Explaining Co-movements Between Stock Markets: the case of US and Germany*

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December 10, 2002

Abstract

We explain co-movements between stock markets by explicitly considering the distinction between interdependence and contagion. We propose and implement a full information approach on data for US and Germany to provide answers to the following questions:

(i) is there long-term interdependence between US and German stock markets ?

(ii) Is there short-term interdependence and contagion between US and German stock markets, i.e do short term fluctuations of the US share prices spill over to German share prices and is such co-movement unstable over high volatility episodes?

Our answers are, respectively, no the the first question and yes to the second one.

JEL Classification: F30, F40, G15

Keywords: Contagion, Stock Market, Interdependence, Structural models

*This paper has benefited from comments of two anonymous referees and seminar participants at INSEAD, CIDE, Bocconi University and University of Naples. We also thanks our discussants Fabio-Cesare Bagliano and Diego Lubian. Roberto Botter contributed comments, discussions and excellent research asssitance. Addres for correspondace: *Carlo A. Favero, IGIER- Università Bocconi, Via Salasco 5, 20136 Milan , Italy, e-mail: carlo.favero@uni-bocconi.it*

1 Introduction

Measuring co-movements between stock markets is a widely debated issue. Academic studies have shown that correlations between international equity markets vary strongly over time[1], and suggested two main distinct explanations for this phenomenon. The first is based on the belief that the transmission mechanism is stable, while the features of shocks (global vs idiosyncratic) vary over time. In some periods global shocks do not occur and equity markets are driven by country-specific factors. As national business cycles are not well synchronized, all markets tend to move independently. In other periods all equity markets are globally affected by the same shocks and therefore their tendency to co-move increases. The alternative explanation relies upon the idea that periods of turbulence are characterized by the occurrence of shocks of unusual dimension, which may come along with structural breaks in their transmission mechanism. The empirical literature on the transmission of financial shocks (Rigobon, 1999) has recently formalized the distinction between the concepts of contagion and interdependence. The latter accounts for the existence of cross-market linkages, while contagion consists in modifications of such linkages during turbulent periods. Consider the case of the US and German stock markets: a strong co-movement of German and US equity prices in presence of unusual fluctuations in the US stock market is compatible both with interdependence and contagion. We have interdependence if the observed comovement is in line with the historically measured simultaneous feedback between the two markets, while we have contagion when a change in the volatility of the US market (the disease) generates a structural break in the parameters measuring interdependence between US and German markets.

Identifying contagion from interdependence is important for their differ-

[1]See Forbes and Rigobon(1998), Karoly and Stultz(1996), Lee and Kim(1993), Lin, Engle and Ito(1994), Longin and Solnik(1995, 2000).

ent implications on asset allocation and on optimal economic policy (see for instance Bernanke and Gertler (1999), Rigobon and Forbes (2002)).

Correlation between stock markets has been traditionally used when measuring co-movements and defining contagion. The earliest studies by King and Wadhwani (1990) and Bertero and Mayer (1990) presented and discussed the evidence of changes in unconditional covariances and correlations between stock returns on high-frequency data around the october 1987 crash. Since then, many authors proposed different ways of testing the stability of (conditional) correlations, such as using ARCH and GARCH models (see Longin and Solnik (1995) and Edwards and Susmel (2000)), cointegration (again, Longin and Solnik (1995), Kasa(1992), Serletis and King(1997)), or switching regimes (see Hassler (1995) and Edwards and Susmel (2000)). This traditional approach has been recently criticized by Rigobon and Forbes (2002). It is easily shown that in a structural model featuring constant interdependence across countries, cross-market correlations are bound to increase in a period of turmoil, when stock market volatility increases. Hence the evidence of changing patterns of correlations cannot be used to directly test for contagion. Rigobon and Forbes consider the 1997 East Asian crisis, the 1994 Mexican Peso crisis and the 1987 US stock market crash to show that unadjusted correlation coefficients support the contagion hypothesis, while tests based on coefficients adjusted for interdependence find virtually no-contagion. Alternative ways of correcting tests on correlations have been suggested, amongst the others, by Boyer et al. (1999) and Loretan and English (2000), that rely on normality of stock returns[2], and by Longin and Solnik (2001), who apply extreme value theory to conditional correlation coefficients and generalize their results for a wide class of returns distributions.

An innovative methodology to test for contagion in presence of interdependence has been proposed by Rigobon (1999) through the implementation

[2]See Corsetti et al. (2002) for a survey and further contributions along this line of research.

of an IV procedure. This strand of research crucially hinges on structural modelling of interdependence, with the adoption of a limited information approach.

This paper extends the limited information approach and test the hypothesis of “no contagion, only interdependence” through the full information estimation of a small co-integrated structural model, built following the LSE econometric methodology (see Hendry (1995)). Our measure of co-movements distinguishes between long-run and short-run dynamics for equity prices on different markets.

We concentrate on US and German stock markets, and consider a sample of monthly data spanning from January 1980 to September 2002. As a first step, we estimate a general reduced form VAR model on six variables (US and German share prices, earnings, and redemption yields on 10-year benchmark bonds). We remove non-normality and heteroscedasticity from residuals by including in the specification a number of point dummies. Having obtained a valid specification for the VAR we perform cointegration to identify long-run equilibria among the selected variables and attribute an exogeneity status to four of them (earnings and long-term interest rates). Subsequently we formulate a bivariate Vector Error Correction model, for the two endogenous variables, i.e. equity prices in US and Germany. Finally, we proceed to specify a structural model of interdependence and test for no contagion. A structural model is identified by assuming a lower triangular pattern of simultaneous feedbacks between US and German stock markets. On this model we test the further restrictions implied by the null of no contagion.

The paper is organized as follows. Section 2 explains our general-to-specific full-information approach to test for contagion, and compares it to alternative methodologies proposed by Forbes and Rigobon (2002), and by Rigobon (1999). Section 3 illustrates our empirical specification and contains a discussion of our analysis of long-run interdependence based on cointegration. Section 4 considers the short-run dynamics and illustrates how we attribute co-movements to interdependence and contagion. Section 5 concludes.

2 Estimating interdependence and contagion with small structural models

We consider the consensus definition of contagion as a change in the international propagation of shocks caused by some country specific factor. In the recent empirical literature on the international propagation of shocks such factor is usually interpreted as a crisis, identified by a local shock of different magnitude (usually paired with a change in the volatility of shocks). Measuring contagion requires some (structural) estimate of the mechanism of international propagation of shocks and the identification of a crisis.

To achieve this purpose we start from a reduced form VAR specification for the logarithms of US and German share prices, $LP_{Ger,t}, LP_{US,t}$ and the vectors of variables candidate to determine their equilibrium: $\mathbf{X}_{Ger,t}$, $\mathbf{X}_{US,t}$. For the sake of exposition, we consider a first order process, although our empirical model features higher order dynamics.

$$\begin{pmatrix} LP_{Ger,t} \\ LP_{US,t} \\ \mathbf{X}_{Ger,t} \\ \mathbf{X}_{US,t} \end{pmatrix} = \begin{pmatrix} \pi_{11} & \pi_{12} & \boldsymbol{\pi}'_{13} & \boldsymbol{\pi}'_{14} \\ \pi_{21} & \pi_{22} & \boldsymbol{\pi}'_{23} & \boldsymbol{\pi}'_{24} \\ \pi_{31} & \pi_{32} & \boldsymbol{\pi}_{33} & \boldsymbol{\pi}_{34} \\ \pi_{41} & \pi_{42} & \boldsymbol{\pi}_{43} & \boldsymbol{\pi}_{44} \end{pmatrix} \begin{pmatrix} LP_{Ger,t-1} \\ LP_{US,t-1} \\ \mathbf{X}_{Ger,t-1} \\ \mathbf{X}_{US,t-1} \end{pmatrix} + \begin{pmatrix} v_{1,t} \\ v_{2,t} \\ \mathbf{v}_{3,t} \\ \mathbf{v}_{4,t} \end{pmatrix}$$

$$\begin{pmatrix} v_{1,t} \\ v_{2,t} & \mid I_{t-1} \\ \mathbf{v}_{3,t} \\ \mathbf{v}_{4,t} \end{pmatrix} \sim \left[\begin{pmatrix} 0 \\ 0 \end{pmatrix}, \Sigma_t \right] \tag{1}$$

Note that residuals from our baseline VAR specification are heteroscedastic. This reflects the presence in our data of observations which correspond to periods of turmoil. By using tests of normality and heteroscedasticity of residuals as a guiding criterion, it is then possible to re-specify (1) as :

$$\begin{pmatrix} LP_{Ger,t} \\ LP_{US,t} \\ \mathbf{X}_{Ger,t} \\ \mathbf{X}_{US,t} \end{pmatrix} = \begin{pmatrix} \pi_{11} & \pi_{12} & \boldsymbol{\pi}'_{13} & \boldsymbol{\pi}'_{14} \\ \pi_{21} & \pi_{22} & \boldsymbol{\pi}'_{23} & \boldsymbol{\pi}'_{24} \\ \pi_{31} & \pi_{32} & \boldsymbol{\pi}_{33} & \boldsymbol{\pi}_{34} \\ \pi_{41} & \pi_{42} & \boldsymbol{\pi}_{43} & \boldsymbol{\pi}_{44} \end{pmatrix} \begin{pmatrix} LP_{Ger,t-1} \\ LP_{US,t-1} \\ \mathbf{X}_{Ger,t-1} \\ \mathbf{X}_{US,t-1} \end{pmatrix} + (I + \Psi D) \begin{pmatrix} u_{1,t} \\ u_{2,t} \\ \mathbf{u}_{3,t} \\ \mathbf{u}_{4,t} \end{pmatrix} \tag{2}$$

$$\begin{pmatrix} u_{1,t} & \\ u_{2,t} & \mid I_{t-1} \\ \mathbf{u}_{3,t} & \\ \mathbf{u}_{4,t} & \end{pmatrix} \sim N\left[\begin{pmatrix} 0 \\ 0 \end{pmatrix}, \Sigma \right]$$

$$\Psi = \begin{pmatrix} \psi_{11} & \psi_{12} & \boldsymbol{\psi}'_{13} & \boldsymbol{\psi}'_{14} \\ \psi_{21} & \psi_{22} & \boldsymbol{\psi}'_{23} & \boldsymbol{\psi}'_{24} \\ \psi_{31} & \psi_{32} & \boldsymbol{\psi}'_{33} & \boldsymbol{\psi}'_{34} \\ \psi_{41} & \psi_{42} & \boldsymbol{\psi}'_{43} & \boldsymbol{\psi}'_{44} \end{pmatrix}$$

$$D = \begin{pmatrix} d_{1,t} & 0 & 0 & 0 \\ 0 & d_{2,t} & 0 & 0 \\ 0 & 0 & \mathbf{d}_{3,t} & 0 \\ 0 & 0 & 0 & \mathbf{d}_{4,t} \end{pmatrix}$$

where the vectors of dummies $\mathbf{d}_{i,t}$ are identified in order to filter non-normality out of the original residuals. The coefficients in the matrix Ψ allow the removal of outliers.

On the basis of this specification we proceed to cointegration analysis and reparameterise our system as follows:

$$\begin{pmatrix} \Delta LP_{Ger,t} \\ \Delta LP_{US,t} \\ \Delta \mathbf{X}_{Ger,t} \\ \Delta \mathbf{X}_{US,t} \end{pmatrix} = \Pi \begin{pmatrix} LP_{Ger,t-1} \\ LP_{US,t-1} \\ \mathbf{X}_{Ger,t-1} \\ \mathbf{X}_{US,t-1} \end{pmatrix} + (I + \Psi D) \begin{pmatrix} u_{1,t} \\ u_{2,t} \\ \mathbf{u}_{3,t} \\ \mathbf{u}_{4,t} \end{pmatrix}$$

where the matrix Π describes the long-run properties of the system. In case of cointegration, there exist stationary combinations of the non-stationary variables. The rank of Π is reduced and equal to the number of cointegrating relationships, and we have $\Pi = \boldsymbol{\alpha\beta}'$. The parameters in $\boldsymbol{\beta}$ describe the long-run equilibria of the system and by analyzing them we are able to address the issue of long-run interdependence. The parameters in $\boldsymbol{\alpha}$ describe the short-run response of the system to disequilibria and by analyzing them we are able to attribute the status of (weak) exogeneity to those variables that do not react to disequilibria. If weak exogeneity applies to the $\mathbf{X}$ variables and there are is a unique cointegrating vectors, we can simplify our general reduced form model in the following Vector Error Correction specification:

$$
\begin{aligned}
\begin{pmatrix} \Delta LP_{Ger,t} \\ \Delta LP_{US,t} \end{pmatrix} &= \begin{pmatrix} \alpha_{11} & \alpha_{12} \\ \alpha_{21} & \alpha_{22} \end{pmatrix} \boldsymbol{\beta}' \begin{pmatrix} LP_{US,t-1} \\ LP_{Ger,t-1} \\ \mathbf{X}_{Ger,t-1} \\ \mathbf{X}_{Ger,t-1} \end{pmatrix} \\
&+ \Pi_1 \begin{pmatrix} \Delta \mathbf{X}_{Ger,t} \\ \Delta \mathbf{X}_{US,t} \end{pmatrix} \\
&+ \left(I + \begin{pmatrix} \psi_{11} & \psi_{12} \\ 0 & \psi_{22} \end{pmatrix} \begin{pmatrix} d_{1,t} & 0 \\ 0 & d_{2,t} \end{pmatrix} \right) \begin{pmatrix} u_{1,t} \\ u_{2,t} \end{pmatrix}
\end{aligned} \tag{3}
$$

Note that the variables contained in the $\mathbf{X}$ vectors are now validly considered as exogenous. Moreover, the specification of the matrix Ψ is designed to match the empirical evidence that there are some German dummy variables that are not significant in the equation for US share prices while the converse is not true. The methodology can be extended to more general specifications for the vector of dummies (see, for example, Favero and Giavazzi (2002)).

The simultaneous presence of dummies in both equations is not informative on the relative importance of contagion and interdependence. This issue cannot be resolved by estimating a reduced form and requires the specification of a structural model. The following structural model, consistent with the reduced form(3) , allows for both contagion and interdependence:

$$\begin{pmatrix} 1 & -\beta_{12} \\ 0 & 1 \end{pmatrix} \begin{pmatrix} \Delta LP_{Ger,t} \\ \Delta LP_{US,t} \end{pmatrix} = \begin{pmatrix} \gamma_{11} & \gamma_{12} \\ \gamma_{21} & \gamma_{22} \end{pmatrix} \boldsymbol{\beta}' \begin{pmatrix} LP_{US,t-1} \\ LP_{Ger,t-1} \\ \mathbf{X}_{Ger,t-1} \\ \mathbf{X}_{Ger,t-1} \end{pmatrix} + \tag{4}$$

$$+\Gamma_2 \begin{pmatrix} \Delta \mathbf{X}_{Ger,t} \\ \Delta \mathbf{X}_{US,t} \end{pmatrix} +$$

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

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

In(4) , we assumed triangularity in the simultaneous relationship between US and German stock prices. Such assumption characterizes our main identifying restrictions, based on the belief that US stock market is not simultaneously influenced by fluctuations in the German Stock markets, while we assume that the converse is not true. Note that in our empirical work we shall impose more restrictions on Γ_2, whose validity is testable as they are over-identifying restrictions. The presence of contagion is described by $a_{12} \neq 0$, because this indicates that modelling interdependence by explicitly allowing $\beta_{12} \neq 0$ is not enough to describe the way shocks are transmitted across countries in periods of turmoil.

The null hypothesis of no contagion can then be tested as an over-identifying restriction for our specification. In particular, the hypothesis of interdependence only and no contagion is parametrized as $\mathrm{H}_0 : a_{12} = 0$, which implies the following overidentifying restriction:

$$\psi_{12} = \beta_{12} a_{22}$$

Under H_0, turmoil in country 2 propagates to country 1 only through interdependence, as described by β_{12}.

As extensively discussed in Favero-Giavazzi(2002), our full-information approach to test for contagion can be compared to the limited information approach, based on the IV method proposed by Rigobon (2000) to estimate β_{12} and control for interdependence in order to detect contagion. Rigobon's methodology hinges on splitting the sample into high and low volatility periods. Based on this distinction, an instrument is constructed whose validity is warranted under the null of no contagion, then tests of validity of instruments are used as a test of contagion. The beauty of this approach depends on the fact that it does not require variables other than endogenous to implement the IV estimator. In fact instruments are constructed by taking transformation of the endogenous variables based on the presence of different regimes in volatility. Avoiding the estimation of a structural model of interdependence has the obvious benefit of imposing milder identifying restrictions than those necessary to implement our full-information procedure. The limited information approach has the advantage of identifying the system even when the traditional just-identifying restrictions are not valid. The main limit is that it is less powerful. 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 very few observations: asymptotic results along the dimension of the full sample size are still applicable while obviously none applies along the dimension of the high volatility sub-sample. This is not a problem when daily or intra-daily high-frequency data are considered. However, it might become a problem when the potential importance of the role of fundamentals calls for the use of lower frequency data. In such a situation our methodology, based on a full information estimation on the whole sample with the inclusion of dummies for the high-volatility periods, is still applicable. Obviously, when the size of the high-volatility and low-volatility sub-samples are sufficiently long and the just-identifying restrictions in the structural model are validly imposed, the limited and full information approaches should both produce consistent estimators and, therefore, the same results.

3 A statistical model for German and US share prices, earnings and long-term interest rates

Our statistical analysis of the relevance of contagion hinges on modelling both short-run and long-run interdependence between stock markets. We model long-run interdependence via cointegration and short-run interdependence via a small simultaneous structural model.

To investigate more closely the nature of the possible long-run equilibria, we consider the following VAR specification as our baseline statistical model:

$$\begin{pmatrix} LP_{US,t} \\ LE_{US,t} \\ R_{US,t} \\ LP_{Ger,t} \\ LE_{Ger,t} \\ R_{Ger,t} \end{pmatrix} = \mathbf{A}_0 + \sum_{i=1}^{4} \mathbf{A}_i \begin{pmatrix} LP_{US,t-i} \\ LE_{US,t-i} \\ R_{US,t-i} \\ LP_{Ger,t-i} \\ LE_{Ger,t-i} \\ R_{Ger,t-i} \end{pmatrix} + \begin{pmatrix} e_{1t} \\ e_{2t} \\ e_{3t} \\ e_{4t} \\ e_{5t} \\ e_{6t} \end{pmatrix}, \tag{5}$$

where LP_{US} and LP_{Ger} are the logs of the share price indexes, LE_{US} and LE_{Ger} the logs of I/B/E/S analysts forecasts of earnings, R_{US} and R_{Ger} the yields to maturity of ten-year benchmark bonds for US and Germany. Some discussion of our choice of variables and lag specification is in order.

Our choice of variables allows us to evaluate a number of different hypotheses recently adopted in the literature for the specification of long-run equilibria. Recent studies (Lander et al.(1997)), following the time honoured contribution by Graham and Dodd(1962), have chosen to construct an equilibrium for stock markets by concentrating on long-term interest rates and the earning-price ratio. Long-term interest rate feature a much stronger comovement with price-earning ratios than the short-term interest rate. Such evidence can be rationalized by considering that the long-term interest rates contain an element of risk premium which is absent in the short-term interest rates. Studies concentrating on the relationship between the short-term interest rates and dividend or earning yields have found empirical evidence of a sizeable and strongly persistent risk premium (see Blanchard (1983) and Wadhwani (1998)), which induces a rather weak long-run relationships among these variables.

Kasa(1992) has applied cointegration analysis to find a single common stochastic trend (and hence four cointegrating vectors) among the G5 stock market indexes. Serletis-King(1997) perform cointegration analysis in a framework similar to that of Kasa(1992) on ten EU stock markets. They measure the degree of convergence by applying time-varying parameter techniques to the vector of loadings measuring the short-run response of variables to disequilibria with respect to the cointegrating relationship(s). Our six variables VAR allows to test the validity of the alternative long-run equilibria proposed by these authors on our data sets Moreover we can also investigate the importance of long-run interdependence between stock markets, by evaluating the relative importance of domestic and international factors in the determination of long-run equilibria.

Turning to the data[3], some graphical evidence on a sample of monthly data over the period 1980-2002 is provided in Figures 1-3, where we report yields to maturity on 10-year German and US Treasury bonds along with the (log) of earning/price ratio for the US and German stock markets. .

Insert Figure 1-3 here

The time-series behaviour of the reported variables suggests that long-term interest rates and US price-earning ratio might share a common stochastic trend while the existence of such common trend is more dubious for the German case, in which deviations from trend tend to be more pronounced

[3]Our data-set comes from DATASTREAM. The stock price indexes are the Datastream all market indexes for US and Germany, the price earning ratios are from the same source and they are based on expected I/B/E/S analysts forecasts for end-of-year earnings. Finally, we considered yield-to-maturity for 10-year benchmark Treasury bonds. All data and an exact description of the Datastream stock market indexes are available from the website http://www.igier.uni-bocconi.it/personal/favero/homepage.htm

and more persistent. Very little evidence in favour of the hypothesis of common international stochastic trends in stock markets seem to emerge from our data.

Turning to the lag selection and the VAR specification, we have chosen the length of the distributed lags relying on the traditional likelihood based criteria. Note that, when we test for normality, heteroscedasticity and autocorrelation, strong evidence of non-normality emerges. Table 1 reports tests of the null hypothesis of residuals normality, both at the single equation and at the system level, proposed by Doornik and Hansen (1994).

Insert Table 1 here

The null of normality is rejected at the one per cent confidence level for all the equations in the system. As a consequence, also normality of the vector of VAR residuals is strongly rejected. These diagnostic tests, which are in general important to detect misspecification and to ensure validity of inference, take additional importance in our context. In fact, non-normality is possibly determined by the presence of outliers, capturing the occurrence of those periods of turmoils that are crucial for detecting contagion. In order to ensure congruency of our statistical model and be able to exploit the information contained in the periods of turmoil, we proceed to include a number of point dummies in our specification. More precisely, we use an automatic criterion and construct a point dummy (taking a value 1 for the relevant observation and zero everywhere else) for each estimated residual lying outside the ± 2.5 standard deviation interval[4].

As witnessed by the results reported in Table 2, the introduction of dummies largely solves the non-normality problems for all equations in our system, with the exception of equations for earnings, where non-normality is not

[4]The threshold has been chosenon the basis ofthe normality of residuals after the dummies have been included in the specification. Our results are robust to modification of such threshold in the range of 2-3 standard deviations

attributable to specific large outliers but to a consistent number of outliers of moderate dimension.

Insert Table 2 here

After controlling for outliers, we consider the following VAR as the baseline statistical model for our investigation:

$$\begin{pmatrix} LP_{US,t} \\ LE_{US,t} \\ R_{US,t} \\ LP_{Ger,t} \\ LE_{Ger,t} \\ R_{Ger,t} \end{pmatrix} = \mathbf{A}_0 + \sum_{i=1}^{4} \mathbf{A}_i \begin{pmatrix} LP_{US,t-i} \\ LE_{US,t-i} \\ R_{US,t-i} \\ LP_{Ger,t-i} \\ LE_{Ger,t-i} \\ R_{Ger,t-i} \end{pmatrix} + \mathbf{B} * \mathbf{DUM} + \begin{pmatrix} e_{1t} \\ e_{2t} \\ e_{3t} \\ e_{4t} \\ e_{5t} \\ e_{6t} \end{pmatrix}, \quad (6)$$

where **DUM** is a vector of thirty-three dummies, taking value of one when the outlier occurs and zero anywhere else.

Endowed with model (6), we address the first issue of our interest: long-run equilibrium and interdependence between US and German stock markets.

Re-parameterize (6) as follows:

$$
\begin{pmatrix} \Delta LP_{US,t} \\ \Delta LE_{US,t} \\ \Delta R_{US,t} \\ \Delta LP_{Ger,t} \\ \Delta LE_{Ger,t} \\ \Delta R_{Ger,t} \end{pmatrix} = \mathbf{A}_0 + \sum_{i=1}^{3} \Pi_i \begin{pmatrix} \Delta LP_{US,t-i} \\ \Delta LE_{US,t-i} \\ \Delta R_{US,t-i} \\ \Delta LP_{Ger,t-i} \\ \Delta LE_{Ger,t-i} \\ \Delta R_{Ger,t-i} \end{pmatrix} + \Pi \begin{pmatrix} LP_{US,t-1} \\ LE_{US,t-1} \\ R_{US,t-1} \\ LP_{Ger,t-1} \\ LE_{Ger,t-1} \\ R_{Ger,t-1} \end{pmatrix} +
$$

$$
+ \mathbf{B} * \mathbf{DUM} + \begin{pmatrix} e_{1t} \\ e_{2t} \\ e_{3t} \\ e_{4t} \\ e_{5t} \\ e_{6t} \end{pmatrix}, \tag{7}
$$

$$
\Pi_i = -\left(I - \sum_{j=1}^{i} \mathbf{A}_j \right),
$$

$$
\Pi = -\left(I - \sum_{i=1}^{3} \mathbf{A}_i \right),
$$

where the matrix Π describes the long-run properties of our system. In particular, the rank of Π determines the number of cointegrating vectors. Whenever the rank of Π is reduced, the following decomposition applies $\Pi = \boldsymbol{\alpha}\boldsymbol{\beta}'$, where the matrix $\boldsymbol{\beta}$ contains the parameters in the cointegrating vector and the matrix $\boldsymbol{\alpha}$ contains the loadings describing the adjustment of each variable to disequilibria with respect of the long-run equilibrium of the system. We analyze the rank of Π and its decomposition by using the statistical framework proposed by Johansen(1995). Results are reported in Table 3.

Insert Table 3 here

The Table reports the sequence of estimated eigenvalues of the long-run matrix along with the test for the rank of Π based on the trace-statistic and the maximum eigenvalue statistics, which points toward the existence of a unique cointegrating vector.[5]

Having fixed the rank of Π to one, we test alternative hypotheses on the specification of the long-run relationship. We consider four alternative hypotheses on the specification of the long-run equilibrium. H_1 postulates a long-run relation between the log of US price-earning and yields to maturity of US and German long-term bonds. Under H_2 a long-run relation exists between the log of German price-earning and yields to maturity of US and German long-term bonds. H_3 claims a long-run relation links the log of German price-earning ratio to the log of US price-earning ratio, and finally H_4 postulates a long-run relation between the log of German stock price and the log of US stock price. The first two hypotheses reflect a generalized version of the long-run solution based on Graham and Dodd and adopted by Lander et al.(1997): H_1 applies it to the US, while H_2 applies it to Germany. H_3 and H_4 allow explicitly for interdependence among US and German stock markets, following the specification of the cointegrating relations chosen by Kasa (1992) and Serletis-King(1997).

Only hypothesis H_1,implying a long-run relationship between the (log of) US price-earning ratio and the US and German long-term interest rates is not statistically rejected. Moreover, the loadings associated to the cointegrating vector show that only the US and German stock prices significantly react to disequilibrium. Hence earnings and long-term interest rates can be considered as weakly exogenous for the estimation of the parameters of interest when estimating models for share prices. Our cointegrating relationship is directly comparable with that obtained by Lander et al.(1997). In fact, we obtain very similar results except that the long-term interest rates relevant

[5]We report both trace and maximum eigevalue statistics although there is evidence that the trace statistic is preferable as the sequence of trace tests lead to a consistent test procedure, while no such result is available for the max eigenvalue statisatics see (see Doornik and Hendry(2001)). The presence of dummies makes the traditional critical values not appropriate although the different in magnitude of in the sequence of eigenvalue suggests robustness of the evidence in favour of the existence of a unique cointegrating vector.

to our cointegrating vector are some weighted average of the US and German long-term rates. Figure 1 may help the interpretation of such results: over the second part of our sample there is virtually no difference between the two long-term rates, while in the first part of the sample the US long-term rates fluctuate remarkably more than the German ones. Price/earning ratios differ from the nominal long term interest rates in that they are real variables and hence they are less affected by inflation[6]. Cointegration between price/earnings and long-term nominal rates implies stationarity of inflation. Current analysis of U.S. monetary policy generally acknowledges that 1979 marks the beginning of a new policy regime characterized by a strong anti-inflationary stance which allowed a mean reverting relation between effective inflation and the target chosen by the monetary policy authorities[7]. Despite the change in the monetary policy regime, some episodes of "inflation scares" hit the US bond market at the beginning of the new monetary regime. As these episodes remained local, some weighted average of the US and German rates is not so dramatically affected by the temporary jumps in expected inflation and keeps a better balance with the price/earning ratio.

We conclude this section by reporting in Figure 4 the deviation of US share prices from their equilibrium value.

Insert Figure 4 here

The Figure shows twenty episode of mean reversion over twenty years. The Figure also suggests that the US market was heavily overvalued at the beginning in 1982 and at the end of year 2000, while at the end of September

[6]By inflation here we mean average ten-year inflation. In fact our price-earning ratios, being defined with reference to expected earnings, are indeed affected by short-term, one-period ahead, inflation.

[7]Empirical investigations of the Fed's reaction function confirm this discontinuity. See the widely cited work of Clarida, Gali and Gertler (2000). Cogley and Sargent (2002) also relate the conquest of U.S. inflation to a different behaviour of the monetary policy authority under the Volcker and Greenspan tenures.

2002 share prices fluctuated at thirty per cent discount with respect to their equilibrium value.

4 Measuring short-run interdependence and contagion

To describe short-run interdependence and assess contagion we need a structural model. We build it starting from simplifying the baseline statistical model into a bivariate Vector Error Correction model for US and German share prices, where, on the basis of the statistical evidence on the loadings reported in Table 3, long-term interest rates and earnings are taken as weakly exogenous :

$$\begin{pmatrix} \Delta LP_{US,t} \\ \Delta LP_{Ger,t} \end{pmatrix} = \mathbf{B}_0 + \mathbf{B}_1 \begin{pmatrix} \Delta LP_{US,t-1} \\ \Delta LP_{Ger,t-1} \end{pmatrix} + \sum_{i=0}^{1} \mathbf{C}_i \begin{pmatrix} \Delta R_{US,t-i} \\ \Delta R_{Ger,t-i} \\ \Delta LE_{US,t-i} \\ \Delta LE_{Ger,t-i} \end{pmatrix} + \tag{8}$$

$$+ \mathbf{D} \left(LP_{US,t-1} - LP^{*}_{US,t-1} \right) + \mathbf{F} \left(dum \right) + \begin{pmatrix} u_{1t} \\ u_{2t} \end{pmatrix},$$

$$LP^{*}_{US,t-1} = LE_{US,t-1} - 0.10 R_{US,t-1} - 0.111 R_{Ger,t-1} + 4.4,$$

The vector of dummies dum is a sub-vector of the one containing thirty-three dummies used for the general system used for the cointegration analysis. The dynamics of the system considering earnings and long-term rates as exogenous is much shorter as first order dynamics is now selected by the optimal lag selection criteria. There are twelve outliers, among which nine are common to both equations and three are specific to the equation for the German share price. The common dummies correspond to episodes of US stock market turmoil. In 1987:10, 1998:08, 2001:02, 2001:03, 2001:09 2002:7 and 2002:09, we observed downward movements respectively of twenty-four, twelve, eleven, nine, seven and a half and nine per cent, while in 1987:01 and

1998:10 equity prices jumped up by thirteen and twelve per cent. Country specific movements in German equity prices are accounted for by the dummies respectively of 1990:09 and 1997:08, and of 1999:12, when the market fell by nineteen and eleven and rose by thirteen per cent. The diagnostic tests[8] reported in Table 4 show that the null of absence of residuals correlation, homoscedasticity and normality cannot be rejected for (8) .

Insert Table 4 here

On the basis of this reduced form, we proceed to estimate two structural models. As discussed in section 2, we consider a more general one allowing for both short-run interdependence and contagion, and a more restrictive model consistent with the hypothesis of "only interdependence, no contagion"[9]. Both structural models impose some testable over-identifying restrictions on our reduced form and we can therefore use the outcome of the tests to discriminate between the cases of interest. The estimated structural models are reported in Table 5. Both models show that the fluctuations in local fundamentals, such as earnings and the long-term interest rates, determine fluctuations in share prices. The US market also react very significantly to deviation of US share prices from their long-run equilibrium. Such variables also affect the fluctuations in German prices although the effect is quantitatively smaller and just marginally statistically significant. Model 1 in Table 5 is consistent with the hypothesis of the existence of contagion between the US and European stock markets. In fact, in the case of interdependence only, when a simultaneous feedback is allowed from US to European stock markets, the dummies capturing turmoil periods in the US market should not enter significantly the equation for German stock prices. We observe

[8]All the tests have performed at system level using PC-FIML, for a detailed description see Doornik-Hendry(1997)

[9]Note that our discussion in section 2 introduces multiplicative dummies on the residuals, this is equivalent to the introduction of shift dummies in the structural model. In fact, with point dummies multiplicative effect are observationally equivalent to shift effects.

that not only do such dummies enter significantly, but their inclusion also renders the simultaneous feedback between German and US stock markets not significantly different from zero. Importantly, the model is supported by the data in that the tests for the validity of the ten over-identifying restrictions imposed by Model 1 on the general reduced form (8) does not lead to the rejection of the null hypothesis of interest.

The results from the estimation of the structural model implicit in the hypothesis of "no contagion, only interdependence"are reported in the same Table under the label of Model 2. The validity of over-identifying restrictions is now rejected. As we have nine dummies for the US stock market, our test for the null of no-contagion is distributed as a χ^2_{29}, with nine more degrees of freedom than the statistic used to test the validity of Model 1. Interestingly, as a consequence of the omission of dummies, the significance of the simultaneous feedback increases drastically and might mislead the inference whenever Model 2 is estimated without reference to the general model (8).

Insert Tables 5

To allow comparison of our results with the IV based approach we have created an instrument w_t,which is equal to $-\Delta LP_{US,t}/261$ for all observations in our sample except for 1987:01 ,1987:10, 1998:08, 1998:10, 2001:02, 2001:03, 2001:09 and 2002:09, where it takes value $\Delta LP_{US,t}/9$. We report in Table 6 the results of the regression showing the validity of w_t as an instrument for $\Delta LP_{US,t}$ along with the augmented regression for the German share prices which allows to implement the Hausman- type test for the validity of instrument, suggested by Rigobon as a test for contagion.

Insert Table 6 here

As the coefficient on $\hat{u}_t$ is significantly different from zero, the null of no-contagion is rejected and our results are confirmed by the implementation of the IV procedure.

5 Conclusions

In this paper we have proposed a methodology to disentangle interdependence from contagion in co-movements between stock markets and applied it to the case of the German and US stock markets. We assessed the relative importance of contagion and interdependence within the framework of an explicit structural model, using cointegration analysis to separate long-run equilibria from short term dynamics. We constructed our long-run equilibria by tsting different possible specification and favouring the hypothesis of cointegration between the (log of) US earning-price ratio and long-term interest rates. Within such framework, we found that the hypothesis of no long-run interdependence between the two markets cannot be rejected. We then used our Vector Error Correction Model as a baseline reduced form and constructed a structural model to assess the relative importance of interdependence and contagion in determining the short-run dynamics of the two markets. Our structural model shows that the effect of fluctuations of US stock market on the German stock market is captured by a non-linear specification. Normal fluctuations in the US stock market have virtually no effect on the German market, while such effect becomes sizeable and significant for abnormal fluctuations. Such non-linearity is clearly consistent with the relevance of contagion, in that it amounts to a modification of short run interdependence in periods of turmoil. Our results are proven to be consistent with those obtained by applying the Instrumental Variable methodology proposed by Rigobon (1999). We believe that the à-touts of our approach originate from the specification of a cointegrated structural model along the lines of the LSE strategy. In particular, this minimizes the risks of mis-specification, allows to distinguish between short-run and long-run interdependence and naturally leads to the identification of turmoil episodes crucial to assess the relevance of contagion.

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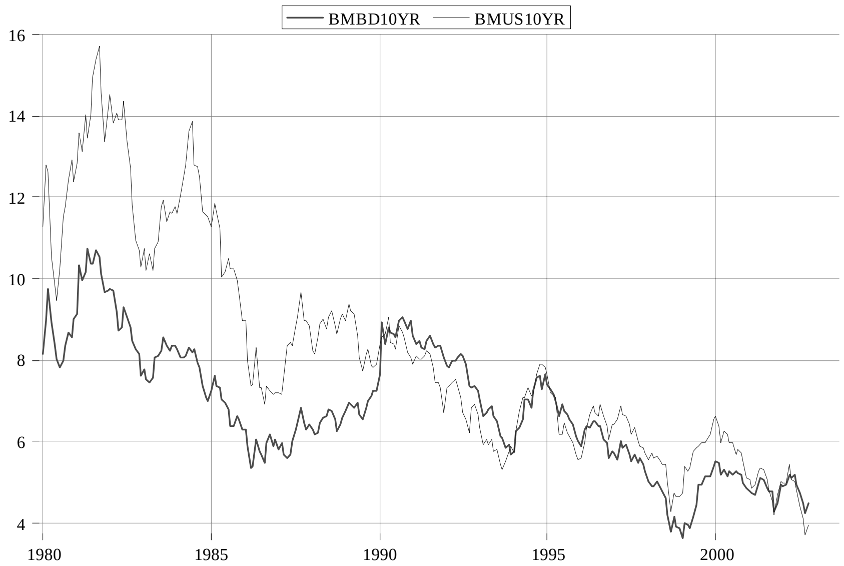

Figure 1: US (BMUS10YR) and German (BMBD10YR) long-term interest rates. Source: Datastream.

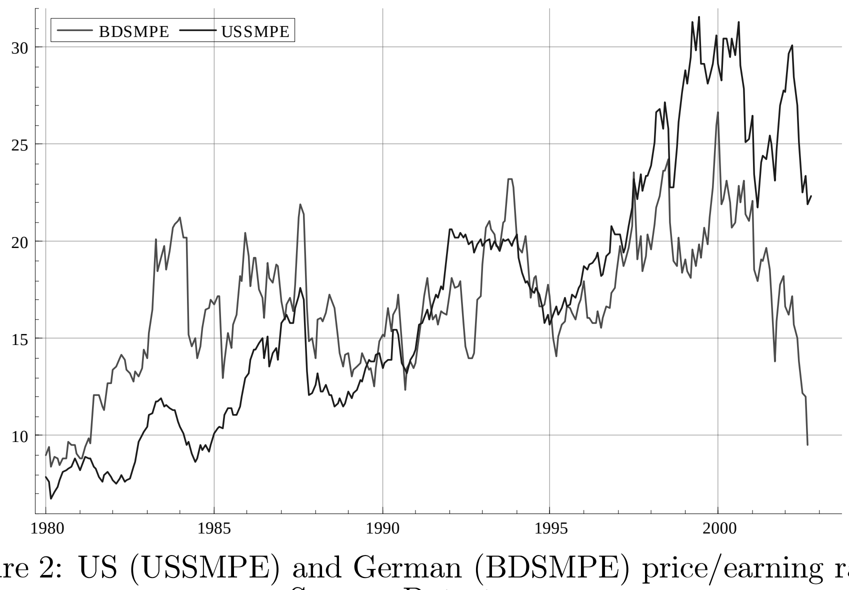

Figure 2: US (USSMPE) and German (BDSMPE) price/earning ratios. Source: Datastream.

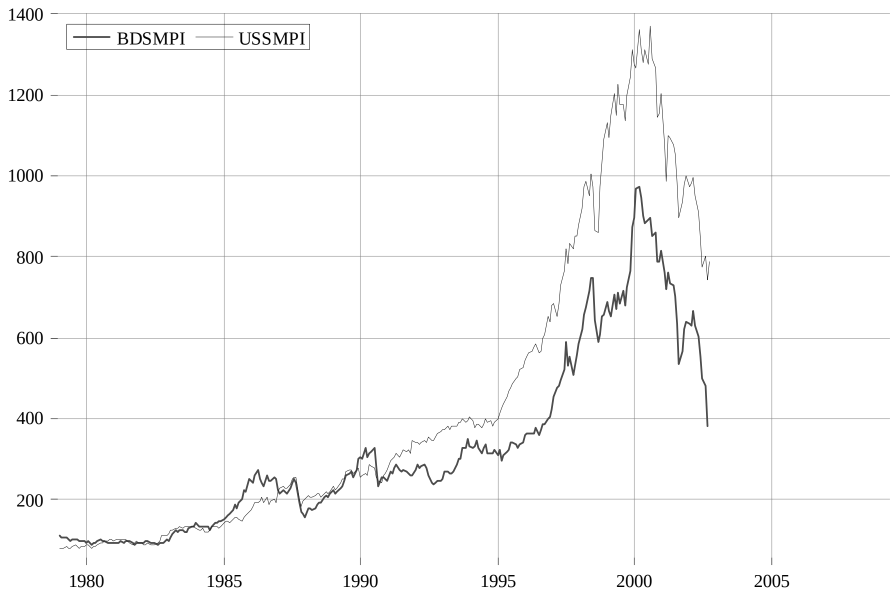

Figure 3: US(USSMPI) and German(BDSMPI) Datastream all market share price indexes

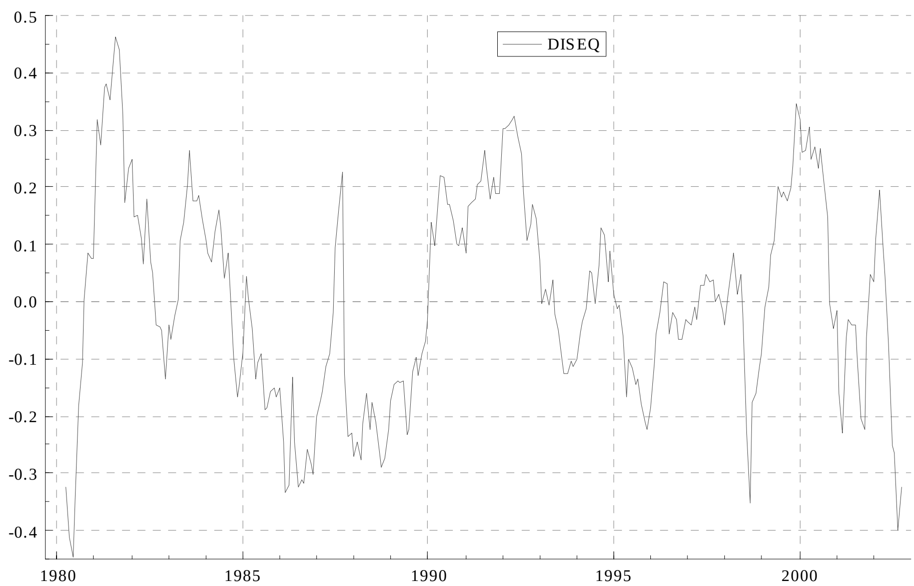

Figure 4: Deviation from long-run equilibrium of US share prices (0.x indicate a 10*x per cent deviation)

Table 1: Testing normality of residuals in the VAR system		
Single equation	without dummies	with dummies
$LP_{US,t}$	35.951**	7.27*
$LE_{US,t}$	67.194**	21.22**
$R_{US,t}$	1.282	4.71
$LP_{Ger,t}$	34.276**	2.08
$LE_{Ger,t}$	83.044**	73.24**
$R_{Ger,t}$	34.502**	2.33

The estimated model is $\mathbf{Y}_t = \mathbf{A}_0 + \sum_{i=1}^{4} \mathbf{A}_i \mathbf{Y}_{t-i} + \mathbf{u}_t$, with

$Y_t = \left[LP_{US,t}, LE_{US,t}, R_{US,t}, LP_{Ger,t}, LE_{Ger,t}, R_{Ger,t}\right]'$,

dummies are introduced to eliminate outliers, defined as observed residuals
with an absolute value larger larger than 2.5 time their standard deviation.
The test statistics reported are based on Hansen-Doornik(1994) and distributed as a χ^2_2
* and ** indicate rejection respectively at 5 and 1 per cent significance level.

TABLE 2: Testing the number of cointegrating vectors			
Variable	$\mathrm{H}_0 : rank\ \Pi = p$	Trace	Max Eig.
0.184	$p = 0$	110.3**	54.6**
0.066	$p \leq 1$	55.7	18.54
0.054	$p \leq 2$	37.13	14.98
0.045	$p \leq 3$	22.15	12.57
0.023	$p \leq 4$	9.58	6.43
0.01	$p \leq 5$	3.15	3.15

Eigenvalue column reports the estimated eigenvalues of Π. Trace and Max Eig. columns reports the values of the trace and. maximum eigenvalue statistics for the null $rank\ \Pi = p$, i.e. there are at most p cointegrating vectors

TABLE 3: Testing hypothesis on the long-run equilibrium					
	H_1		H_2	H_3	H_4
Variable	β	α	β	β	β
$LP_{US,t}$	1	-0.05 (0.015)	0	1	1
$LE_{US,t}$	-1	0.005 (0.006)	0	-1	-1
$R_{US,t}$	0.10 (0.016)	-0.13 (0.15)	-0.013 (0.032)	0	0
$LP_{Ger,t}$	0	-0.056 (0.018)	1	-6.33 (2.68)	0
$LE_{Ger,t}$	0	-0.039 (0.15)	-1	6.33 (2.68)	0
$R_{Ger,t}$	0.10 (0.03)	-0.17 (0.09)	-0.10 (0.058)	0	0
Testing restrictions	$\chi^2_3 = 10.78$		$\chi^2_3 = 30.92^{**}$	$\chi^2_4 = 34.86^{**}$	$\chi^2_5 = 36.57^{**}$

The Table reports test of the validity of restrictions on the unique cointegrating vectors. We consider four hypothesis

H_1 postulates a long-run relation between the log of US price-earning and yields to maturity of US and German long-term bonds.

H_2 postulates a long-run relation between the log of German price-earning and yields to maturity of US and German long-term bonds.

H_3 postulates a long-run relation between the log of German price-earning ratio and the log of US price-earning ratio.

H_4 postulates a long-run relation between the log of German stock price and the log of US stock price

For each hypothesis we report estimated parameters in the cointegrating vector, with the associated standard error, and, whenever the validity of the over-identifying restrictions is not rejected, the estimating loading of the cointegrating vectors in equations associated to each variable in the VAR.

Table 4: Testing congruency of the bivariate VECM			
	Normality	Autocorrelation 1-7	Heteroscedasticity
$\Delta LP_{US,t}$	$\chi^2_2 = 11.1(0.01)$	$F_{7,237} = 0.8(0.63)$	$F_{46,204} = 1(0.43)$
$\Delta LP_{Ger,t}$	$\chi^2_2 = 3.6(0.17)$	$F_{7,237} = 0.6(0.76)$	$F_{46,204} = 1.3(0.11)$
System	$\chi^2_4 = 4.9(0.30)$	$F_{28,472} = 0.7(0.87)$	$F_{138,606} = 1.1(0.22)$

The table reports tests for Normality, Heteroscedasticity and Autocorrelation of residuals for the following model:

$$\begin{pmatrix} \Delta LP_{US,t} \\ \Delta LP_{Ger,t} \end{pmatrix} = \mathbf{B}_0 + \mathbf{B}_1 \begin{pmatrix} \Delta LP_{US,t-1} \\ \Delta LP_{Ger,t-1} \end{pmatrix} + \sum_{i=0}^{1} \mathbf{C}_i \begin{pmatrix} \Delta R_{US,t-i} \\ \Delta R_{Ger,t-i} \\ \Delta LE_{US,t-i} \\ \Delta LE_{Ger,t-i} \end{pmatrix} +$$

$$+ \mathbf{D} \left(LP_{US,t-1} - LP^*_{US,t-1} \right) + \mathbf{F} \left(\mathbf{dum} \right) + \begin{pmatrix} u_{1t} \\ u_{2t} \end{pmatrix},$$

$$LP^*_{US,t-1} = LE_{US,t-1} - 0.10 R_{US,t-1} - 0.111 R_{Ger,t-1} + 4.4,$$

where the vector dum contains dummies for periods 87:1, 87:10, 90:9, 97:8, 98:8, 98:10, 99:12 01:2, 01:3, 01:9, 02:9.

Rows two and three of the Table report the relevant statistics for each equation with p-values in parentheses, while row four does the same for the system. Normality test, based on Hansen and Doornik (1994), is rejected for the first equation but neither for thesecond nor for the system. LM test for autocorrelation up to the seventh order andWhite (1980) test for heteroscedasticity of residuals are rejected both for singleequations and for the entire system

Table 5: Structural models for US and European stock prices				
	Model 1		Model 2	
	$\Delta LP_{US,t}$	$\Delta LP_{Ger,t}$	$\Delta LP_{US,t}$	$\Delta LP_{Ger,t}$
Constant	$\mathbf{0.009}$ (0.002)	$\mathbf{0.008}$ (0.004)	$\mathbf{0.006}$ (0.002)	-0.0007 (0.003)
$\Delta LP_{US,t}$		-0.14 (0.271)		$\mathbf{0.90}$ (0.109)
ECM_{t-1}	$\mathbf{-0.053}$ (0.012)	$\mathbf{-0.036}$ (0.021)	$\mathbf{-0.055}$ (0.012)	0.026 (0.017)
$\Delta R_{US,t}$	$\mathbf{-0.032}$ (0.005)		$\mathbf{-0.024}$ (0.005)	
$\Delta R_{Ger,t}$		$\mathbf{-0.048}$ (0.011)		-0.019 (0.010)
$\Delta LE_{US,t}$	$\mathbf{0.19}$ (0.094)		0.18 (0.111)	
$DUM8701$	$\mathbf{0.108}$ (0.035)	$\mathbf{-0.125}$ (0.056)	0.061 (0.034)	
$DUM8710$	$\mathbf{-0.259}$ (0.035)	$\mathbf{-0.29}$ (0.083)	$\mathbf{-0.264}$ (0.034)	
$DUM9009$		$\mathbf{-0.168}$ (0.038)		$\mathbf{-0.176}$ (0.044)
$DUM9708$		$\mathbf{-0.083}$ (0.039)		-0.082 (0.044)
$DUM9808$	$\mathbf{-0.141}$ (0.035)	$\mathbf{-0.19}$ (0.056)	$\mathbf{-0.150}$ (0.034)	
$DUM9810$	$\mathbf{0.104}$ (0.035)	0.047 (0.053)	$\mathbf{0.088}$ (0.034)	
$DUM9912$		$\mathbf{0.104}$ (0.039)		$\mathbf{0.112}$ (0.044)
$DUM0102$	$\mathbf{-0.126}$ (0.035)	-0.095 (0.054)	$\mathbf{-0.120}$ (0.034)	
$DUM0103$	$\mathbf{-0.101}$ (0.035)	-0.09 (0.056)	$\mathbf{-0.098}$ (0.034)	
$DUM0109$	$\mathbf{-0.112}$ (0.035)	$\mathbf{-0.19}$ (0.056)	$\mathbf{-0.130}$ (0.034)	
$DUM0207$	$\mathbf{-0.11}$ (0.033)	$\mathbf{-0.14}$ (0.056)	$\mathbf{-0.141}$ (0.034)	
$DUM0209$	$\mathbf{-0.107}$ (0.035)	$\mathbf{-0.28}$ (0.056)	$\mathbf{-0.141}$ (0.034)	
LR test	$\chi^2_{19} = 20.73(0.37)$		$\chi^2_{28} = 101.94(0.000)$	

Model 1 reflects the hypothesis of "interdependence and contagion" among US and German stock markets. Model 2 reflects the hypothesis of no contagion. The LR test is a statistic for the valifity of the over-identifying restrictions imposed by each model on the Vector Error Correction Reduced form reported in Table 4

TABLE 6: Testing contagion by the limited information approach		
	$\Delta LP_{US,t}$	$\Delta LP_{Ger,t}$
w_t	**0.010** (0.002)	
Constant	**7.884** (0.877)	-0.007 (0.005)
ECM_{t-1}		-0.030 (0.025)
$\Delta LE_{Ger,t}$		0.281 (0.463)
ΔR_{Ger}		$-$**0.056** (0.019)
$\Delta LP_{US,t}$		**1.137** (0.181)
$\widehat{u}_t$		$-$**1.488** (0.469)
$DUM9009$		$-$**0.195** (0.057)
$DUM9708$		-0.154 (0.081)
$DUM9912$		**0.164** (0.061)

Column 2 reports OLS coefficients from the regression of $\Delta LP_{US,t}$ on w_t, the instrument suggested by Rigobon, which equals $\frac{\Delta LP_{US,t}}{9}$ for the observations corresponding to US-specific dummies and $-\frac{\Delta LP_{US,t}}{264}$ elsewhere. Colum 3 reports coefficients estimated by IV. . Instruments are: $\Delta LE_{US,t}, \Delta R_{US,t}, \Delta LE_{Ger,t}, \Delta R_{Ger,t}, ECM_{t-1}$, DUM8701, DUM8710, DUM9009, DUM9708, DUM9808, DUM9810, DUM9912, DUM0102 DUM0103, DUM0109, DUM0207, DUM0209, u_t are estimated residuals from the regression in column2. Standard errors are reported in parentheses and significant coefficients in bold.