced-form VAR that describes the joint process generating the 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 $s_1 = R_2 - R_1$ and $s_2 = R_3 - R_1$, the spreads, which reflect expectations of exchange rate depreciation. The conditional distribution of $s_1$and s_2 is described by the following reduced form:

$$\begin{pmatrix} s_{1,t} \\ s_{2,t} \end{pmatrix} = \begin{pmatrix} \pi_{11} & \pi_{12} \\ \pi_{21} & \pi_{22} \end{pmatrix} \begin{pmatrix} s_{1,t-1} \\ s_{2,t-1} \end{pmatrix} + \begin{pmatrix} u_{1,t} \\ u_{2,t} \end{pmatrix} \tag{2.1}$$

$$\left(\begin{matrix} u_{1t} \\ u_{2t} \end{matrix} \mid I_{t-1} \right) \sim \left[\begin{pmatrix} 0 \\ 0 \end{pmatrix}, \begin{pmatrix} \sigma_{1t}^2 & \sigma_{12,t} \\ \sigma_{12,t} & \sigma_2^2 \end{pmatrix} \right]$$

In order to detect the presence, in our dataset, of observations which correspond to episodes of market turbulence, we look for residuals of the above VAR model that are non-normal and heteroscedastic: we do this by running tests of normality and heteroscedasticity of the residuals. We then identify such datapoints by means of a set of dummies, $\mathbf{d}$. The dummies thus identify the periods of crisis. It is then possible to re-specify (2.1) as :

$$\begin{aligned} \begin{pmatrix} s_{1,t} \\ s_{2,t} \end{pmatrix} &= \begin{pmatrix} \pi_{11} & \pi_{12} \\ \pi_{21} & \pi_{22} \end{pmatrix} \begin{pmatrix} s_{1,t-1} \\ s_{2,t-1} \end{pmatrix} + \\ & \left(I + \begin{pmatrix} c_{11} & c_{12} \\ c_{21} & c_{22} \end{pmatrix} \begin{pmatrix} d_{1,t} & 0 \\ 0 & d_{2,t} \end{pmatrix} \right) \begin{pmatrix} u_{1,t}^l \\ u_{2,t}^l \end{pmatrix} \end{aligned} \tag{2.2}$$

$$\left(\begin{matrix} u_{1t}^l \\ u_{2t}^l \end{matrix} \mid I_{t-1} \right) \sim N \left[\begin{pmatrix} 0 \\ 0 \end{pmatrix}, \begin{pmatrix} \sigma_1^2 & \sigma_{12} \\ \sigma_{12} & \sigma_2^2 \end{pmatrix} \right]$$

where $\begin{pmatrix} d_{1,t} \\ d_{2,t} \end{pmatrix}$, the vector of dummies, is partitioned in two blocks according to whether the event generating the turmoil occurred in country 1 or in country 2. Note that the introduction of dummy variables eliminates non-normality and heteroscedasticity from the residuals. We consider additive dummies because the market turbulence periods last at most four consecutive observations.

Having identified the episodes of market turbulance, we then proceed to estimate a structural model of interdependence. Consider, continuing with our example, the following two-equation structural model which we shall assume to be just-identified by the restriction that, in each equation, the own lagged dependent variable is sufficient to capture the structural dynamics:

$$\begin{pmatrix} 1 & -\beta_{12} \\ -\beta_{21} & 1 \end{pmatrix} \begin{pmatrix} s_{1,t} \\ s_{2,t} \end{pmatrix} = \begin{pmatrix} \gamma_{11} & 0 \\ 0 & \gamma_{22} \end{pmatrix} \begin{pmatrix} s_{1,t-1} \\ s_{2,t-1} \end{pmatrix} + \left(I + \begin{pmatrix} a_{11} & a_{12} \\ a_{21} & a_{22} \end{pmatrix} \begin{pmatrix} d_{1,t} & 0 \\ 0 & d_{2,t} \end{pmatrix} \right) \begin{pmatrix} \epsilon_{1,t} \\ \epsilon_{2,t} \end{pmatrix} \tag{2.3}$$

$$\left(\begin{matrix} \epsilon_{1,t} \\ \epsilon_{2,t} \end{matrix} \mid I_{t-1} \right) \sim N \left[\begin{pmatrix} 0 \\ 0 \end{pmatrix}, \begin{pmatrix} \sigma^2_{\epsilon 1} & 0 \\ 0 & \sigma^2_{\epsilon 2} \end{pmatrix} \right]$$

Contagion happens when either $a_{12} \neq 0$, or $a_{21} \neq 0$, that is when, during periods of turmoil, the parameters which describe interdependence (β_{12} and β_{21}) are not enough to describe the transmission of shocks across countries. Note however that under the null of interdependence only, no contagion turmoils affects all reduced form relationships.

Within the framework of our example it is easy to make two related points. First, simple correlations are the wrong indicator to detect contagion. As shown by Rigobon [9] and by Forbes and Rigobon [3], a change in the (conditional and unconditional) correlation between $s_{1,t}$ and $s_{2,t}$ does occur during a crisis, quite independently of contagion. Consider, for example, conditional correlations and define:

$$\rho \equiv \frac{Cov\left(s_{1,t}, s_{2,t} \mid I_{t-1}\right)}{\sqrt[2]{Var\left(s_{1,t} \mid I_{t-1}\right) Var\left(s_{2,t} \mid I_{t-1}\right)}}$$

the correlation between the two interest rates spreads. In the low volatility periods, that is outside the observations identified by the **d** dummies, we have:

$$\begin{pmatrix} s^l_{1,t} \\ s^l_{2,t} \end{pmatrix} - E \left(\begin{matrix} s^l_{1,t} \\ s^l_{2,t} \end{matrix} \mid I_{t-1} \right) = \begin{pmatrix} u^l_{1,t} \\ u^l_{2,t} \end{pmatrix}$$

while in the high volatility period:

$$\begin{aligned} \begin{pmatrix} s^h_{1,t} \\ s^h_{2,t} \end{pmatrix} - E \left(\begin{matrix} s_{h1,t} \\ s_{h2,t} \end{matrix} \mid I_{t-1} \right) &= \left(I + \begin{pmatrix} c_{11} & c_{12} \\ c_{21} & c_{22} \end{pmatrix} \begin{pmatrix} d_{1,t} & 0 \\ 0 & d_{2,t} \end{pmatrix} \right) \begin{pmatrix} u^l_{1,t} \\ u^l_{2,t} \end{pmatrix} \\ &= \begin{pmatrix} u^h_{1,t} \\ u^h_{2,t} \end{pmatrix} \end{aligned}$$

Next compute ρ^l, and ρ^h, the conditional correlations, in the high and low volatility periods respectively. It is straightforward to check that *even under the null of no contagion* $\rho^h > \rho^l$. Rigobon [9], for example, considers the simple case in which $a_{11} = a_{21} = 0$, i.e. all turmoil is generated in country 2. The null of no contagion implies $a_{12} = 0$. In this case we have:

$$
\begin{aligned}
\rho^l &= \frac{\beta_{12}\sigma_{\epsilon 2}^2}{\sqrt[2]{\left[(b_{12})^2 \sigma_2^2 + \sigma_{\epsilon 1}^2\right] \sigma_{\epsilon 2}^2}} \\
&= \frac{1}{\sqrt[2]{1 + \frac{\sigma_1^2}{\beta_{12}^2 \sigma_{\epsilon 2}^2}}}
\end{aligned}
$$

$$
\begin{aligned}
\rho^h &= \frac{\beta_{12}\sigma_{\epsilon 2}^2 \left(1 + a_{22} d_{2t}\right)^2}{\sqrt[2]{\left[(\beta_{12})^2 \sigma_{\epsilon 2}^2 \left(1 + a_{22} d_{2t}\right)^2 + \sigma_{\epsilon 1}^2\right] \left(\sigma_{\epsilon 2}^2 \left(1 + a_{22} d_{2t}\right)^2\right)}} \\
&= \frac{1}{\sqrt[2]{1 + \frac{\sigma_{\epsilon 1}^2}{\beta_{12}^2 \sigma_{\epsilon 2}^2 (1 + a_{22} d_{2t})^2}}}
\end{aligned}
$$

and $\rho^h > \rho^l$.

Second, we can compare our full-information approach with the instrumental variables method proposed by Rigobon. [10] In that method one controls by estimating β_{12} and β_{21} using a limited-information technique for interdependence in order to detect contagion. The technique hinges on splitting the sample into high and low volatility periods: an instrument is then constructed whose validity is guaranteed under the null of no contagion. The test for contagion is then simply a test of the validity of the instruments. To illustrate this procedure within our example, consider again the simple case in which $a_{11} = a_{21} = 0$. In this case the dependent variable is split into high and low volatility observations in the following way:

$$
\begin{aligned}
\begin{pmatrix} 1 & -\beta_{12} \\ -\beta_{21} & 1 \end{pmatrix} \begin{pmatrix} s_{1,t}^h \\ s_{2,t}^h \end{pmatrix} &= \begin{pmatrix} \gamma_{11} & 0 \\ 0 & \gamma_{22} \end{pmatrix} \begin{pmatrix} s_{1,t-1}^h \\ s_{2,t-1}^h \end{pmatrix} + \\
& \left(I + \begin{pmatrix} 0 & a_{12} \\ 0 & a_{22} \end{pmatrix} \begin{pmatrix} d_{1,t} & 0 \\ 0 & d_{2,t} \end{pmatrix} \right) \begin{pmatrix} \epsilon_{1,t} \\ \epsilon_{2,t} \end{pmatrix}
\end{aligned}
\tag{2.4}
$$

$$\begin{pmatrix} 1 & -\beta_{12} \\ -\beta_{21} & 1 \end{pmatrix} \begin{pmatrix} s_{1,t}^{l} \\ s_{2,t}^{l} \end{pmatrix} = \begin{pmatrix} \gamma_{11} & 0 \\ 0 & \gamma_{22} \end{pmatrix} \begin{pmatrix} s_{1,t-1}^{l} \\ s_{2,t-1}^{l} \end{pmatrix} + + \begin{pmatrix} \epsilon_{1,t} \\ \epsilon_{2,t} \end{pmatrix} \tag{2.5}$$

Consider now the following instrument for s_2 :

$$\mathrm{w}_t = \begin{pmatrix} \frac{\mathbf{s}_{2,t}^{h}}{T^h} \\ -\frac{\mathbf{s}_{2,t}^{l}}{T^l} \end{pmatrix},$$

Using w_t as an instrument for s_2 leads to the following just-identified instrumental variables estimator of the interdependence parameter β_{12}:

$$\hat{\beta}_{12} = (w' s_2)^{-1} w' s_1$$

The two usual conditions for validity and consistency of the IV estimator are checked by looking at the probability limits of $(w' s_2)$ and $(w' \epsilon_1^*)$, where $\boldsymbol{\epsilon}_1^* = a_{12} d_{2,t} \boldsymbol{\epsilon}_2 + \boldsymbol{\epsilon}_1$

$$\begin{aligned} p\lim (w' s_2) &= p\lim \frac{1}{T^h} s_{2,t}^{h} s_{2,t}^{h} - p\lim \frac{1}{T^l} s_{2,t}^{l\prime} s_{2,t}^{l} \\ &= \left(-\frac{\beta_{21} a_{12}}{-1 + \beta_{12}\beta_{21}} d_{2t} - \frac{a_{22} d_{2t}}{-1 + \beta_{12}\beta_{21}} \right)^2 \sigma_{\epsilon 2}^2 \\ p\lim (w' \boldsymbol{\epsilon}_1^*) &= \left(\frac{\beta_{21} a_{12}}{1 - \beta_{21}\beta_{12}} d_{2t}^2 - \frac{a_{12} a_{22}}{1 - \beta_{21}\beta_{12}} d_{2t}^2 \right) \sigma_{\epsilon 2}^2 \end{aligned}$$

The validity of w_t as an instrument is guaranteed under $\mathrm{H}_0\,(a_{12} = 0)$, while its efficiency depends on the degree of heteroscedasticity between low and high volatility observations . Within this framework, contagion can be tested applying a Hausman [5]-type test for the validity of the instruments. The beauty of this approach depends on the fact that it does not require variables other than $s_{2,t}$ to implement the IV estimator.

Skipping the estimation of a structural model from interdependence has obvious benefits; however, it also implies a non-negligible cost. There are cases in which the configuration of parameters in the structural model is such that the IV procedure would lead to reject the null of contagion when it is true. In our example, for instance, $a_{12} = 0$ is not the unique solution of $p\lim (w' \epsilon_1^*) = 0$.

3. Looking for contagion in the ERM

In this section we implement the technique outlined above to study the propagation of financial shocks across European money markets during the ERM. Are the normal channels of interdependence enough to explain the way such schocks were transmitted from one market to another, or is there evidence of contagion? 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 contagion.

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 raw data on Euro spreads are shown in Figure 1. Only German and Dutch rates appear to share a common stochastic trend; the other spreads show a remarkable degree of persistence (this characteristic of the data will be confirmed by the econometrics.)

Insert Figure 1

The estimation of a first-order VAR for interest rates spreads, as described in 2.1, features 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 2.2. Estimates of the reduced-from coefficients are shown in tables 1 and 2 where report the coefficients on the twenty-six dummies separately. The residuals obtained from a VAR that includes the dummies, reported in Figure 2, show no apparent evidence of correlation, nor of heteroscedasticy (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, thus confirming the visual impression of Figure 1. 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[7] procedure, we find evidence in favour for the stationarity of the spread only for Holland [1]–a result which suggests, except for Holland, a low credibility of the ERM over our sample.

The observations identified by each of the twenty-six dummies can be traced to a piece of news relevant for financial markets in the ERM: we describe each piece of news in Table 2 below the corresponding observation/s. Based on this information, shocks are defined "local" or "common", depending on whether they hit a single country, 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 2 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 variable that identifies the Dutch shock of May 17, 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 could signal contagion, but could also simply be the effetct of normal interdependence.

To detect contagion we thus estimate a structural model to control for interdependence. This will allow us to test the significance of local shocks outside the country where they have originated.

3.2. Modelling interdependence and testing for contagion

To model interdependence we estimate a structural simultaneous model for the determination of interest rate spreads.

We achieve identification by restricting the lag structure of the model.[2] Our initial identifying assumption is that only the lagged value of the dependent variable is allowed to enter each equation, while we allow for all possible simultaneous feedbacks. The validity of this set of just-identifying restrictions that is using

[1] 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[8].

[2] 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,Stock and Watson[11]).

the lagged dependent variables as instruments for the contemporaneous values–is supported by the evidence in our reduced form which showed that only the own lagged dependent variable is significant in each equation. We then move from a just-identified structure to an over-identified model by restricting to zero all the contemporaneous effects that are not significantly different from zero. Estimates of our structural model of interdependence are reported in Tables 3 and 4 (as for the reduced form we report separately, in Table 4, the coefficients on the dummies).

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 contagion is contained in Table 4. Under the null of no contagion the dummies associated with local shocks should be significant only in the country where the shock originates. The null is rejected in eleven episodes of local shocks. For instance, interdependence is not enough to explain the transmission to Denmark, Italy, Holland and Belgium of the local Swedish shocks of November-December 1990. The same is true, for the transmission to Spain of the local Italian and Dutch shocks occurring, respectively, in March and May 1989. With the exception of France, contagion affects all countries in our sample. Note that contagion implies a change in the international transmission of shocks, which normally amounts to a stronger effect in the same direction, but occasionally implies a significant effect in the opposite direction. For instance, on the occasion of the Italian shock of July 8th, 1992, which rasied the Italian spread by 135 basis points, the Dutch spread narrowed by 17 basis points.

4. Conclusions

This paper proposes a framework to test for contagion in situations where it is possible to specify and estimate a model for interdependence across financial markets. We share with Forbes and Rigobon[4] the view on the importance of modelling interdepedence in order to test for contagion. We show, however, that the limited information framework proposed by Rigobon [10] may not reject the null of no contagion when it is false. This is because the limited information approach does not allow to test explicitly the over-identifying restrictions imposed by the structural model under the null of interest.

Using data on three-month interest rate spreads on German rates for seven European countries over the period 1988-1992, we were able to reject the null of no contagion. We identify a number of country-specific shocks, whose effects on other European markets were significantly different from those predictable from the estimated channels of interdependence. Our evidence suggests that contagion within the ERM was a general phenomenon, not limited to a subset of weaker countries, with the only exception of France.

This evidence is consistent with a large variety of models that describe alternative mechanisms through which contagion may occurr: multiple equilibria due to expectations shifts, liquidity effects, herd behaviour, liquidity problems faced by foreign investors, macroeconomic similarities.

While mute as to the question of what lies behind these episodes of contagion, our findings run against the view that evidence of contagion is only the result of the application of poor statistical techniques.

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 (0.043)	0.63 (0.06)	−0.02 (0.03)	0.001 (0.009)	0.006 (0.01)	−0.04 (0.02)	−0.02 (0.008)	0.10 (0.04)
s_t^{FR}	−0.02 (0.06)	−0.17 (0.09)	0.90 (0.04)	0.013 (0.013)	0.015 (0.015)	0.028 (0.023)	−0.02 (0.012)	0.06 (0.05)
s_t^{IT}	0.03 (0.11)	−0.21 (0.16)	−0.06 (0.07)	1.03 (0.03)	−0.04 (0.03)	0.07 (0.04)	0.01 (0.02)	−0.06 (0.10)
s_t^{ES}	0.09 (0.07)	−0.2 (0.10)	0.2 (0.05)	0.1 (0.2)	0.96 (0.02)	−0.0004 (0.03)	−0.0004 (0.01)	0.09 (0.06)
s_t^{DK}	0.03 (0.08)	−0.10 (0.12)	−0.018 (0.05)	−0.002 (0.02)	0.02 (0.02)	0.97 (0.03)	−0.03 (0.02)	0.04 (0.07)
s_t^{SW}	−0.07 (0.12)	0.06 (0.2)	−0.03 (0.08)	−0.06 (0.03)	0.004 (0.03)	0.10 (0.05)	0.95 (0.03)	−0.17 (0.10)
s_t^{BG}	−0.05 (0.05)	−0.02 (0.08)	−0.05 (0.03)	0.01 (0.01)	0.01 (0.01)	0.04 (0.02)	−0.02 (0.01)	0.90 (0.05)
Testing for vector autocorrelation of the residuals (lags 1 to 7): $F(343{,}984) = 0.97$ [0.61]								
Testing for vector heteroscedasticity 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

Note: the model also includes the twenty-six dummy variables reported in Table 2.

Table 2: Dummies in the reduced form model of Table 1							
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** (0.10)	0.67** (0.15)	−0.33 (0.28)	0.81** (0.18)	0.59** (0.20)	0.51** (0.25)	0.54** (0.14)
common shock	Bundesbank raises policy rates						
08/03/89	0.07 (0.10)	0.024 (0.15)	0.94** (0.28)	0.53** (0.18)	0.29 (0.20)	0.50** (0.25)	0.11 (0.14)
local shock	Bank of Italy raises rates after bad trade deficit data						
03/05/89	0.29** (0.10)	−0.0212 (0.15)	−0.09 (0.28)	0.39** (0.18)	−0.03 (0.20)	0.02 (0.25)	−0.03 (0.14)
local shock	Dutch government resigns						
17/05/89	−0.10 (0.10)	−0.20 (0.15)	−0.14 (0.28)	−0.72** (0.18)	−0.33 (0.20)	−0.56** (0.25)	−0.22 (0.14)
local shock	Spain announces sharp cuts in public spending						
05/07/89	0.023 (0.10)	0.34** (0.15)	−0.10 (0.28)	0.64** (0.18)	0.34 (0.20)	−0.05 (0.25)	−0.16 (0.14)
common shock	US dollar collapses						
11/10/89	0.18 (0.10)	0.093 (0.15)	−0.02 (0.28)	0.10 (0.18)	0.47** (0.20)	0.03 (0.25)	0.61** (0.14)
common shock	Bundesbank raises interest rates						
18/10/89	−0.17 (0.10)	0.035 (0.15)	0.07 (0.28)	−0.23 (0.18)	2.60** (0.20)	−0.42** (0.25)	−0.04 (0.14)
local shock	Danish Krona hits the bottom of the ERM band						
25/10/89	−0.14 (0.10)	0.012 (0.15)	0.96** (0.28)	0.12 (0.18)	−1.62** (0.20)	−0.24 (0.25)	0.03 (0.14)
local shock	Bundesbank intervenes to prop up the Krona						
01/11/89	−0.12 (0.10)	−0.087 (0.15)	−1.48** (0.28)	−0.26 (0.18)	0.08 (0.20)	−0.37 (0.25)	−0.28 (0.14)
common shock	Bundesbank injects liquidity in the system						
21/03/90	−0.065 (0.10)	−0.018 (0.15)	0.08 (0.28)	0.14 (0.18)	−0.03 (0.20)	2.19** (0.25)	0.09 (0.14)
local shock	Sveriges Riksbank raises rates to stem capital outflows						
14/11/90	−0.16 (0.10)	−0.092 (0.15)	0.71** (0.28)	−0.04 (0.18)	−0.27 (0.20)	1.86** (0.25)	−0.19 (0.14)
21/11/90	0.33** (0.10)	−0.08 (0.15)	0.81** (0.28)	0.15 (0.18)	0.14 (0.20)	1.21** (0.25)	0.15 (0.14)
05/12/90	−0.19** (0.10)	−0.07 (0.15)	−0.44 (0.28)	−0.15 (0.18)	−0.40** (0.20)	−0.94** (0.25)	0.33** (0.14)
12/12/90	0.46** (0.10)	0.02 (0.15)	0.27 (0.28)	0.15 (0.18)	0.44** (0.20)	0.69** (0.25)	0.23 (0.14)
local shock	Swedish recession, exchange rate pressure & interest rates hike						

Table 2 (continued)							
Dummies	dep.var.						
	s_t^{NL}	s_t^{FR}	s_t^{IT}	s_t^{ES}	s_t^{DK}	s_t^{SW}	s_t^{BG}
27/03/91	0.10 (0.10)	−0.04 (0.15)	−0.19 (0.28)	−0.58** (0.18)	0.09 (0.20)	0.45** (0.25)	0.01 (0.14)
local shock	Bank of Spain intervenes buying French Francs						
11/12/91	0.065 (0.10)	0.07 (0.15)	−0.08 (0.28)	0.08 (0.18)	0.023 (0.20)	1.5 ** (0.25)	0.13 (0.14)
18/12/91	0.10 (0.10)	−0.08 (0.15)	0.01 (0.28)	0.03 (0.18)	0.12 (0.20)	0.66** (0.25)	0.03 (0.14)
25/12/91	0.14 (0.10)	0.30** (0.15)	−0.11 (0.28)	−0.06 (0.18)	0.026 (0.20)	0.94** (0.25)	0.10 (0.14)
02/01/92	0.056 (0.10)	0.24 (0.15)	0.21 (0.28)	0.11 (0.18)	0.25 (0.20)	1.02** (0.25)	0.01 (0.14)
08/01/92	0.19** (0.10)	−0.12 (0.15)	−0.36 (0.28)	0.13 (0.18)	−0.03 (0.20)	−0.81** (0.25)	−0.12 (0.14)
local shock	Swedish exchange rate crisis						
08/07/92	−0.17** (0.10)	0.03 (0.15)	1.18** (0.28)	0.05 (0.18)	0.06 (0.20)	−0.10 (0.25)	−0.10 (0.14)
22/07/92	0.018 (0.10)	0.04 (0.15)	2.27** (0.28)	0.25 (0.18)	0.25 (0.20)	0.28 (0.25)	0.07 (0.14)
29/07/92	−0.028 (0.10)	−0.03 (0.15)	−1.58** (0.28)	0.25 (0.18)	0.02 (0.20)	0.28 (0.25)	−0.07 (0.14)
05/08/92	0.104 (0.10)	−0.09 (0.15)	−1.85** (0.28)	−0.06 (0.18)	−0.04 (0.20)	0.28 (0.25)	0.01 (0.14)
local shock	Italian political and fiscal crisis						
26/08/92	0.056 (0.10)	0.31** (0.15)	0.49 (0.28)	0.23 (0.18)	0.07 (0.20)	1.62** (0.25)	−0.05 (0.14)
09/09/92	0.076 (0.10)	0.16 (0.15)	2.5 ** (0.28)	0.05 (0.18)	0.031 (0.20)	9.35** (0.25)	0.04 (0.14)
common shock	ERM crisis						

Table 3: 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 (0.01)	0.71 (0.04)							0.047 (0.01)
s_t^{FR}	0.0005 (0.02)	0.93 (0.02)					0.038 (0.01)		
s_t^{IT}	0.056 (0.05)	0.98 (0.01)							
s_t^{ES}	0.03 (0.03)	0.98 (0.01)							
s_t^{DK}	0.03 (0.01)	0.97 (0.01)							
s_t^{SW}	0.08 (0.045)	0.93 (0.01)					0.06 (0.02)		
s_t^{BG}	−0.02 (0.01)	0.93 (0.02)					0.035 (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

Note: The model also includes the twenty six dummy variables reported in Table 4.

Table 4: Dummies in structural model of Table 3

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 (0.10)	0.74 (0.14)		0.88 (0.17)	0.65 (0.18)	0.45 (0.24)	0.57 (0.12)
common							
08/03/89			0.73 (0.25)	0.34 (0.15)			
local							
03/05/89	0.30 (0.09)			0.37 (0.15)			
local							
17/05/89				−0.55 (0.15)			
local							
05/07/89		0.38 (0.12)		0.63 (0.15)			
common							
11/10/89					0.33 (0.16)		0.51 (0.10)
common							
18/10/89					2.70 (0.16)	−0.37 (0.22)	
local							
25/10/89			1.09 (0.26)		−1.62 (0.17)		
local							
01/11/89			−1.3 (0.25)				
common							
21/03/90						2.19 (0.2)	
local							
14/11/90			0.86 (0.25)			2.07 (0.2)	
local							
21/11/90	0.2 (0.09)		0.82 (0.25)			1.00 (0.2)	0.44 (0.11)
local							
05/12/90	−0.25 (0.09)				−1.35 (0.18)	−0.8 (0.2)	0.39 (0.10)
local							
12/12/90	0.31 (0.08)					0.45 (0.28)	
local							

Table 4 (continued)							
Dummies	dep.var.						
	s_t^{NL}	s_t^{FR}	s_t^{IT}	s_t^{ES}	s_t^{DK}	s_t^{SW}	s_t^{BG}
27/03/91				-0.63 (0.15)		0.43 (0.21)	
local							
11/12/91						0.96 (0.21)	
local							
18/12/91						0.55 (0.21)	
local							
25/12/91		0.24 (0.12)				0.87 (0.21)	
local							
02/01/92						0.91 (0.21)	
local							
08/01/92	0.18 (0.08)					-0.88 (0.21)	
local							
08/07/92	-0.17 (0.08)		1.35 (0.25)				
local							
22/07/92			2.35 (0.25)				
local							
29/07/92			-1.37 (0.26)				
local							
05/08/92			-1.68 (0.25)				
local							
26/08/92		0.25 (0.12)				1.6 (0.21)	
common							
09/09/92			2.7 (0.25)			9.04 (0.21)	
common							

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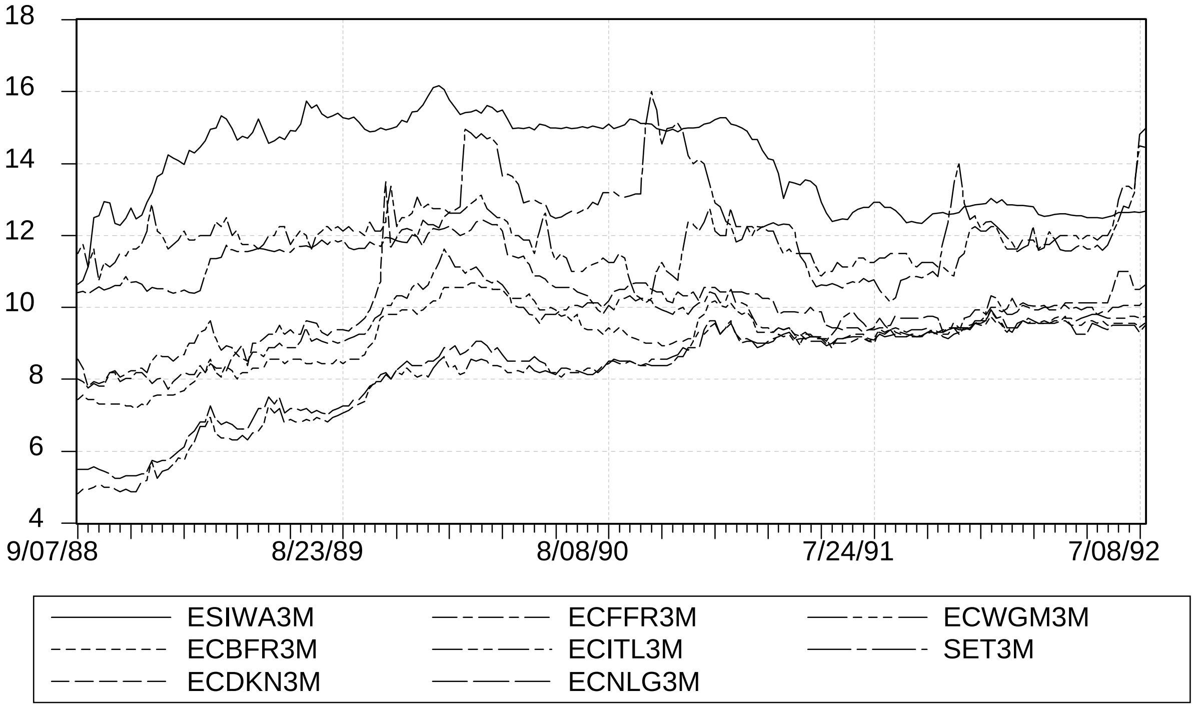

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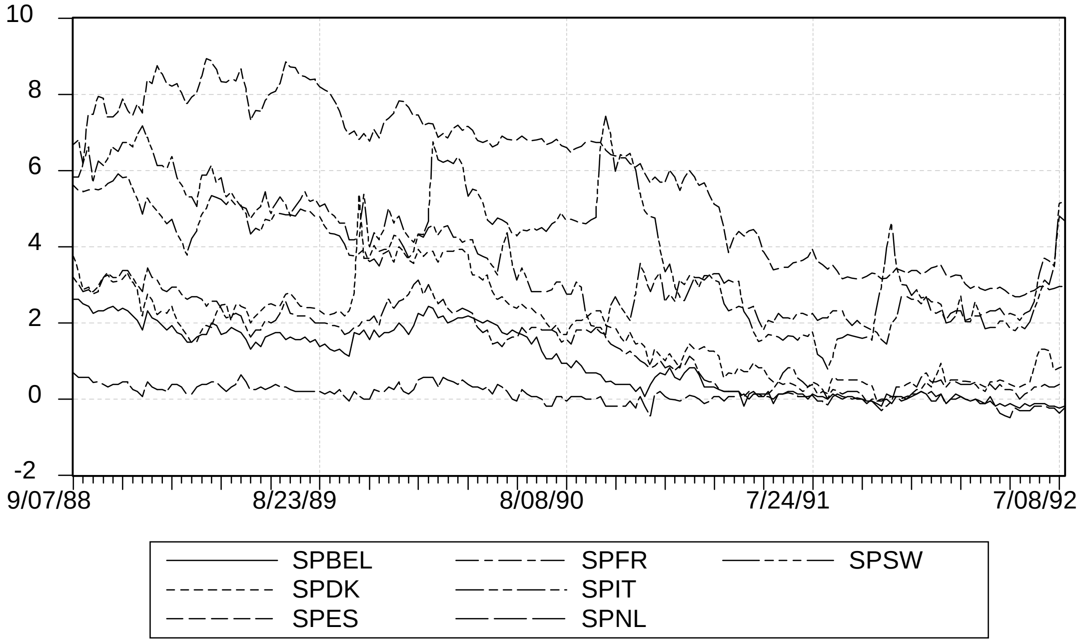

Figure 2: the spreads on 3month Euro-DM interest rates. 1988-1992

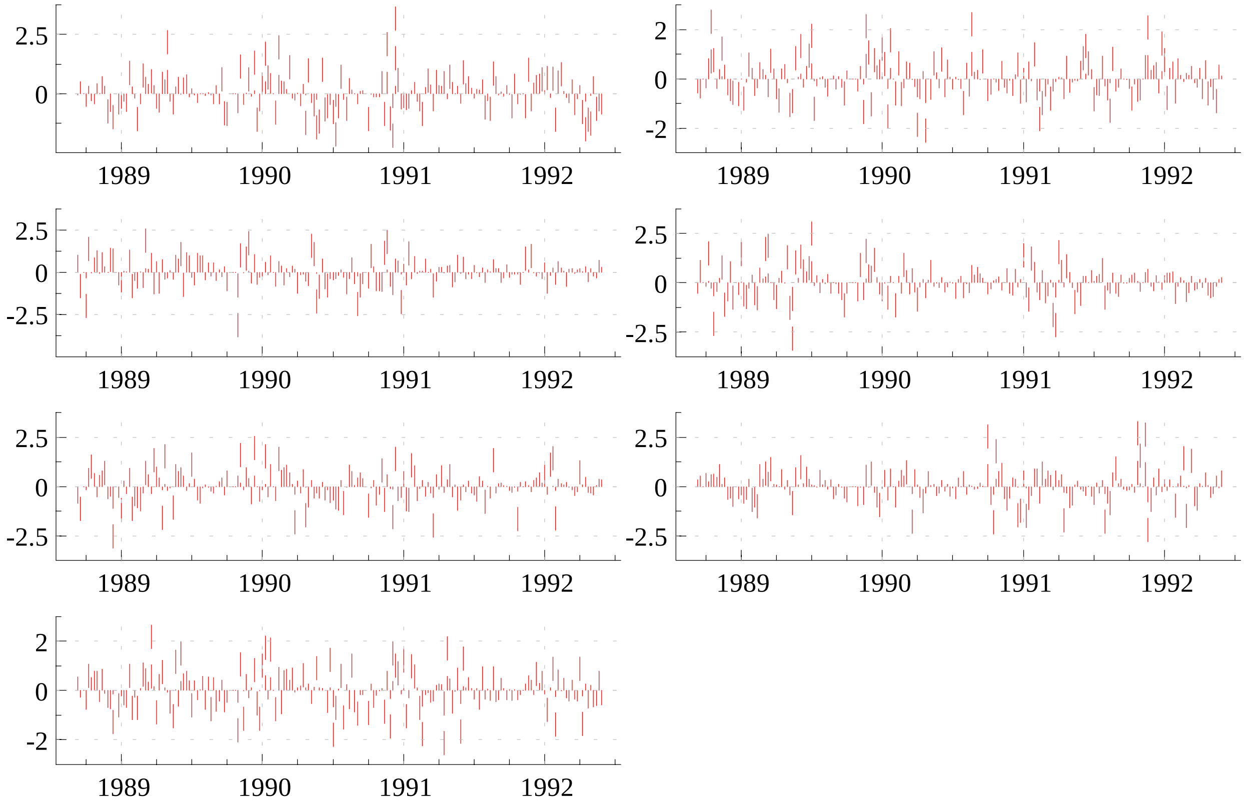

Figure 3: Residuals form the reduced form for European interest rates spreads (The Netherlands, France, Italy, Spain Sweden, Denmark and Belgium) on German rates