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Bank Credit in the EU: Long-run Independence or Integration with Germany?

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Abstract
This study examines the long-run role of Germany in the determination of bank credit in the EU. For a sample of nine EU members, the Johansen procedure is employed to test for common trends and bivariate cointegrating relationships with Germany. Long-run causality is investigated through tests which are based on the speed of adjustment towards long-run equilibrium. Long-run equilibrium relationships with Germany are identified in most cases. However, the results lend limited support to the German dominance hypothesis since credit provision appears to be more interactive in nature.

1. Introduction

A number of studies (see, inter alia, Koedijk and Kool (1992), Hafer and Kutan (1994), Katsimbris and Miller (1995)) have found that financial integration in the European Union (EU) has increased over the Exchange Rate Mechanism (ERM) period and that Germany, while the key player in EU monetary policy, is itself influenced by the stance of non-German monetary policy. While most studies of financial integration have focused on interest rates, monetary bases or monetary aggregates, little attention has been paid to the provision of bank credit which is a key consideration in an assessment of monetary conditions and the provision of finance. Using cointegration analysis and causality tests, this study examines the provision of bank credit for a sample of EU members and, in particular, the role played by Germany. Three questions are of direct interest. First, is there evidence that the provision of bank credit in EU countries bears a long-run relationship with that of Germany? Second, does the sensitivity to Germany depend on the degree of commitment to ERM membership? Third, is there evidence that the German Dominance Hypothesis (GDH) holds in this context, i.e. where the provision of credit in Germany drives the provision of credit elsewhere in the ERM? The extent to which the existing findings concerning financial integration may apply to the provision of bank credit is an open issue. Given that interest rates are an important determinant of credit market activity and the ERM has facilitated closer interest rate linkages in the EMS, it could be argued that increased integration in EU credit markets is inevitable. However, a number of considerations may lead one to expect that integration is limited. First, although interest rates have become more closely related, sensitivities to given interest rate changes may differ across EU members. Second, other key determinants of credit market activity include economic activity and attitudes to risk across the EU. With regard to the former case, Serletis and Krichel (1992) find limited evidence that EU output growth rates have converged. Finally, it can be argued that cross-border banking activity is relatively limited where domestic borrowers have a preference for indigenous institutions (Holmes, 1992).

Such a study is important for a number of reasons. First, there is an additional dimension to examining financial integration in the EU in the context of national credit markets. For the reasons mentioned above, there may be a strong case for arguing that the experiences in terms of credit markets may be at variance with the experiences noted under the more general heading of financial integration. This study offers a more defined examination in terms of one important aspect of financial integration. Second, the extent to which national credit markets are integrated has implications for the ability of national governments to regulate the provision of credit. Third, increased debt levels within member countries may have implications for relative sensitivities of economic activity to given interest rate fluctuations. Given that interest rates are a key instrument in maintaining designated exchange rate bands, the closer integration of credit provision may indicate similarities in the macroeconomic costs of interest rate adjustments. Fourth, such an investigation enables us to examine the provision of credit in an international context which, given an increased commitment to maintaining exchange rate bands, the easing of capital controls and restrictions on the movement of goods, is now more appropriate than the standard closed economy analyses.[2]

2. Modelling credit provision in the EU

Assume that bank lending in each country is determined by the following long-run relationships.

$$C_t^{D,i} = a(r_t^i,\ \lambda_t^i) \qquad (1)$$

$$C_t^{S,i} = b(r_t^i,\ G_t^i,\ D_t^i,\ \Phi_t^i) \qquad (2)$$

$$C_t^{D,*} = g(r_t^*,\ \lambda_t^*) \qquad (3)$$

$$C_t^{S,*} = h(r_t^*,\ G_t^*,\ D_t^*,\ \Phi_t^*) \qquad (4)$$

$$b_1,\ h_1,\ \geq 0,\ a_1,\ g_1 \leq 0$$

where i refers to country i and * refers to the foreign or base country, i.e. Germany. C is the log of total real bank credit. Equations (1) and (3) are demand functions for bank credit while (2) and (4) are supply functions. r is the interest rate attached to bank lending, λ incorporates a set of arguments which influence credit demand. These include an indicator of economic activity, expected inflation, private sector wealth, non-price terms which are attached to credit and so on. The supply of credit is positively influenced by interest rates, by G which is government restrictions on the supply of bank credit, by D which is the level of deposit taking activity and Φ which captures a range of other factors which influence the supply such as risk considerations.

The extent of integration across EU credit markets largely depends on the extent to which the explanatory variables included in (1)-(4) are related. On the supply-side, the moves towards a single market have encouraged greater cross-border competition among financial institutions thereby providing the opportunity for domestic quantity-based credit controls to be circumvented (see, for example, Llewellyn and Holmes (1991)). The extent to which interest rates can be independently used to regulate monetary growth has, for many members, been moderated by the commitment to the ERM bands of exchange rate fluctuations. In terms of demand side factors, studies by Koedijk and Kool (1992) and Hafer and Kutan (1994) have confirmed the presence of close interrelationships for interest rates across EU members. Under the composite null hypothesis of the integration of the explanatory variables in (1)-(4) and market clearing across EU credit markets, it can be shown that a positive long-run relationship exists between C^i and C^*:

$$C_t^i = \zeta_0 + \zeta_1 C_t^* \tag{5}$$

where $\zeta_1 \geq 0$. With no integration of demand- or supply-side factors then $\zeta_1 = 0$. An alternative scenario is where capital controls are effective and bank credit is supply constrained through the use of quantity-based controls. This would facilitate a greater independence for domestic credit markets and therefore would be likely to give rise to a relatively small value for ζ_1. In the extreme case of zero capital mobility and freely floating exchange rates, we might expect the domestic credit market to be insulated from external events and so $\zeta_1 = 0$.

3. Data and Estimation

Equation (5) forms the basis of estimation. Data on real bank credit for France, Germany, Greece, Ireland, Italy, the Netherlands, Portugal, Spain and the UK are taken from the IFS database (line 22d) and then deflated by national price indices[3] and finally transformed into natural log form. The series are quarterly and cover the ERM period, 1979Q1-93Q4 thereby providing 60 observations. Figures 1 and 2 provide plots for each country.[4] Each figure comprises Germany and four of the remaining sample. It can be seen that six of the series, particularly Italy and the Netherlands, are on an upward trend during the period of study while Greece, Ireland and Spain appear to be fairly static.

Estimation is based on the following stages. First, the full sample of EU members are included and estimated in a system without defining which is the base country. Using the Johansen cointegrating procedure, this approach enables us to identify the number of common shared trends in the data. A necessary condition for complete integration across EU members is evidence of a single shared trend. The second approach to estimating (5) is to look for bilateral relationships between domestic credit and German credit with Germany defined as the base country. Given any possible interdependencies in the data series, the Johansen cointegration methodology is appropriate. Finally, the GDH is investigated through weak exogeneity tests on the terms which represent the speed of adjustment towards long-run equilibrium.

More specifically, suppose $\mathbf{z}_t$ is a vector of n potentially endogenous variables, we may model $\mathbf{z}_t$ as an unrestricted VAR involving up to k lags of $\mathbf{z}$.

$$\mathbf{z}_t = \mathbf{A}_1\mathbf{z}_{t-1} + \mathbf{A}_2\mathbf{z}_{t-2} + \ldots + \mathbf{A}_k\mathbf{z}_{t-k} + \mathbf{u}_t \tag{6}$$

where $\mathbf{u}_t \sim NI(0,\Sigma)$. We can reformulate (6) into the following vector error correction model

$$\Delta\mathbf{z}_t = \Gamma_1\Delta\mathbf{z}_{t-1} + \ldots + \Gamma_2\Delta\mathbf{z}_{t-k+1} + \Pi\mathbf{z}_{t-k} + \mathbf{u}_t \tag{7}$$

where $\Gamma_i = -(\mathbf{I}-\mathbf{A}_1-\ldots-\mathbf{A}_i)$, ($i$=1, 2,.., k-1) and $\Pi = -(\mathbf{I}-\mathbf{A}_1-\ldots-\mathbf{A}_k)$. We may write

$$\Pi = \alpha\beta' \tag{8}$$

where α represents the speed of adjustment to equilibrium, while β is a matrix of long-run coefficients such that βz_{t-k} represents up to (n-1) cointegrating relationships in the multivariate model which ensure that the elements in $\mathbf{z}_t$ converge to their long-run steady state solutions. We may test for causality by examining the nature of α. Suppose $\mathbf{z}'_t = [y_{1t}, y_{2t}, y_{3t}]$, the number of cointegrating vectors, r=2 and k=1. We may

Figure 1. Bank Credit in the EU, 1979-93.

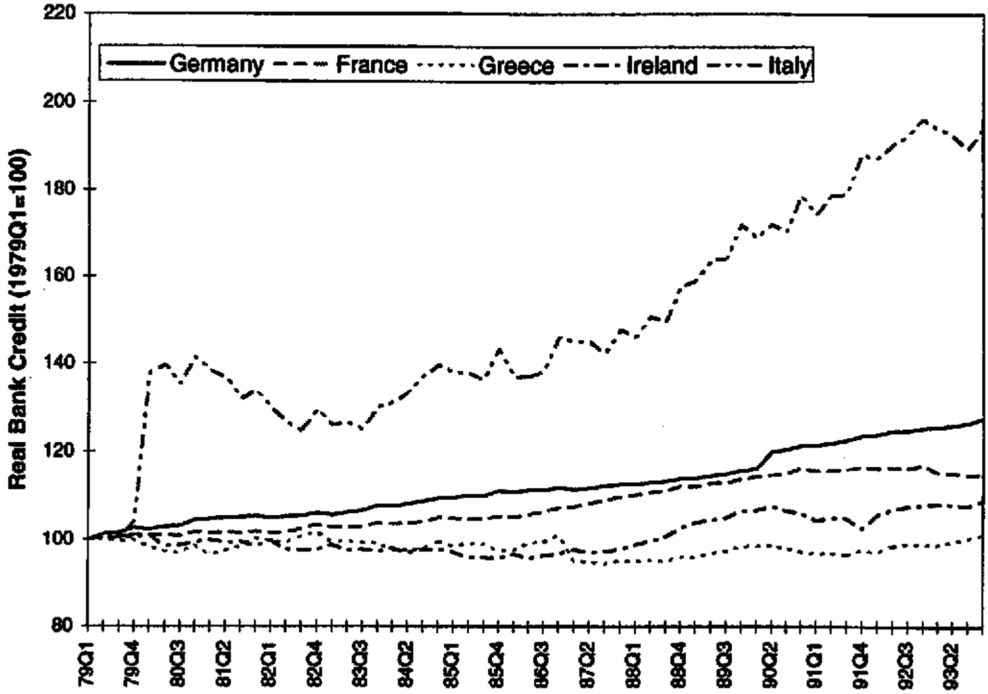
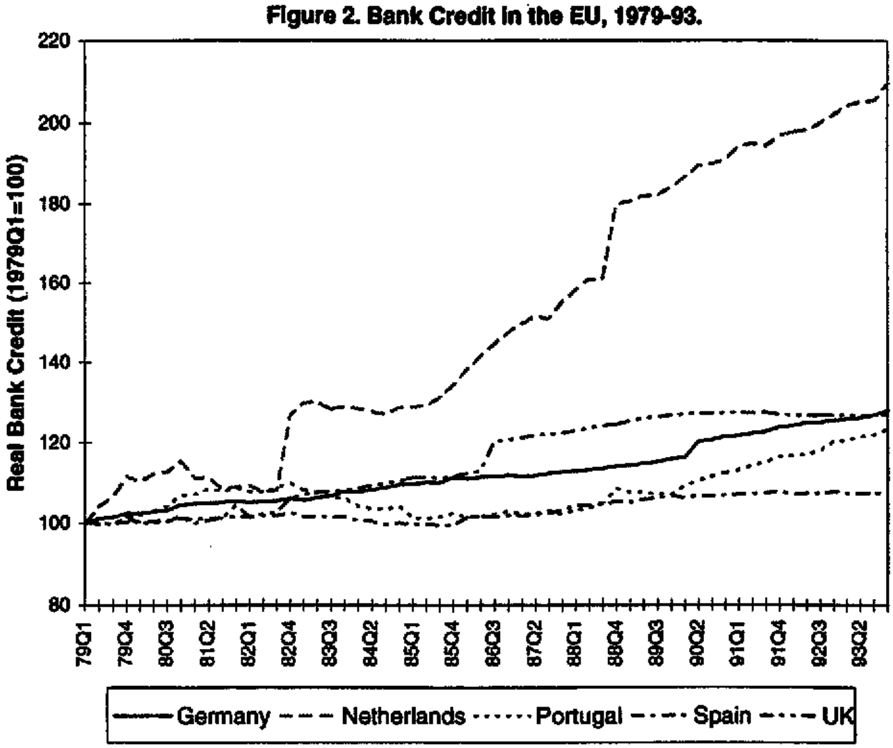


Figure 2. Bank Credit in the EU, 1979-93.



write the vector error correction model as

$$\begin{bmatrix} \Delta y_{1t} \\ \Delta y_{2t} \\ \Delta y_{3t} \end{bmatrix} = \Gamma_1 \begin{bmatrix} \Delta y_{1t-1} \\ \Delta y_{2t-1} \\ \Delta y_{3t-1} \end{bmatrix} + \begin{bmatrix} \alpha_{11} & \alpha_{12} \\ \alpha_{21} & \alpha_{22} \\ \alpha_{31} & \alpha_{32} \end{bmatrix} \begin{bmatrix} \beta_{11} & \beta_{21} & \beta_{31} \\ \beta_{12} & \beta_{22} & \beta_{32} \end{bmatrix} \begin{bmatrix} y_{1t-1} \\ y_{2t-1} \\ y_{3t-1} \end{bmatrix}. \tag{9}$$

If $\alpha_{3j} = 0$ for $j = 1,2$, then the equation for Δy_{3t} contains no information about the long-run since the cointegrating relationship does not enter the equation. The variable y_{3t} is therefore said to be weakly exogenous with respect to the system (see Johansen and Juselius (1992), Hall and Wickens (1993) and Harris (1995)). Toda and Phillips (1993) show that this type of test is preferable the standard VAR-based Granger-causality tests conducted in error correction models. With complications where there exists stochastic trends and cointegration in the system, the latter tests suffer from nuisance parameter dependencies where Wald tests for causality are not always valid asymptotic criteria. Johansen-type error correction models, however, deliver optimal estimates in cointegration space and provide a more promising basis than VARs for the sequential testing procedures that are needed to adequately test causality hypotheses.

4. The Results

Prior to estimating the model each time series was tested for the order of integration. A time series is said to be integrated of order $d[I(d)]$ if it must be differenced d times to become stationary. Table A1 presents the results from testing for unit roots for each of the series for the study period. The null hypotheses of non-stationarity are tested using Phillips-Perron tests. These non-parametric tests have an advantage over the standard Dickey-Fuller tests in that nuisance parameters are (asymptotically) eliminated if disturbances are not independent and identically distributed. In all cases, the null hypothesis of non-stationarity is accepted when the series are in levels, but is rejected when the series are in first differences. It is thus concluded that these series are $I(1)$ or first-difference stationary processes and therefore valid for examining the possibility of long-run relationships.

Bernard and Durlauf (1995) define convergence in multivariate output as when the long-run forecasts of output are equal at some time, t. If countries do not converge in this sense, they may still respond to the same long-run driving processes, i.e. they may face the same permanent shocks with different long-run weights. Moreover, countries contain a single common trend if the long-run forecasts of output are *proportional* at some time, t. A necessary but not sufficient condition for convergence is the presence of a single common shared trend. Hafer and Kutan (1994) argue that the existence of r cointegrating vectors among n variables where $r<n$ implies the existence of n-r shared trends. Should n-$r = 1$ then we have evidence of a *single* shared trend and therefore evidence of (weak) convergence. On this basis, stronger evidence of convergence would be implied if the null hypothesis of homogeneous cointegrating vectors is accepted. If r=0 (Π has a zero rank) then we have n stochastic trends but no shared trends thus no long-run convergence of policy, i.e. bank credit is unrelated across countries. If $0<r<n-1$ then we have more than one shared trend which could be interpreted as evidence of just *partial* convergence.[5]

Table A2 reports the tests for cointegrating relationships using the full sample of nine countries. There is evidence that six cointegrating vectors are present. This in turn implies that three common shared trends exist. The results reported in table 2 clearly reject the hypothesis of a single common shared trend. Given the disparate experiences among the sample, one would not expect to see evidence of a single shared trend. Italy and the UK have conducted a somewhat more lax membership of the ERM, Greece and Portugal remained outside, while France, Germany and the Netherlands conducted a relatively close monetary arrangement within the ERM. The implications of this latter point may be investigated if we consider a sub-sample comprising these latter countries. Table A3 suggests there is evidence of two shared common trends with a single cointegrating vector. Since the presence of a single common shared trend is strongly rejected, it would appear that even a relatively close commitment to ERM membership is unable to provide at least weak evidence in favour of convergence.

Using the full sample, we may test for weak exogeneity using the procedure outlined above. Table A4 reports that in all cases except Greece, the null of weak exogeneity with respect to the system, i.e. the full sample, is rejected at the 5 per cent level. With the exception of Greece, this means that we have evidence that bank credit provision in all countries is to some extent interrelated with other EU members and that no individual member can be described as truly independent. A necessary but not sufficient condition for German dominance (with respect to the system) would be the acceptance of weak exogeneity in the case of Germany. This is clearly rejected at the 5 per cent level with $\chi^2(6) = 17.233$. In line with earlier findings concerning financial integration in general, Germany plays a somewhat interactive role in EU financial markets. Another implication of these results is that peripheral EU members with a history of limited participation in the ERM, such as Portugal, Spain and the UK, are influenced by the stance of EU credit provision. A possible reason for this result might be that in the long run, credit policy is geared to controlling domestic inflation which, for competitive reasons, should not be too out of line with that prevailing in the rest of the EU. In the case of Greece, outside of the ERM for the entire study period, it would appear that even these considerations have not fostered closer credit market links with her EU partners.

The next stage of the empirical analysis is to investigate bivariate relationships with Germany. Table A5 reports that at the 5 per cent level of significance, there is a single cointegrating vector, or long-run equilibrium relationship, between France, Italy, the Netherlands, Portugal and Spain in turn and Germany. However, there is no long-run cointegrating relationship in the cases of Greece, Ireland and the UK. As mentioned, Greece has remained outside of the ERM. Ireland has been a member of the ERM but has employed capital controls and engaged in realignments, and the UK was a formal member of the ERM for only the period 1990Q4-1992Q3 operating with the wider ±6 per cent band of permitted exchange rate fluctuations. In the case of the five countries where a single cointegrating vector can be identified, it must be the case that causality runs in at least one direction. As before, a necessary but not sufficient condition for German dominance is that causality runs from Germany but not *vice-versa*. Table A6 reports evidence that long-run causality runs from Germany to each of the other five EU members. However, long-run causality also runs from France, Italy, the Netherlands and Portugal to Germany. Thus it is only in the case of Spain where the null of German weak exogeneity is accepted at the 5 per cent level with $\chi^2(1) = 1.390$. With the exception of the Netherlands, the null of weak exogeneity on the part of each country with respect to Germany is more strongly rejected than when the reverse hypothesis is considered. Thus while the results in Table A5 provide general evidence against German dominance, causality may be slightly stronger in terms of running from Germany.

In five out of eight cases, therefore, Germany causes bank credit provision in other member countries. Table A7 reports the normalised vectors identified from Table A5 where German casualty is confirmed. Following equation (5), increased values of ζ_1 imply increased sensitivity to credit market conditions in Germany while $\zeta_1 = 1$ would be suggestive of domestic and German credit markets moving closely in tandem. Alternatively, a much more limited influence from Germany would be confirmed with $\zeta_1 \cong 0$. Table A7 reports that estimates range from 0.599 in the case of Spain to 1.271 in the case of Italy. It is difficult to argue that commitment to ERM membership has a direct bearing on the closeness of ζ_1 to unity. Further analysis of the bivariate relationships can be made in terms of the estimated speeds of adjustment towards long-run equilibrium (α). Table A7 reports that France, Italy and the Netherlands have much faster speeds than Portugal and Spain. These differences might be attributable to the relative degree of development in domestic credit markets which allows supply and demand to respond more rapidly to given shocks.

Having established that the role of Germany in EU credit markets is interactive rather than dominant, the final stage of the analysis is to consider whether Germany can be regarded as the linchpin of credit market activity across the EU. This can be done through a comparison with the presence, if any, of cointegrating relationships among non-German EU members. Table A8 indicates where bivariate cointegrating relationships are present in the full sample. Germany, for which there are five cointegrating relationships present, has the highest number long-run relationships present. In contrast, one might note the very limited number of cointegrating relationships which involve Greece, Ireland and the UK.

5. Summary and Conclusion

Following an analysis of common trends and causality tests, this study indicates that future studies of credit provision in the EU should adopt an international rather than purely domestic focus. In the multivariate setting, the presence of three common shared trends among nine countries is evidence against convergence during the ERM period. However, with the exception of Greece, no country is weakly exogenous with respect to all other countries. With regard to Germany, this finding of interaction offers evidence against German dominance. In the bivariate setting, long-run equilibrium relationships with Germany are identified for France, Italy, the Netherlands, Portugal and Spain. With the exception of Spain, causality appears to run both ways rather than just from Germany. An examination of bivariate relationships between all countries in the sample are suggestive of Germany being at the centre of EU credit market activity. An implication of this study is that France, Italy and the Netherlands are in the strongest position to proceed towards EMU. The credit markets of these economies are closely and interactively related to Germany's and are characterised by speeds of adjustment that allow a relatively rapid adjustment to shocks.

Endnotes

1. Department of Economics, Loughborough University. I would like to acknowledge the helpful comments of two anonymous referees and accept full sole responsibility for any remaining errors.

2. See, for example, Moore and Threadgold (1980), Cuthbertson (1985), Hartropp (1992).

3. Data availability means that national GDP deflators are used in all cases except for producer prices in the cases of Greece and the Netherlands and

4. For the purposes of graphical exposition, all series are indexed to 1979Q1=100.

5. Hall *et al* (1992) argue that cointegration may have a tendency to reject convergence in the face of structural change. This study is concerned with the evaluation of convergence and integration for the ERM period as a whole rather than the evolution of convergence and integration during this time period.

Appendix

Table A1: Unit root tests

	$I(0)$	$I(1)$
France	-1.129	-8.063**
Germany	-0.750	-8.813**
Greece	-2.475	-8.733**
Ireland	-1.341	-8.812**
Italy	-2.425	-8.094**
Netherlands	-2.095	-7.610**
Portugal	-2.135	-7.686**
Spain	-1.340	-7.996**
UK	-1.027	-8.579**

These are Phillips-Perron unit root tests whih include a constant and trend. ** indicates rejection of the null hypothesis of non-stationarity at the 5% level of significance with critical value of -3.49. Estimates obtained through *Shazam v7.0*.

Table A2: Tests for cointegration: full sample

H_0	*Statistic*	*95% critical value*
$r=0$	508.9**	192.9
$r\leq1$	355.0**	156.0
$r\leq2$	228.1**	124.2
$r\leq3$	158.8**	94.2
$r\leq4$	104.8**	68.5
$r\leq5$	52.83**	47.2
$r\leq6$	28.84	29.7
$r\leq7$	11.26	15.4
$r\leq8$	0.185	3.8

Tests for cointegration are Johansen's likelihood ratio test based on the trace of the stochastic matrix (see Johansen and Juselius (1990) for more details) where r refers to the number of cointegrating vectors, and H_0 refers to the null hypothesis. Maximum lag length of 4 in the VAR determined by information criteria. Following the application of the Pantula principle, estimates include an unrestricted constant and seasonal dummies. ** indicates rejection of the null at the 5% level of significance. Significance levels taken from Osterwald-Lenum (1992), table 1. Estimates obtained through *PC-FIML*.

Table A3: Tests for cointegration: France, Germany and the Netherlands

H_0	$r=0$	$r\leq1$	$r\leq2$
Statistic	36.14**	10.16	0.277
95% critical value	29.7	15.4	3.8

See notes for table A2. Maximum lag length in VAR is 4. Estimates include unrestricted seasonals and constant.

Table A4: Tests for weak exogeneity

	Fr.	*Ger.*	*Greece*	*Irel'd*	*Italy*	*Neth'ds*	*Port.*	*Spain*	*UK*
$\chi^2(6)$	40.967**	17.233**	10.189	16.615**	80.660**	26.977**	44.865**	21.992**	39.157**

These tests are conducted on the α matrix (based on the results reported in Table 2) and are distributed as $\chi^2(r)$ on the null hypothesis of weak exogeneity where r refers to the number of significant cointegrating vectors. ** denotes rejection of the null hypothesis at the 5% level of significance with critical value of $\chi^2(6)$. Estimates obtained through *PC FIML*.

Table A5: Tests for cointegratin: bivarate case with Germany								
	Fr.	*Greece*	*Irel'd*	*Italy*	*Neth'ds*	*Port.*	*Spain*	*UK*
k	4	4	3	5	6	9	9	3
r=0	53.64**	8.558	12.22	46.61**	28.93**	28.55**	24.75**	4.874
$r\leq1$	9.054	1.192	4.768	5.732	8.712	4.513	8.165	0.720

See notes for Table 2. k refers to the lag length of the VAR. Following the Pantula principal, estimates for France, Germany, Ireland, Portugal and the UK include an unrestricted constant while remaining estimates include a restricted constant. Following Osterwald-Lenum (1992), ** denotes significance at the 5% level.

Table A6: Tests for weakexogeneity: bivariate case			
H_0	χ^2	H_0	χ^2
Germany → France	21.708**	Netherlands → Germany	20.501**
France → Germany	17.792**	Germany → Portugal	16.836**
Germany → Italy	35.025**	Portugal → Germany	4.438**
Italy → Germany	5.688**	Germany → Spain	6.740**
Germany → Netherlands	14.597**	Spain → Germany	1.390

See notes for table 4. $X \rightarrow Y$ denotes the null that country Y is weakly exogenous with respect to country X. ** denotes rejection of the null of weak exogeneity at the 5% level of significance with the critical value $\chi^2(1)$=3.84.

Table A7: Long-run equilibrium relationships with Germany					
	France	*Italy*	*Netherlands*	*Portugal*	*Spain*
ζ_1	1.022	1.271	1.203	1.166	0.599
α	-0.029	-0.079	-0.056	-0.015	-0.003

Cointegrating vectors obtained in table 5 are normalised with respect to France, Italy, the Netherlands, Portugal and Spain. ζ_1 as defined in equation (5), α is the speed of adjustment towards long-run equilibrium as defined in equation (8).

Table A8: Bivariate cointegrating relationships among full sample

	Fra	*Ger*	*Gre*	*Ire*	*Ita*	*Nth*	*Por*	*Spa*	*UK*
Fra	-	Y	x	x	Y	x	Y	x	x
Ger		-	x	x	Y	Y	Y	Y	x
Gre			-	x	x	x	x	x	x
Ire				-	Y	x	x	x	x
Ita					-	Y	x	x	Y
Nth						-	x	x	x
Por							-	Y	x
Spa								-	x
UK									-

Y confirms rejection of the null of no cointegration while x indicates acceptance of the null at the 5% level of significance. Johansen procedure applied throughout, see table 2 for details.

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