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Core and Periphery in the European Monetary Union: Bayoumi and Eichengreen 25 Years Later*

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Abstract: Bayoumi-Eichengreen (1993) establish a EMU core-periphery pattern using 1963-1988 data. We use same methodology, sample, window length (1989-2015), and a novel over-identifying restriction test to ask whether the EMU strengthened or weakened the core-periphery pattern. Our results suggest the latter.

JEL classification: E32, E63, F02
Keywords: Business cycle synchronization, Structural VAR, European Monetary Union, Core-periphery

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1. Introduction

The seminal paper by Bayoumi and Eichengreen (1993) highlights the existence of a core-periphery pattern in the run-up to the European Monetary Union (EMU). If persistent, this pattern would be detrimental to the EMU project. Using pre-EMU data to estimate the degree of supply shocks synchronization, they argue that there is a core (Germany, France, Belgium, Netherlands and Denmark) where shocks are highly correlated and a periphery (Greece, Ireland, Italy, Portugal, Spain and UK) where synchronisation is significantly lower.

The objective of this paper is to revisit Bayoumi and Eichengreen (1993) in order to evaluate the effect of the EMU on the core-periphery pattern they find using 1963-1988 data. We use the same estimation methodology, sample, and time window (25 years) to replicate their results for 1989-2015. We ask whether the EMU strengthened or weakened the core-periphery pattern. Based on a new over-identifying restriction test, our results suggest that the core-periphery pattern has actually weakened.

2. Theory

The main research question driving the scholarship on optimal currency areas (OCA) regards the costs and benefits of sharing a currency (Alesina and Barro, 2002). The main cost is the loss of monetary policy autonomy, while the main benefits are transaction costs and exchange rate uncertainty reductions, and increasing price transparency, trade and competition. OCA theory stresses labour mobility, product diversification and trade openness as criteria while debating the endogeneity of currency unions (Frankel and Rose, 1998). Recent work highlights the role of credibility shocks: with varying degrees of commitment (time inconsistency), countries with dissimilar credibility shocks should join currency unions (Chari et al 2015). A second relevant recent strand highlights situations in which OCA criteria are modelled as interdependent. For instance, Farhi and Werning (2015) focus on interactions between openness and mobility. Recent econometric evidence showing the absence of a robust effect of currency unions on trade raises caveats to the discussion above (Glick and Rose, 2016).

3. Estimation

The methodology used by Bayoumi and Eichengreen (1993) is an extension of the Blanchard and Quah (1989) procedure for decomposing permanent and temporary shocks. Consider a system where the true model is represented by an infinite moving average of a (vector) of variables, X_t, and shocks, ϵ_t. Using the lag operator L, a bi-variate VAR featuring real GDP and its deflator can be written as an infinite moving average representation of demand and supply disturbances:

$$X_t = A_0\epsilon_t + A_1\epsilon_{t-1} + A_2\epsilon_{t-2} + A_3\epsilon_{t-3} + \cdots = \sum_{i=0}^{\infty} L^i A_i \epsilon_t \quad (1)$$

where $X_t = [\Delta y_t, \Delta p_t]$ and the matrices A represent the impulse response functions of the shocks to the elements of X. It follows that

$$\begin{bmatrix} \Delta y_t \\ \Delta p_t \end{bmatrix} = \sum_{i=0}^{\infty} L^i \begin{bmatrix} a_{11i} & a_{12i} \\ a_{21i} & a_{22i} \end{bmatrix} \begin{bmatrix} \epsilon_{dt} \\ \epsilon_{st} \end{bmatrix} \quad (2)$$

where $y_t$and p_t represent the logarithm of output and prices and ϵ_t are $i.i.d.$ disturbances, which identify supply and demand shocks (Ramey, forthcoming). For the i-th country, a_{11i} represents element a_{11}, in matrix A_i and so on.

This framework implies that supply shocks have permanent effects on output, while demand shocks have temporary effects. Both have permanent (opposite) effects on prices. The cumulative effect of demand shocks on the change in output must be zero:

$$\sum_{i=0}^{\infty} a_{11i} = 0 \tag{3}$$

Using the standard relation between the VAR's residuals (e_t) and demand and supply shocks, i.e. $e_t = C\epsilon_t$ for each country, exact identification of the C matrix requires four restrictions. Two are normalizations, which define the variance of the shocks ϵ_{dt} and ϵ_{st}. The third restriction is from assuming that demand and supply shocks are orthogonal to each other. The fourth that demand shocks have only temporary effects on output (equation 3).

Based on the standard AD-AS model, there is one restriction that Bayoumi and Eichengreen (1993) do not impose as their model was exactly identified. Here we extend their framework by imposing a fifth, additional over-identifying restriction and we explicitly test for a permanent effect of supply shocks on output by imposing $\sum_{i=0}^{\infty} a_{12i} = \gamma$, where $\gamma > 0$. Accordingly, demand in each country is restricted to respond to supply shocks qualitatively (sign) and quantitatively (size) in the same way. In terms of the structural VAR analysis, this implies:

$$\sum_{i=1}^{\infty} \begin{bmatrix} d_{11i} & d_{12i} \\ d_{21i} & d_{22i} \end{bmatrix} \begin{bmatrix} c_{11} & c_{12} \\ c_{21} & c_{22} \end{bmatrix} = \begin{bmatrix} 0 & \gamma \\ . & . \end{bmatrix} \tag{4}$$

We do not restrict γ *a priori*; instead, we vary γ in the interval [0.1, 2] and choose its value optimally, as explained below (the number we chose to report is $\gamma = 1$.)

3.1 Testing for over-identifying restriction

In order to test for the over-identifying restriction described above, we estimate Bayoumi and Eichengreen (1993) SVAR model. Differently from them, we bootstrap the original VAR residuals in a *i.i.d.* fashion and generate K = 10.000 data sets. For each of the k-*th* samples we test for the over-identifying restriction based on a LR-test. We record the number of rejections (NoR) of the over-identifying restriction test at each bootstrap replication, and calculate

$$NoR_i = 100 \times \frac{\sum_{k=1}^{K} \left\{ NoR = 1 \middle| -2(L_r - L_u) > \chi^2_{q-\left(\frac{n^2-n}{2}\right)} \right\}_{i,k}}{K} \qquad (5)$$

where L_u and L_r are the maximized values of the (Gaussian) log likelihood function of the unrestricted and restricted regressions, respectively. Under H_0, the LR statistic has an asymptotic distribution with degrees of freedom equal to the number of long-run restrictions (q) minus $(n^2 - n)/2$, where n is the VAR-dimension (in this case $n = 2$). We calculate NoR_i for different values of γ.

Based on the results in Table 2A (Cf. Appendix), we chose the value of γ which minimizes the total number of rejections in our sample. Demand and supply shocks are then retrieved by bootstrap, specifically by recalculating the VAR parameters (K = 10.000), identifying the SVAR and considering median values of structural disturbances under $\gamma = 1$.

4. Results

Figure 1 shows our main results. The residuals (median bootstrapped) are retrieved from a Structural VAR with two lags for all countries, no constant, and using yearly data with respect to

Germany closely following Bayoumi and Eichengreen (1993). The over-identifying restriction is imposed and the sample is 1989–2015. As dispersion has decreased compared to the pre-EMU era, we argue the results suggest the core-periphery pattern has weakened after 1989.

[Figure 1 about here]

Based on the bootstrapped VAR, we test for the over-identifying restriction described above where (non) rejection supports classifying the country as periphery (centre). The four countries for which the rejection of the over-identifying restriction is stronger, at conventional significance levels, are Ireland, Spain, Greece and Portugal (Table 2A in the Appendix).[1] Without imposing this over-identifying restriction for these four countries, the core-periphery pattern in Bayoumi and Eichengreen's terms actually weakens even further. When the over-identifying restriction is not imposed, Ireland and Portugal move down the demand-axis and Greece and Spain jump to the left (Figure 2).

[Figure 2 about here]

Overall, our results support a re-interpretation of the core-periphery pattern: after EMU a new, smaller periphery emerges (Spain, Portugal, Ireland and Greece) and its dynamics is systematically different from the rest in that, for these countries, the over-identifying restriction is rejected by the data in most cases.

One important concern is that the relationship between demand and supply may have changed over time and/or the nature of shocks has been altered by the EMU itself. Hence, a structural identification on economic variables that may have changed can be misleading. One can argue that the increase in correlation in supply disturbances may be due to a larger role for

[1] The UK shows an ambiguous development: higher correlation of supply shocks but lower correlation of demand shocks.

oil price shocks in the sample. Proponents of using the nominal price of oil in empirical models of the transmission of oil price shocks tend to conclude that there is no stable dynamic relationship between percent changes in the nominal price of oil and inflation. There is evidence from in-sample fitting exercises, however, of a predictive relationship between suitable nonlinear transformations of the nominal price of oil and real output. The most successful of these transformations is the Net Oil Price Increase (NOPI) measure from Hamilton (2003). Let s_t denote the nominal price of oil in logs, then

$$NOPI_t = \begin{cases} s_t - \max(s_{t-1}, s_{t-37}) & if \;\; s_t - \max(s_{t-1}, s_{t-37}) > 0 \\ 0 & otherwise \end{cases} \tag{6}$$

The net oil price increase is a censored predictor that assigns zero weight to net oil price decreases and singles out oil prices peaks in a 36-month (or shorter) window. To construct a Net Oil Price Index, we use the Brent Europe crude oil price index at a monthly frequency and identify the net increases (Figure 3.) Based on this characterization, we define dummy variables at a yearly frequency. In particular, we identify the following net oil increases $\{1996, 1999, 2000, 2004 \; to \; 2008\}$. When conditioning the VAR on the NOPI, we find little evidence that this is relevant in this framework and that the responses of real GDP and inflation to demand and supply innovations are driven by net oil price increases (results also remain broadly unchanged if we use the change in the price of oil as exogenous variable instead).

[Figures 3 and 4 about here]

5. Conclusions

Bayoumi and Eichengreen (1993) is a seminal paper because, *inter alia*, it is one of the first to point out the risks of an entrenched core-periphery to the then nascent EMU. Their influential diagnostics was based upon data covering 25 years from 1963 to 1988. Using the same methodology, sample, and time window, this paper replicates their results for 1989-2015. We ask whether the EMU strengthened or weakened the core-periphery pattern. Using a new over-identifying restriction test, our results suggest the EMU has significantly weakened the original pattern described in Bayoumi and Eichengreen, in that we find, based on demand and supply shocks, changes in the clustering of countries.

References

Alesina, A. & R. Barro (2002), "Currency Unions", *Quarterly Journal of Economics*, 117 (2): 409-436.

Bayoumi, T. & B. Eichengreen (1993), "Shocking Aspects of European Monetary Integration," in F. Torres and F. Giavazzi (eds), *Adjustment and Growth in the European Monetary Union*, Cambridge University Press.

Blanchard, O. & D. Quah (1989), "The Dynamic Effects of Aggregate Demand and Aggregate Supply Disturbances," *American Economic Review*, 79 (1989), pp. 655–673.

Chari V.V., Dovis, A. & P. Kehoe (2015), "Rethinking Optimal Currency Areas", Federal Reserve Bank of Minneapolis, Research Department Staff Report.

Farhi, E. & I. Werning, (2015) "Labor Mobility in Currency Unions," MIT mimeo.

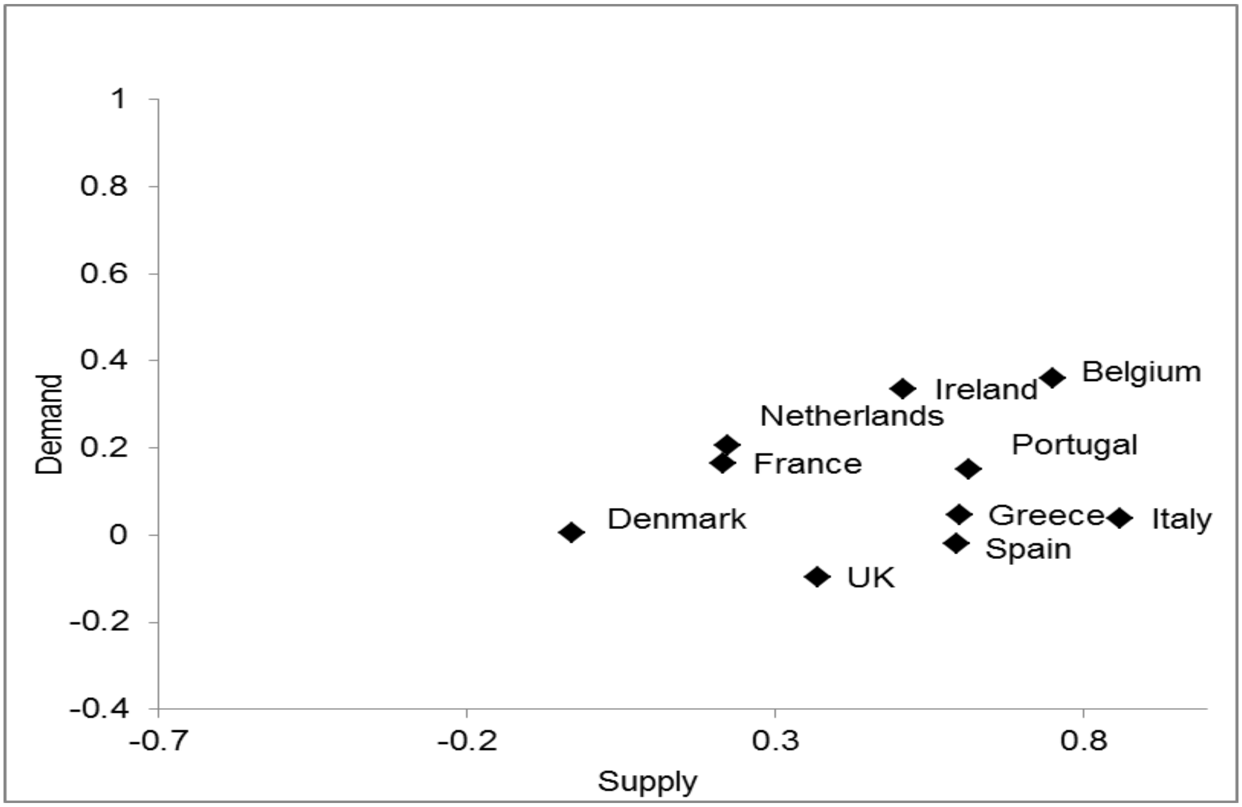
Frankel, J. & A. Rose (1998), "The Endogeneity of the Optimum Currency Area Criteria," *Economic Journal*, 108 (449), 1009-1025.

Glick, R. & A. Rose (2016), "Currency Unions and Trade: A Post-EMU *Mea Culpa*," mimeo.

Hamilton, J. (2003), "What is an Oil Shock?" *Journal of Econometrics*, 113, 363-398.

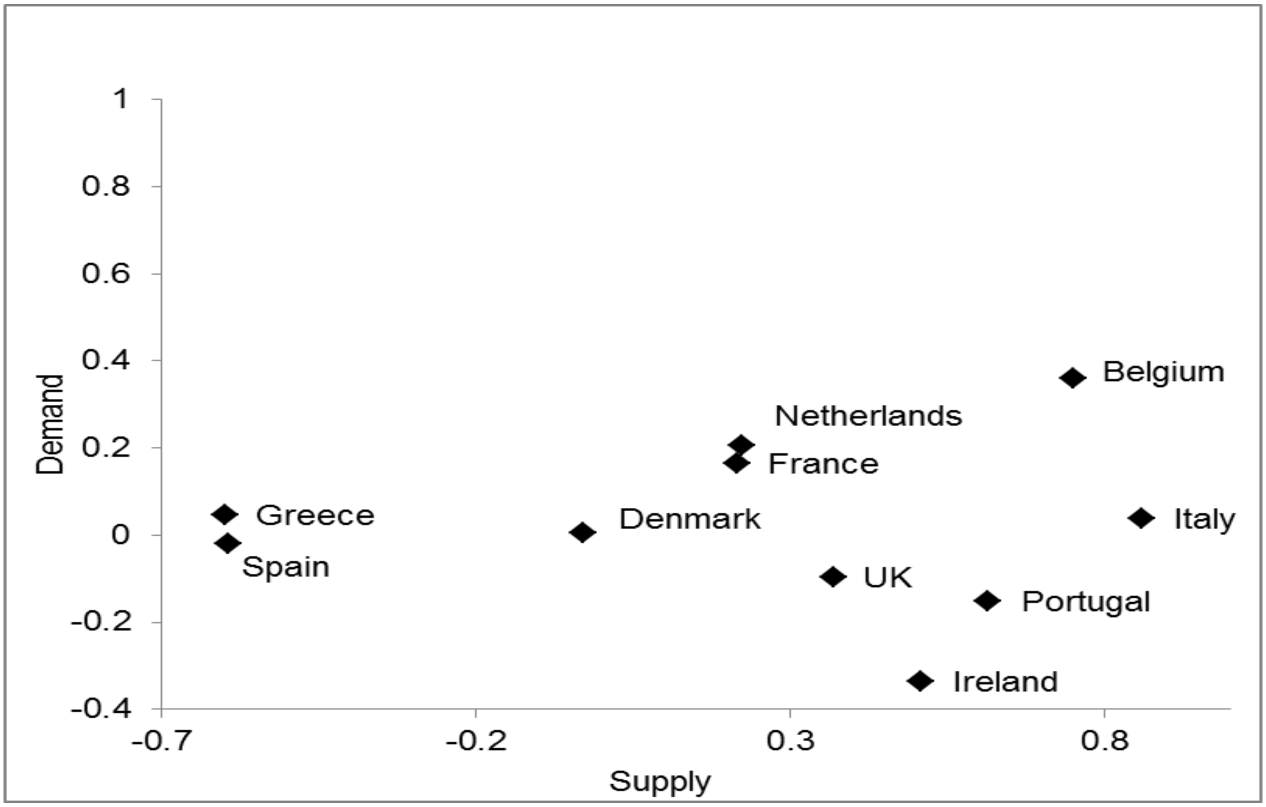
Ramey, V. (forthcoming), "Macroeconomic Shocks and Their Propagation," in J. Taylor and H. Uhlig (eds.), *Handbook of Macroeconomics, Volume 2*, Elsevier.

Figure 1 – Correlation of supply and demand disturbances imposing the over-identifying restriction (bootstrapped residuals – median values)



Note: This figure reports median bootstrapped residuals based on 10.000 VAR replications. Structural residuals are retrieved from a SVAR where the over-identifying restriction above is imposed for all countries. The sample for this SVAR is 1989–2015, with two lags for all countries and no constant as in Bayoumi and Eichengreen (1993). The demand and supply disturbances correlation coefficients vis-à-vis Germany are reported in Appendix Table 3A.

Figure 2 – Correlation of supply and demand disturbances (bootstrapped residuals – median values) relaxing the over-identifying restriction



Note: This figure reports median bootstrapped residuals based on 10.000 VAR replications. Structural residuals are retrieved from a SVAR where the over-identifying restriction above is imposed for all countries, with the exception of Ireland, Spain, Greece and Portugal. The sample for this SVAR is 1989–2015, with two lags for all countries and no constant as in Bayoumi and Eichengreen (1993). The demand and supply disturbances correlation coefficients are reported in Appendix Table 4A.

Figure 3 – Net Oil Price Increases Indicator

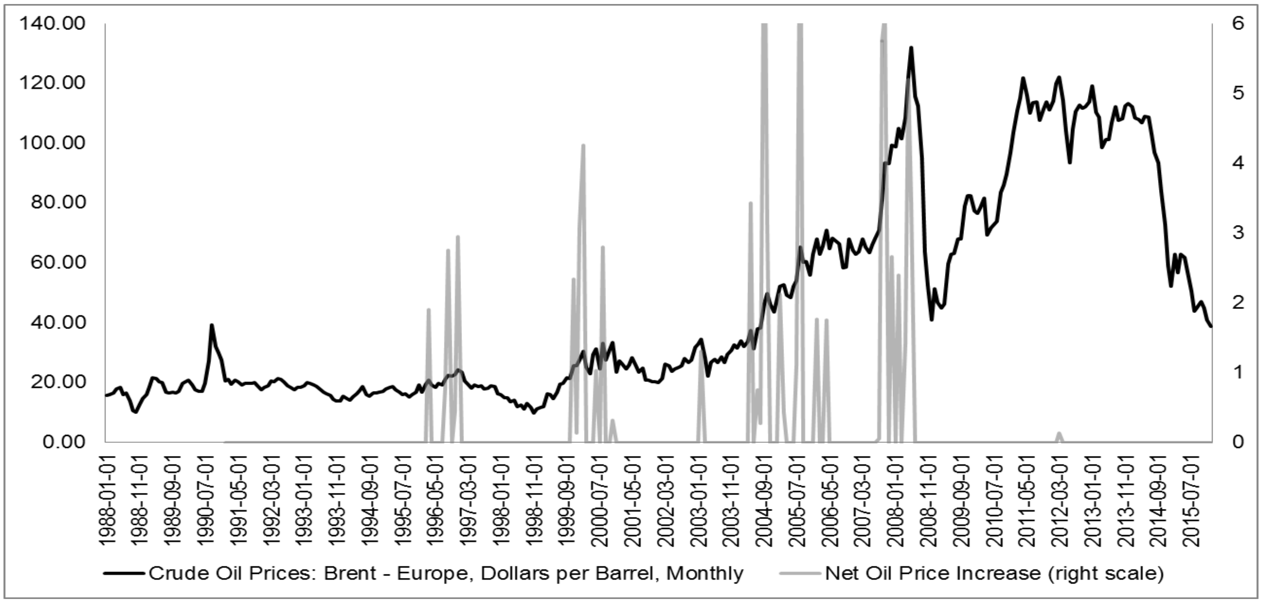
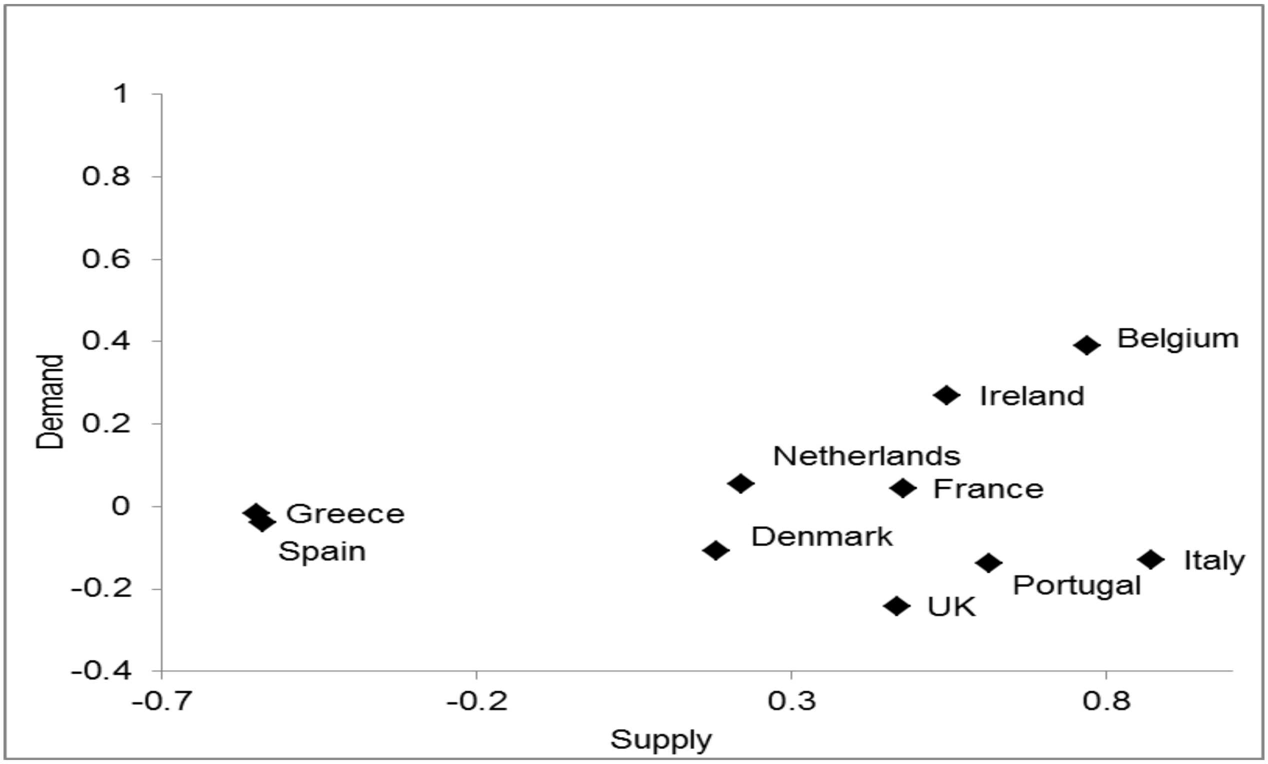


Figure 4 – Correlation of supply and demand disturbances relaxing the over-identifying restriction and conditional on NOPI (bootstrapped residuals – median values)



Note: This figure reports median bootstrapped residuals based on 10.000 VAR replications. Structural residuals are retrieved from a SVAR where the over-identifying restriction above is imposed for all countries, with the exception of Ireland, Spain, Greece and Portugal. The sample for this SVAR is 1989–2015, with two lags for all countries and no constant as in Bayoumi and Eichengreen (1993). The SVAR is conditional on NOPI dummies (Cf. Results' section).

APPENDIX

(Supplementary Material

On-line Appendix

Not for Publication)

APPENDIX 1: Estimation

The methodology used by Bayoumi and Eichengreen (1993) is an extension of the Blanchard and Quah (1989) procedure for decomposing permanent and temporary shocks. Consider a system where the true model is represented by an infinite moving average of a (vector) of variables, X_t, and shocks, ϵ_t. Using the lag operator L, a bi-variate VAR featuring real GDP and its deflator can be written as an infinite moving average representation of demand and supply disturbances:

$$X_t = A_0\epsilon_t + A_1\epsilon_{t-1} + A_2\epsilon_{t-2} + A_3\epsilon_{t-3} + \cdots = \sum_{i=0}^{\infty} L^i A_i \epsilon_t \quad (1.1)$$

where $X_t = [\Delta y_t, \Delta p_t]$ and the matrices A represent the impulse response functions of the shocks to the elements of X. It follows that

$$\begin{bmatrix}\Delta y_t \\ \Delta p_t\end{bmatrix} = \sum_{i=0}^{\infty} L^i \begin{bmatrix} a_{11i} & a_{12i} \\ a_{21i} & a_{22i}\end{bmatrix} \begin{bmatrix}\epsilon_{dt} \\ \epsilon_{st}\end{bmatrix} \quad (1.2)$$

where $y_t$and p_t represent the logarithm of output and prices and ϵ_t are $i.i.d.$ disturbances, which identify supply and demand shocks (Ramey, forthcoming). For the i-th country, a_{11i} represents element a_{11}, in matrix A_i and so on.

This framework implies that supply shocks have permanent effects on output, while demand shocks have temporary effects. Both have permanent (opposite) effects on prices. The cumulative effect of demand shocks on the change in output must be zero:

$$\sum_{i=0}^{\infty} a_{11i} = 0 \quad (1.3)$$

So it can be estimated using a VAR. Each element can be regressed on lagged values of all the elements of X. Using B to represent these estimated coefficients:

$$X_t = B_1 X_{t-1} + B_2 X_{t-2} + \cdots + B_n X_{t-n} + e_t$$

$$= \left(I - B(L)\right)^{-1} e_t \quad (1.4)$$

$$= (I + B(L) + B(L)^2 + \cdots)e_t$$

$$= e_t + D_1 e_{t-1} + D_2 e_{t-2} + D_3 e_{t-3}$$

where e_t represents the residuals from the VAR equations. In order to convert (1.4) into the model in (1.2) under (1.3), the residuals from the VAR, e_t, are transformed into demand and supply shocks. Using the standard relation between the VAR's residuals (e_t) and demand and supply shocks, i.e. $e_t = C\epsilon_t$, it is clear that, for each country, exact identification of the C matrix requires four restrictions. Two are normalizations, which define the variance of the shocks ϵ_{dt} and ϵ_{st}. The third restriction is from assuming that demand and supply shocks are orthogonal to each other. The fourth that demand shocks have only temporary effects on output (equation 1.3).

The standard AD-AS model implies that demand shocks should raise prices in both the short and long run, while supply shocks should lower prices and increase demand permanently.

In order to achieve that, it suffices to impose the additional over-identifying restriction in the VAR that supply shocks have permanent effects on output. We need to impose this restriction in our sample for the demand and supply shocks to be identified. This differs from Bayoumi and Eichengreen (1993) because they do not impose this last restriction, which leaves the model exactly identified. One reason we adopt the proposed over-identifying restriction is that inflation differentials are often considered a ‘normal feature of currency unions. Therefore, we pay particular attention to modelling the effect of shocks on demand. The role of co-movements in output’s cyclical fluctuations is further in line with the business-cycle literature. Since the proposed over-identifying restriction is sufficient to get structural disturbances in line with AD-AS dynamics, any additional long-run restriction may be redundant in this setting.

We test for the above over-identifying restriction, by imposing $\sum_{i=0}^{\infty} a_{12i} = \gamma$, where $\gamma > 0$. Under the latter assumption, demand across each country is restricted to respond qualitative (sign) and quantitative (size) in the same way to supply shocks. In terms of the structural VAR analysis, this implies:

$$\sum_{i=1}^{\infty} \begin{bmatrix} d_{11i} & d_{12i} \\ d_{21i} & d_{22i} \end{bmatrix} \begin{bmatrix} c_{11} & c_{12} \\ c_{21} & c_{22} \end{bmatrix} = \begin{bmatrix} 0 & \gamma \\ . & . \end{bmatrix} \tag{1.5}$$

We do not restrict γ *a priori*; instead, we vary γ in the interval [0.1, 2] as shown in Table 2A.

In order to construct a test for the over-identifying restriction described above, we estimate the SVAR model consistent with Bayoumi and Eichengreen (1993). Differently from the latter, we bootstrap the original VAR residuals in a *i.i.d.* fashion and generate K = 10.000 data sets. For each of the k-*th* samples we proceed with a structural analysis and test for the over-identifying restriction based on a LR-test. We record the number of rejections of the over-identifying restriction test at each bootstrap replication, and calculate

$$NoR_i = 100 \times \frac{\sum_{k=1}^{K} \left\{ NoR = 1 \middle| -2(L_r - L_u) > \chi^2_{q-\left(\frac{n^2-n}{2}\right)} \right\}_{i,k}}{K}$$

where L_u and L_r are the maximized values of the (Gaussian) log likelihood function of the unrestricted and restricted regressions, respectively. Under H_0, the LR statistic has an asymptotic distribution with degrees of freedom equal to the number of long-run restrictions (q) minus $(n^2 - n)/2$, where n is the VAR-dimension (in this case $n = 2$). We calculate NoR_i for different values of γ.

Based on the results in Table 2A, we chose the value of γ which minimizes the total number of rejections in our sample. Demand and supply shocks are then retrieved by bootstrap, in particular recalculating the VAR parameters (K = 10.000), identifying the SVAR and considering median values of structural disturbances under $\gamma = 1$.

APPENDIX 2: Data

Annual data: Annual data on real and nominal GDP spanning the period 1989 - 2015 (Portugal 1989 - 2014) were collected from the OECD Annual National Accounts for the 12 members of the EC. As in Bayoumi and Eichengreen (1993), Germany is used as a numeraire country. For each country growth and inflation were calculated as the first difference of the logarithm of real GDP (OECD base year) and the implicit GDP deflator. In line with BE the deflator was used to measure prices since it reflects the price of output rather than the price of consumption. Some descriptive statistics of the raw data are presented in Table 1A. The series used in the VAR were corrected for different regimes in mean, before 1992 – consistent with the pre-Maastricht period, as well as the British sterling and Italian lira EMS dismissal – and after 2007.

Monthly data: Crude Oil Prices: Brent - Europe, Dollars per Barrel, not seasonally adjusted (Source: Federal Reserve Bank of St. Louis – FRED Database). The series is seasonally adjusted using a standard X12 ARIMA model.

Legend (in alphabetical order)**:**

BE = Belgium

DE = Germany

DK = Denmark

ES = Spain

FR = France

GR = Greece

IE = Ireland

IT = Italy

NL= Netherlands

PT = Portugal

UK = United Kingdom

APPENDIX 3: Results

Table 1A – Standard deviation and correlation coefficients with Germany: Log of raw data

	Growth		Inflation	
	St. dev	Correlation	St. dev	Correlation
BE	1.450	0.749	0.959	0.597
DE	2.125	1	1.225	1
DK	1.915	0.628	1.019	-0.121
ES	2.312	0.528	2.158	0.463
FR	1.472	0.750	0.760	0.325
GR	3.927	0.167	5.732	0.641
IE	3.875	0.452	2.701	-0.045
IT	1.911	0.778	1.923	0.546
NL	1.936	0.730	0.975	-0.089
PT	2.501	0.618	3.290	0.689
UK	1.706	0.330	1.768	0.400

Note: All variables are measured in log percent, so e.g. 2.125 for Germany indicates approximately standard deviation of 2.125 percent.

Table 2A – Test for over-identifying restrictions' count (% of bootstrap replications)

	# of rejections $\gamma = 0.1$	# of rejections $\gamma = 0.5$	# of rejections $\gamma = 1$	# of rejections $\gamma = 1.5$	# of rejections $\gamma = 2$
BE	100.0	66.2	17.4	53.5	83.9
DE	99.8	94.0	25.1	18.3	47.1
DK	100.0	95.6	35.5	16.2	36.8
ES	99.8	99.0	74.2	35.4	21.8
FR	100.0	77.5	20.3	39.3	68.5
GR	94.6	100.0	92.5	63.5	35.8
IE	100.0	100.0	98.4	86.8	64.9
IT	100.0	69.4	14.6	51.8	84.9
NL	100.0	93.7	20.2	17.3	50.0
PT	100.0	99.9	89.2	53.2	24.8
UK	99.8	94.0	50.2	27.2	33.6
Total largest EZ3	99.9	86.7	30.9	32.4	54.5
Total largest EZ5	99.9	88.4	21.9	25.0	55.2
Total EZ9	99.4	88.9	50.2	46.6	53.5
No of countries > threshold	11	11	4	5	4

Note: We bootstrap the original VAR residuals in a i.i.d. fashion and generate 10.000 data sets. For each of the 10.000 samples we recalculate the VAR parameters. At each replication we proceed with the SVAR analysis proposed by Bayoumi and Eichengreen (1993) and further impose the over-identifying restriction by counting the number of rejections. Cut off value is that of a $\chi^2(1)$ with probability 0.999 (10.828). The results are robust if this probability is reduced to 0.99 (6.635). The countries for which this restriction is rejected on average more than in 50.5% of cases are the ones for which the over-identifying restriction is relaxed. For consistency of the results, the number of cases the SVAR does not converge is excluded from the count.

Table 3A – Correlation of supply and demand disturbances *vis-à-vis* Germany imposing the over-identifying restriction

	Supply shocks	Demand shocks
BE	0.750	0.360
DE	1	1
DK	-0.029	0.005
ES	0.594	-0.019
FR	0.216	0.164
GR	0.599	0.047
IE	0.508	0.335
IT	0.859	0.039
NL	0.223	0.205
PT	0.614	0.152
UK	0.368	-0.096

Note: Structural disturbances are retrieved from a SVAR where the over-identifying restriction described in Section 3 is imposed for all countries, with the exception of Ireland, Spain, Greece and Portugal. The reported values are median values based on 10.000 bootstrap replications. The sample is 1989 – 2015, with the SVAR being solved using 2 lags for all countries and no constant as in Bayoumi and Eichengreen (1993).

Table 4A – Correlation of supply and demand disturbances *vis-à-vis* Germany relaxing the over-identifying restriction for Ireland, Spain, Greece and Portugal

	Supply shocks	Demand shocks
BE	0.750	0.360
DE	1	1
DK	-0.029	0.005
ES	-0.594	-0.019
FR	0.216	0.164
GR	-0.599	0.047
IE	0.508	-0.335
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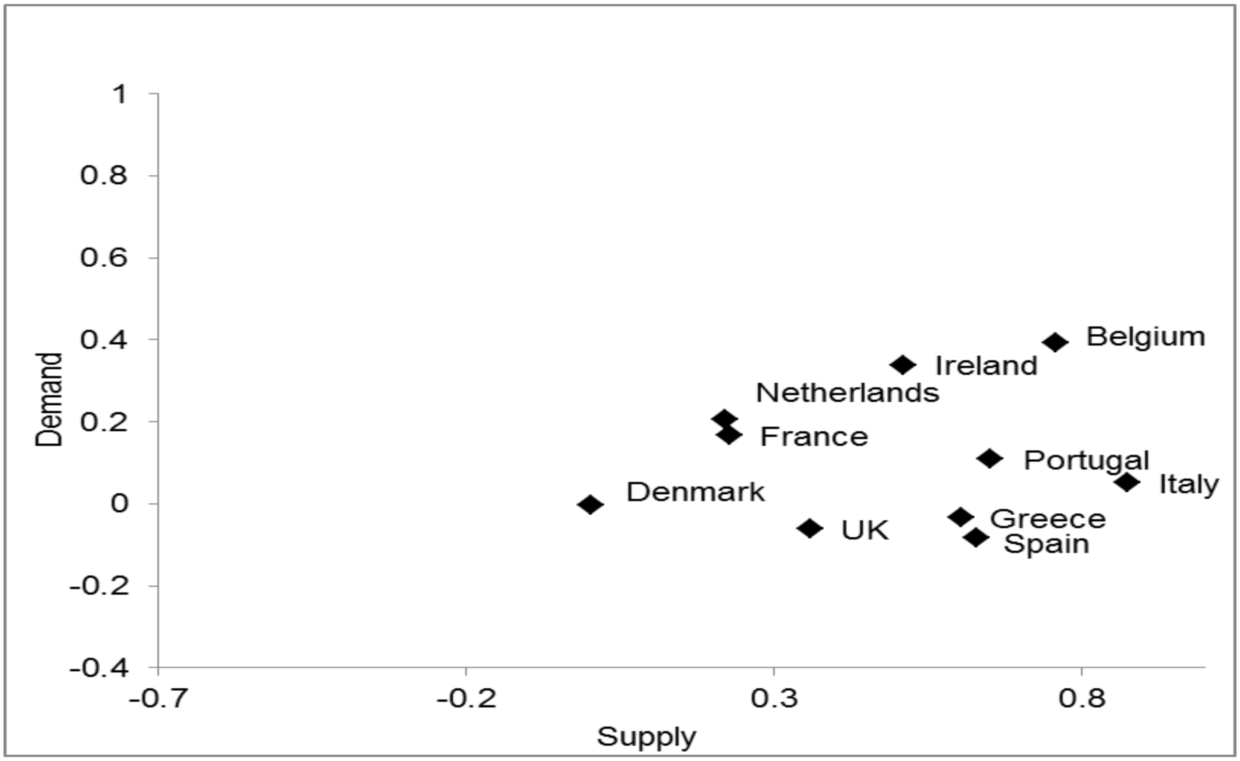
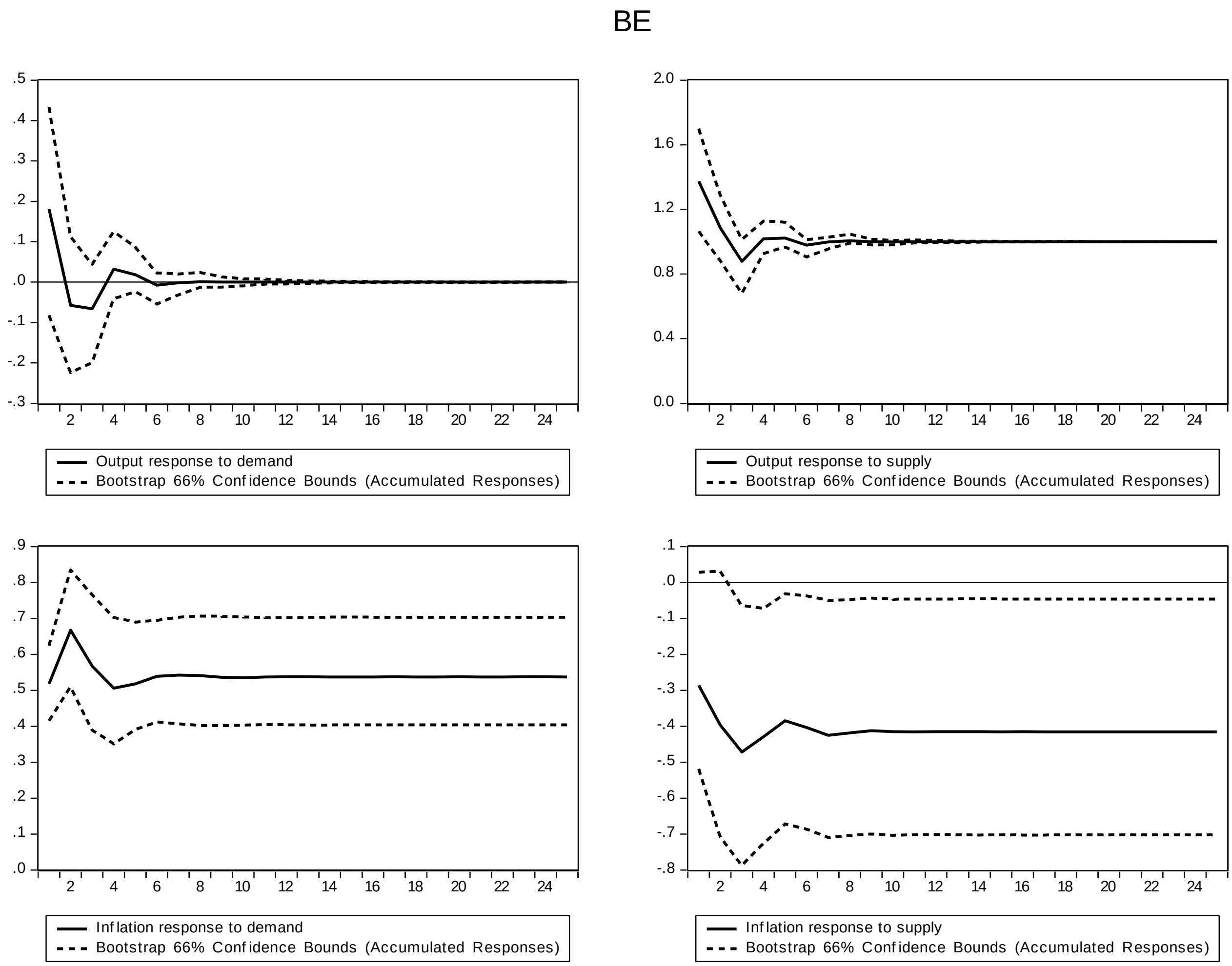
Figure 1A – Correlation of supply and demand disturbances *vis-à-vis* Germany imposing the over-identifying restriction



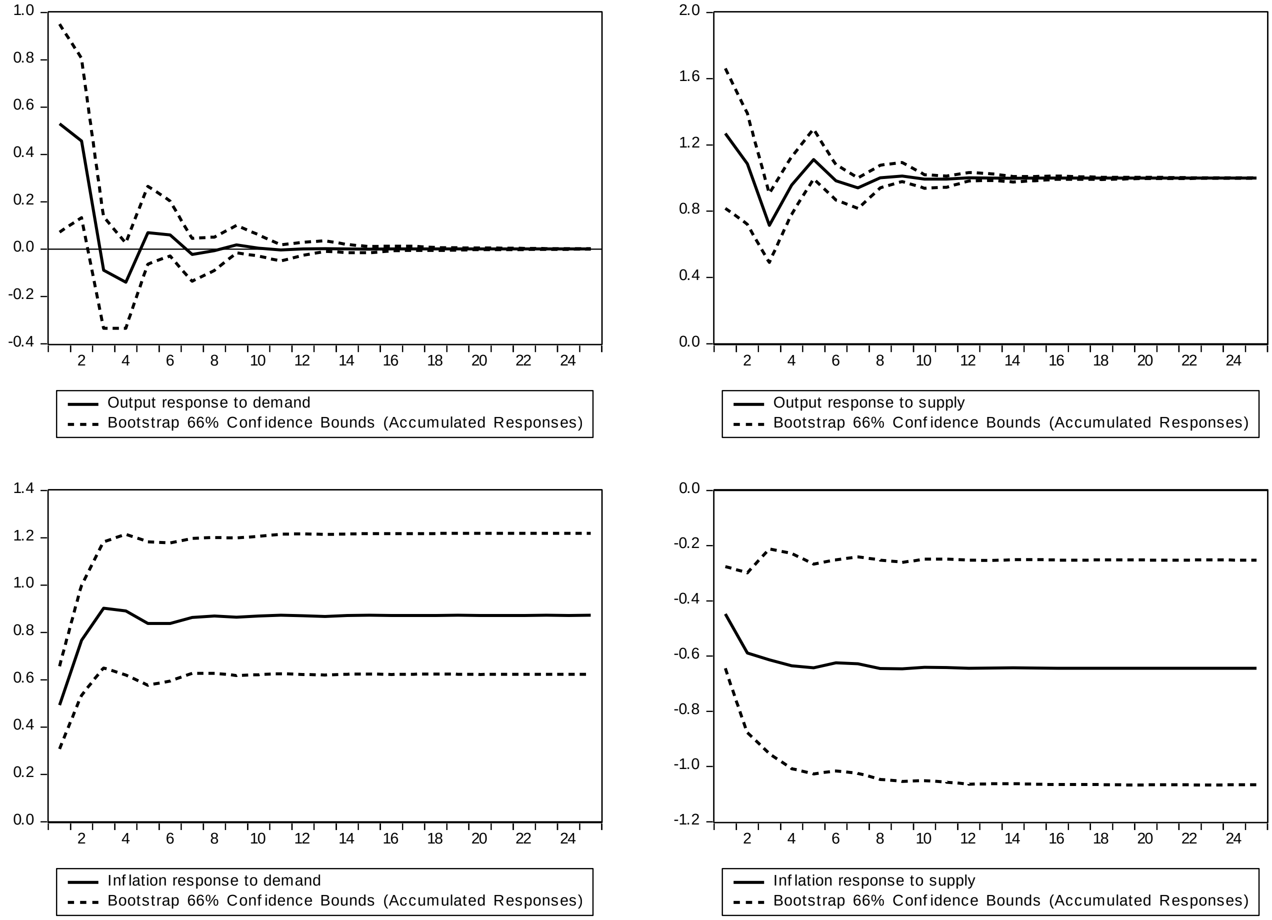
Note: Structural disturbances are retrieved from a SVAR where the over-identifying restriction described in Section 3 is imposed for all countries. The sample is 1989 – 2015, with the SVAR being solved using 2 lags for all countries and no constant as in Bayoumi and Eichengreen (1993). Comparisons of Figure 1 and 1A shows that there are no substantial differences in the results, whether residuals are bootstrapped or not.

Figure 2A – SVAR Impulse Response Functions (cumulated) relaxing the over-identifying restriction for Ireland, Spain, Greece and Portugal



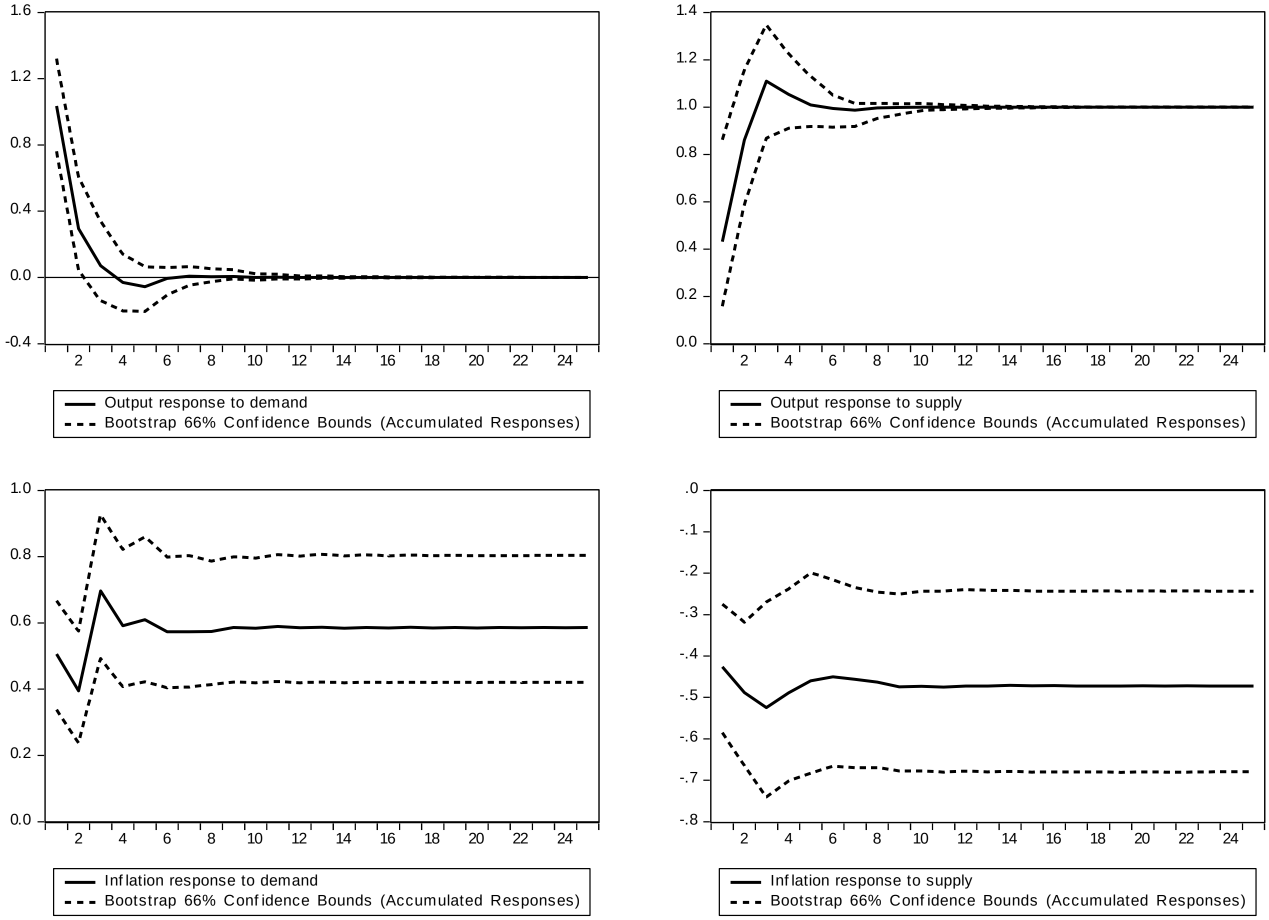
Note: IRFs report based on 10.000 VAR replications. The black line denotes the median IRF, whereas the dotted lines denote its 66% confidence interval. Structural residuals are retrieved from a SVAR where the over-identifying restriction described in Section 3 is imposed for all countries, with the exception of Ireland, Spain, Greece and Portugal. The sample is 1989 – 2015, with the SVAR being solved using 2 lags and no constant as in Bayoumi and Eichengreen (1993).

DE



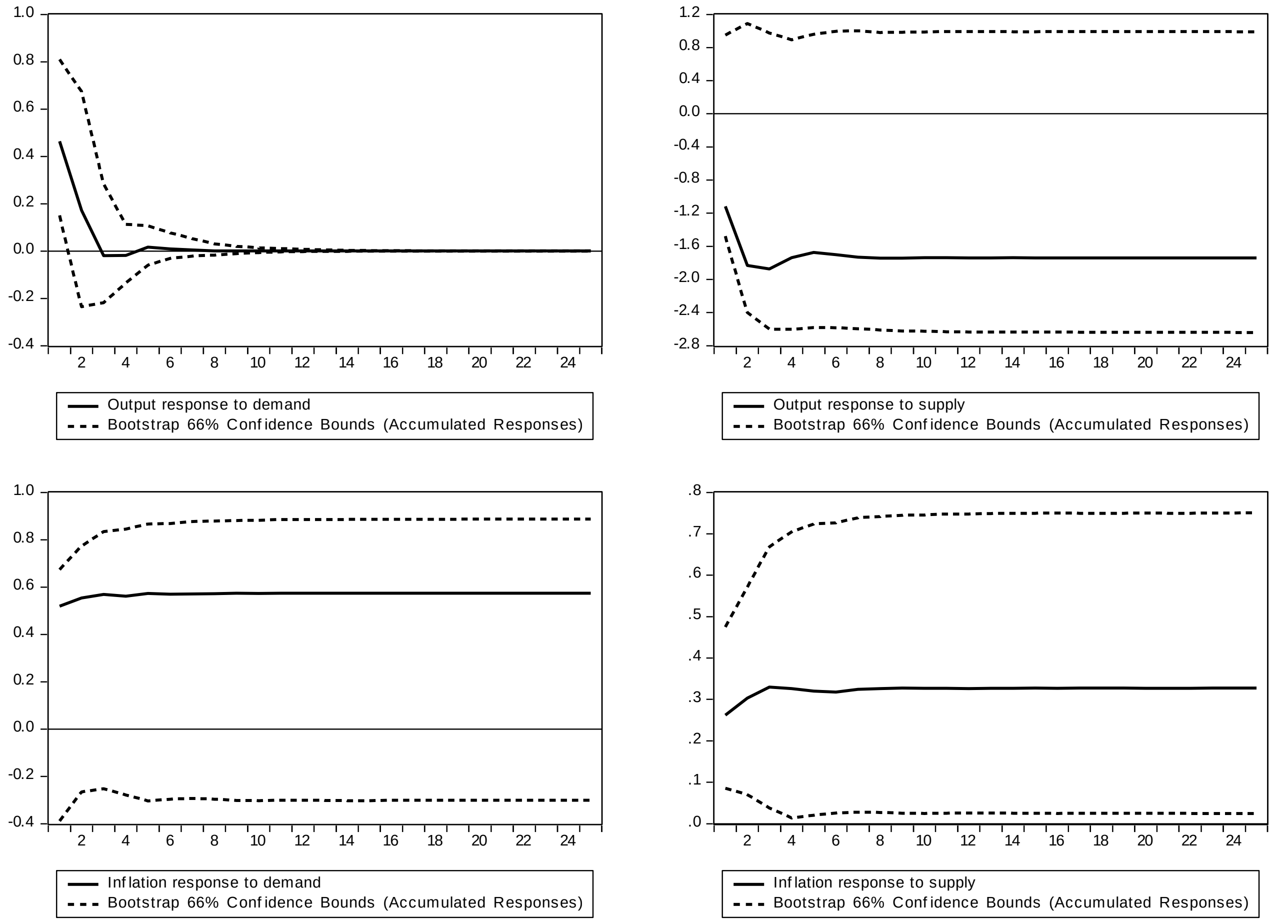
Note: IRFs report based on 10.000 VAR replications. The black line denotes the median IRF, whereas the dotted lines denote its 66% confidence interval. Structural residuals are retrieved from a SVAR where the over-identifying restriction described in Section 3 is imposed for all countries, with the exception of Ireland, Spain, Greece and Portugal. The sample is 1989 – 2015, with the SVAR being solved using 2 lags for all countries and no constant as in Bayoumi and Eichengreen (1993).

DK



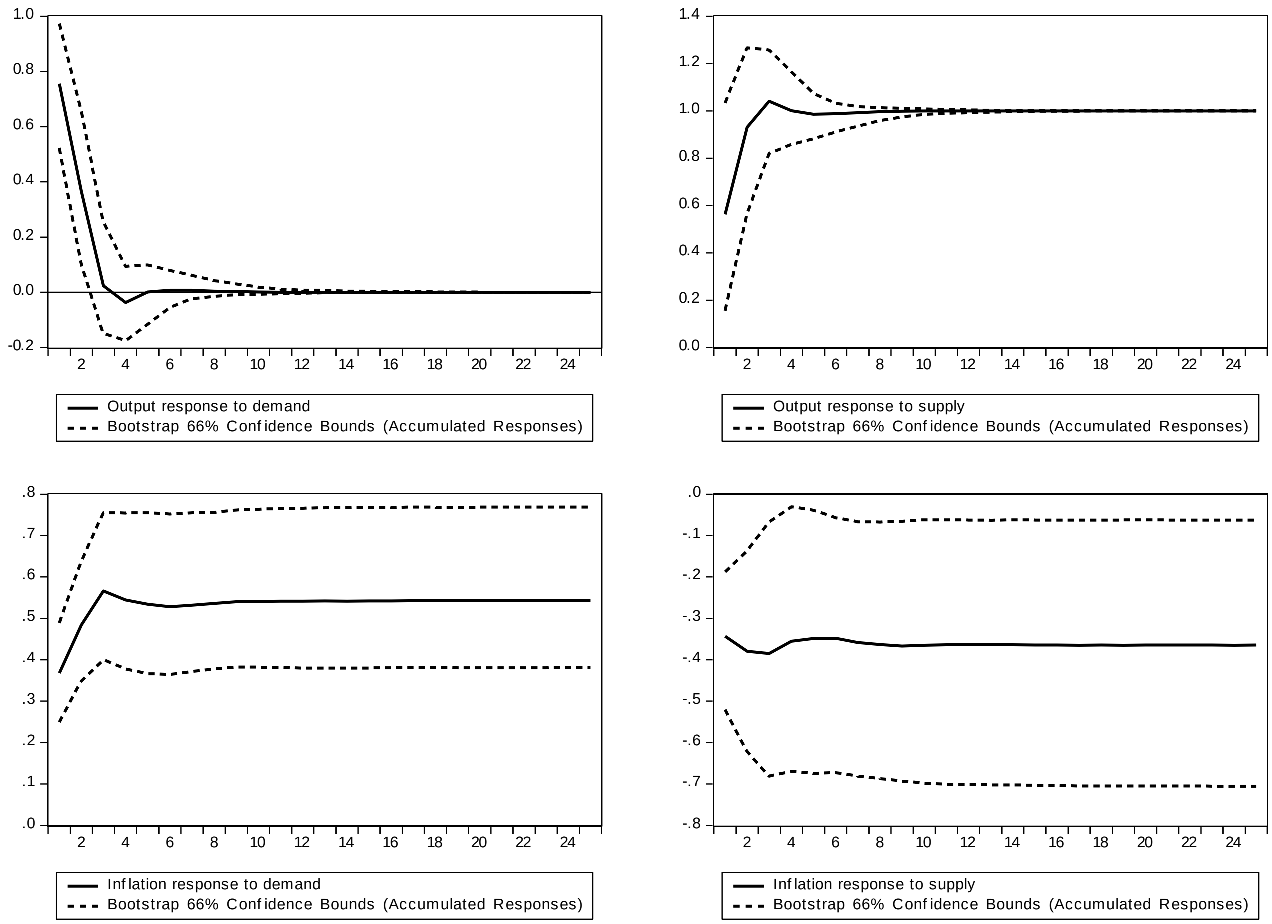
Note: IRFs report based on 10.000 VAR replications. The black line denotes the median IRF, whereas the dotted lines denote its 66% confidence interval. Structural residuals are retrieved from a SVAR where the over-identifying restriction described in Section 3 is imposed for all countries, with the exception of Ireland, Spain, Greece and Portugal. The sample is 1989 – 2015, with the SVAR being solved using 2 and no constant as in Bayoumi and Eichengreen (1993).

ES



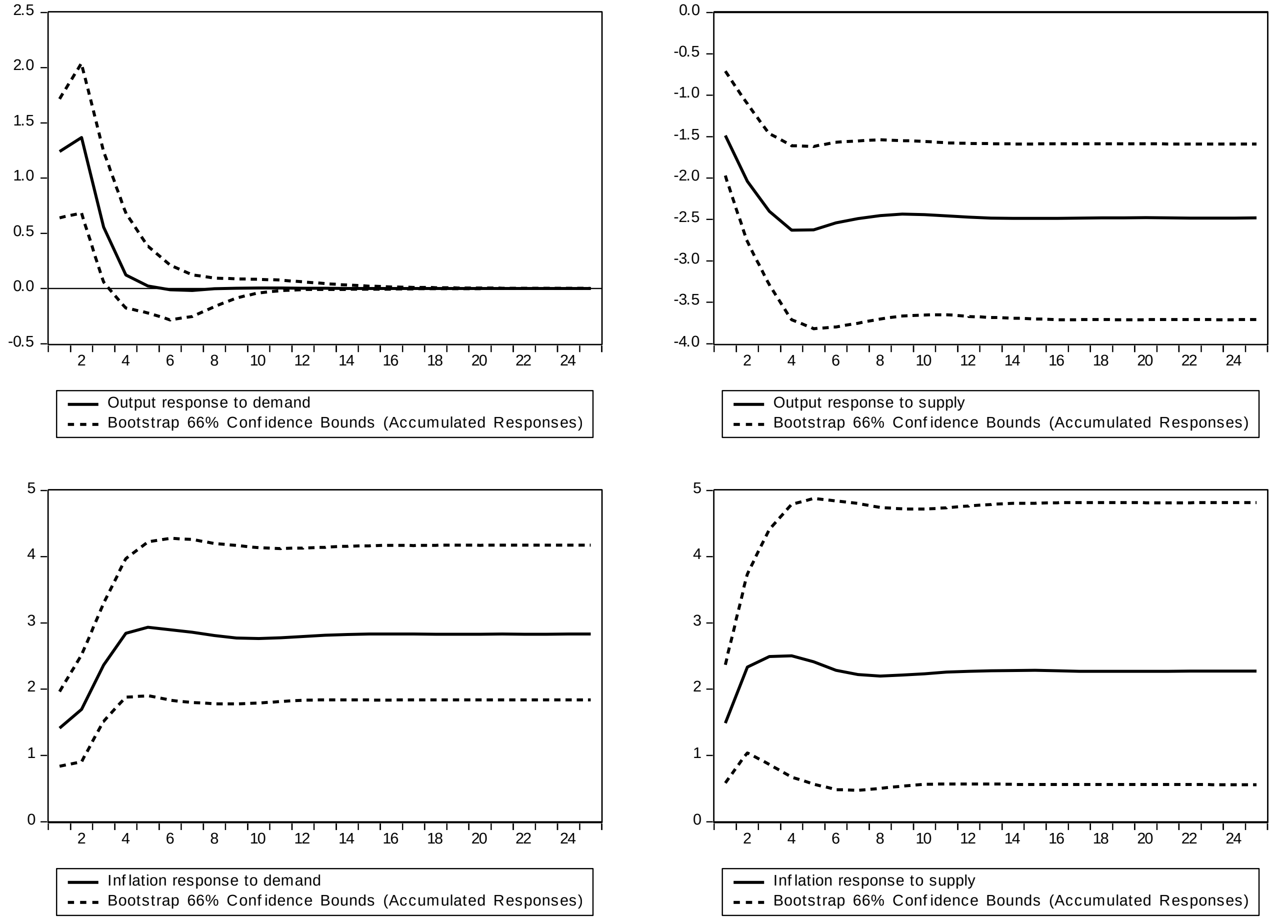
Note: IRFs report based on 10.000 VAR replications. The black line denotes the median IRF, whereas the dotted lines denote its 66% confidence interval. Structural residuals are retrieved from a SVAR where the over-identifying restriction described in Section 3 is imposed for all countries, with the exception of Ireland, Spain, Greece and Portugal. The sample is 1989 – 2015, with the SVAR being solved using 2 lags for all countries and no constant as in Bayoumi and Eichengreen (1993).

FR



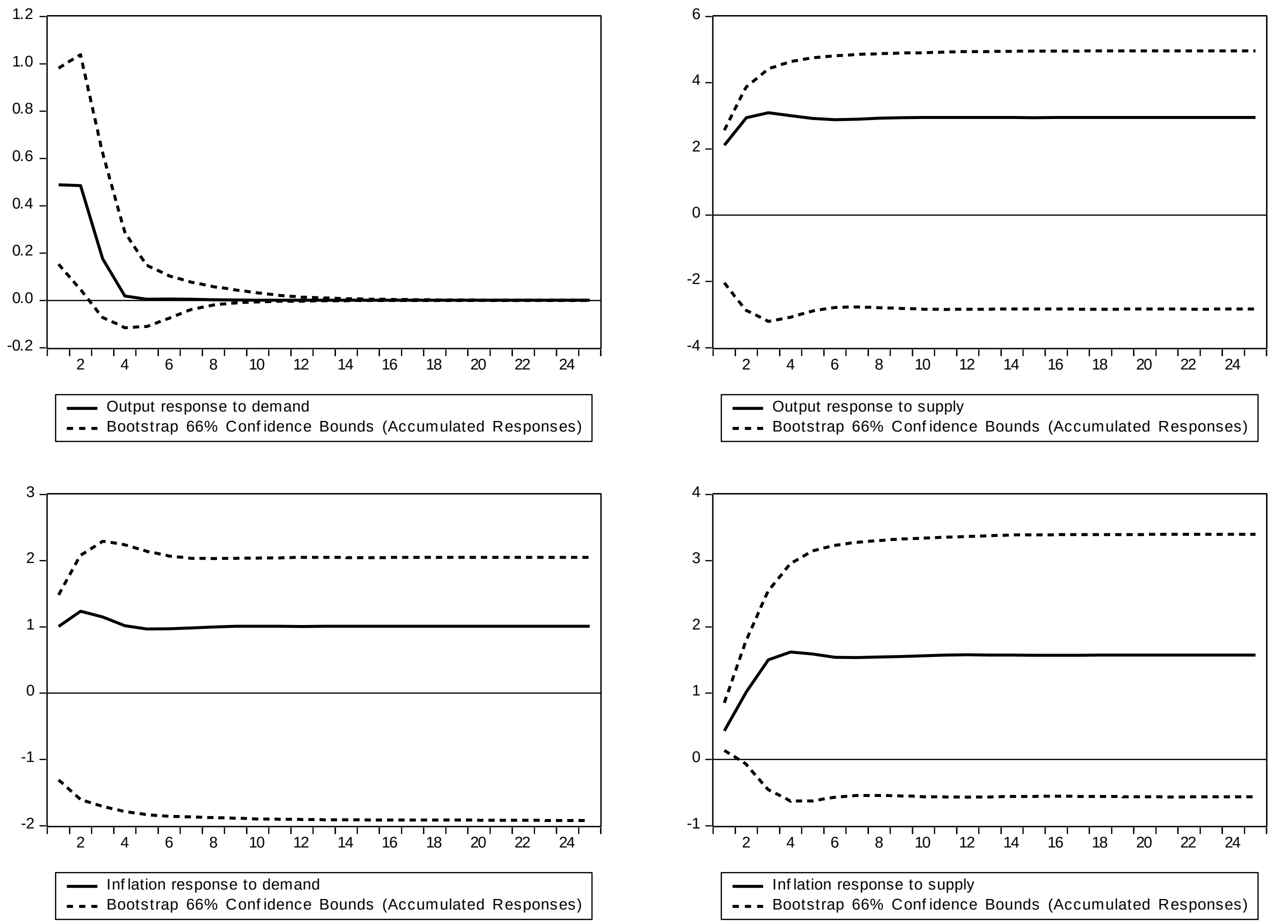
Note: IRFs report based on 10.000 VAR replications. The black line denotes the median IRF, whereas the dotted lines denote its 66% confidence interval. Structural residuals are retrieved from a SVAR where the over-identifying restriction described in Section 3 is imposed for all countries, with the exception of Ireland, Spain, Greece and Portugal. The sample is 1989 – 2015, with the SVAR being solved using 2 lags for all countries and no constant as in Bayoumi and Eichengreen (1993).

GR



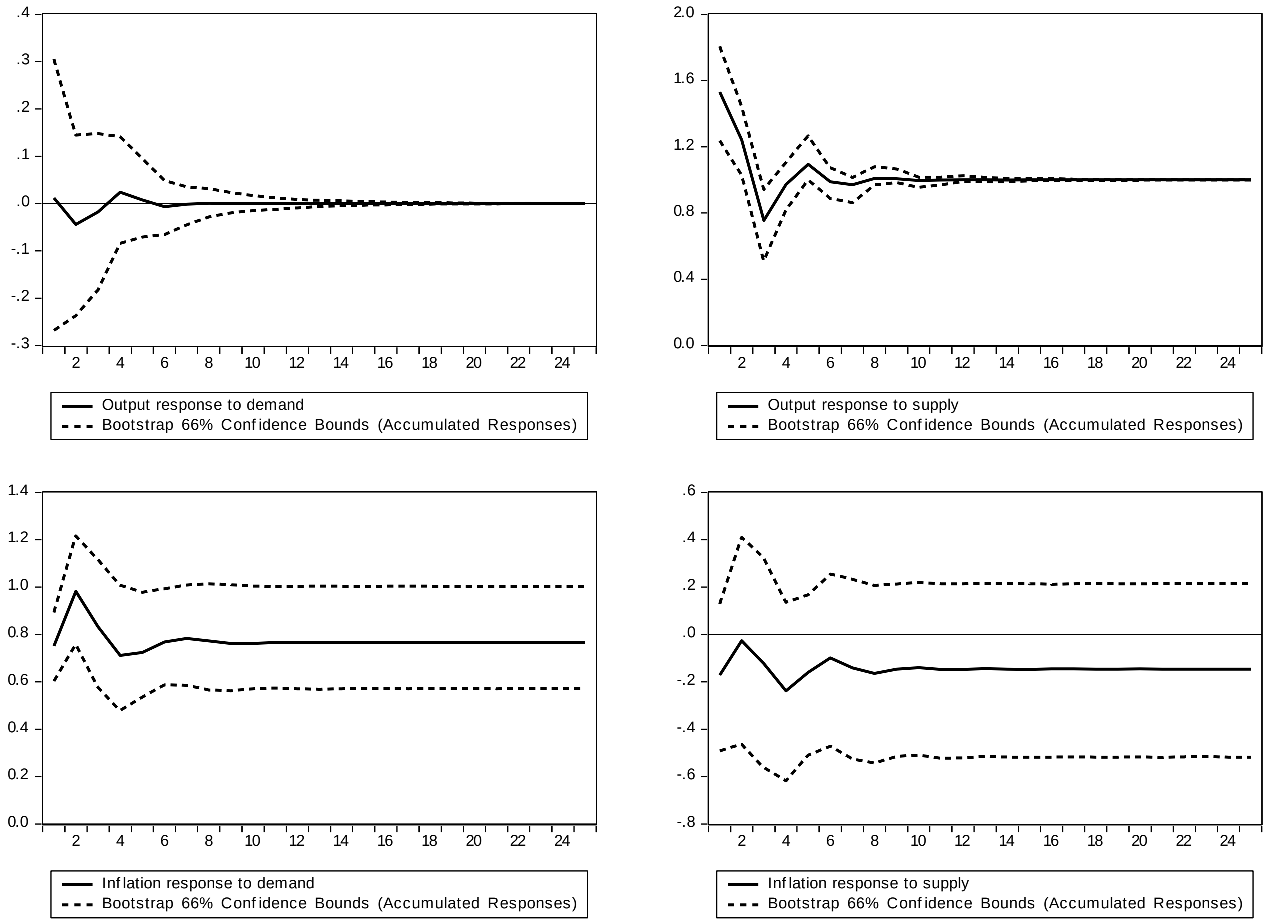
Note: IRFs report based on 10.000 VAR replications. The black line denotes the median IRF, whereas the dotted lines denote its 66% confidence interval. Structural residuals are retrieved from a SVAR where the over-identifying restriction described in Section 3 is imposed for all countries, with the exception of Ireland, Spain, Greece and Portugal. The sample is 1989 – 2015, with the SVAR being solved using 2 lags for all countries and no constant as in Bayoumi and Eichengreen (1993).

IE



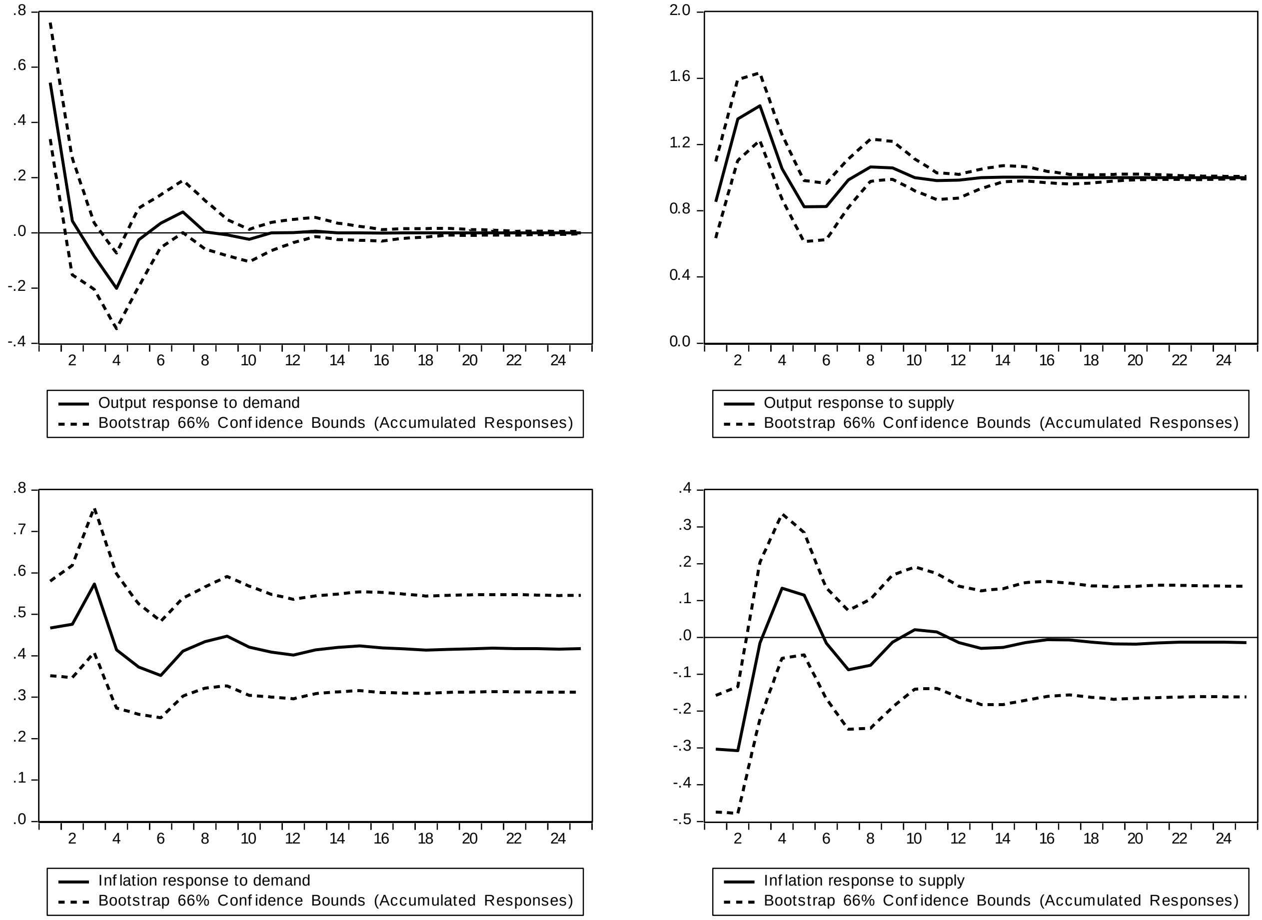
Note: IRFs report based on 10.000 VAR replications. The black line denotes the median IRF, whereas the dotted lines denote its 66% confidence interval. Structural residuals are retrieved from a SVAR where the over-identifying restriction described in Section 3 is imposed for all countries, with the exception of Ireland, Spain, Greece and Portugal. The sample is 1989 – 2015, with the SVAR being solved using 2 lags for all countries and no constant as in Bayoumi and Eichengreen (1993).

IT



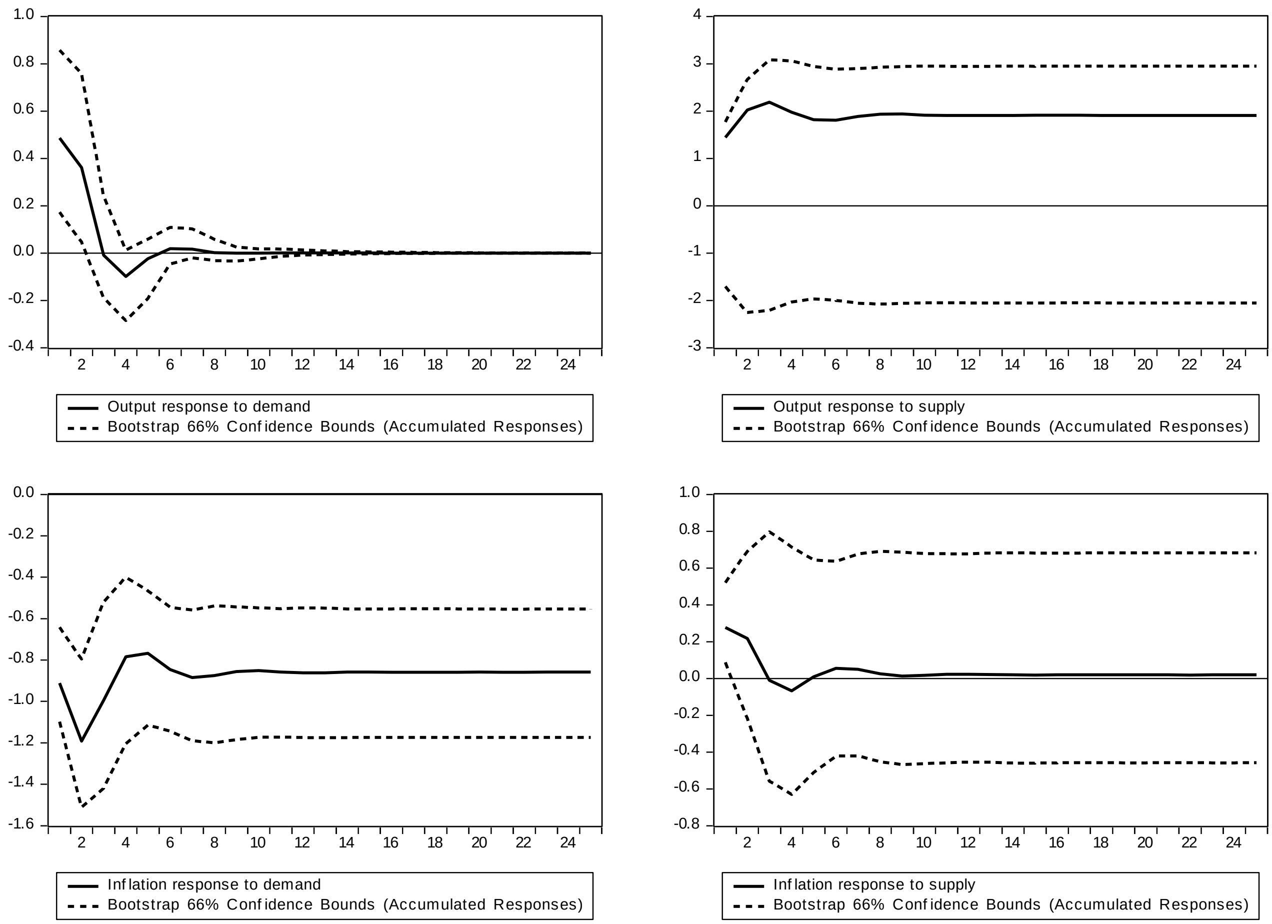
Note: IRFs report based on 10.000 VAR replications. The black line denotes the median IRF, whereas the dotted lines denote its 66% confidence interval. Structural residuals are retrieved from a SVAR where the over-identifying restriction described in Section 3 is imposed for all countries, with the exception of Ireland, Spain, Greece and Portugal. The sample is 1989 – 2015, with the SVAR being solved using 2 lags for all countries and no constant as in Bayoumi and Eichengreen (1993).

NL



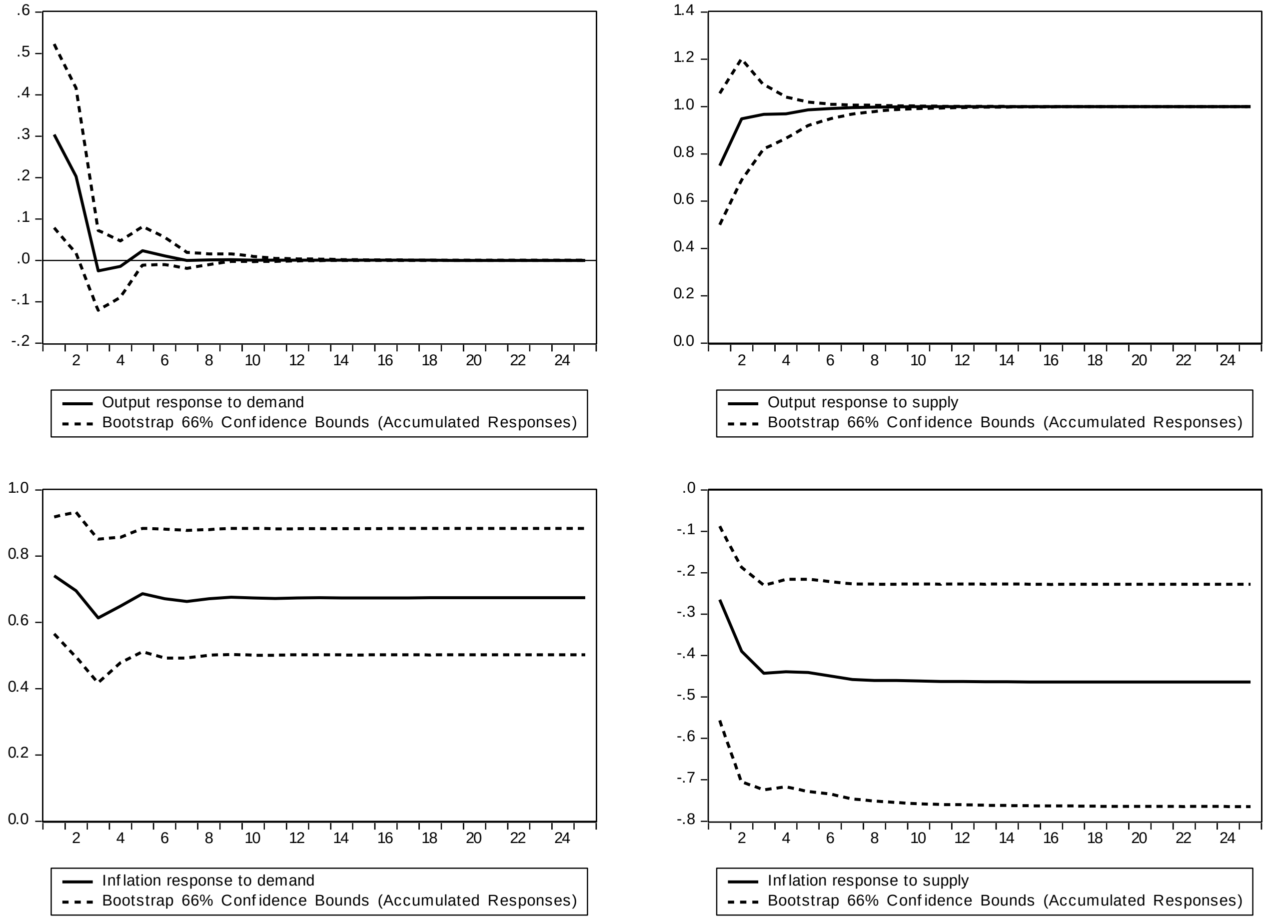
Note: IRFs report based on 10.000 VAR replications. The black line denotes the median IRF, whereas the dotted lines denote its 66% confidence interval. Structural residuals are retrieved from a SVAR where the over-identifying restriction described in Section 3 is imposed for all countries, with the exception of Ireland, Spain, Greece and Portugal. The sample is 1989 – 2015, with the SVAR being solved using 2 lags for all countries and no constant as in Bayoumi and Eichengreen (1993.

PT



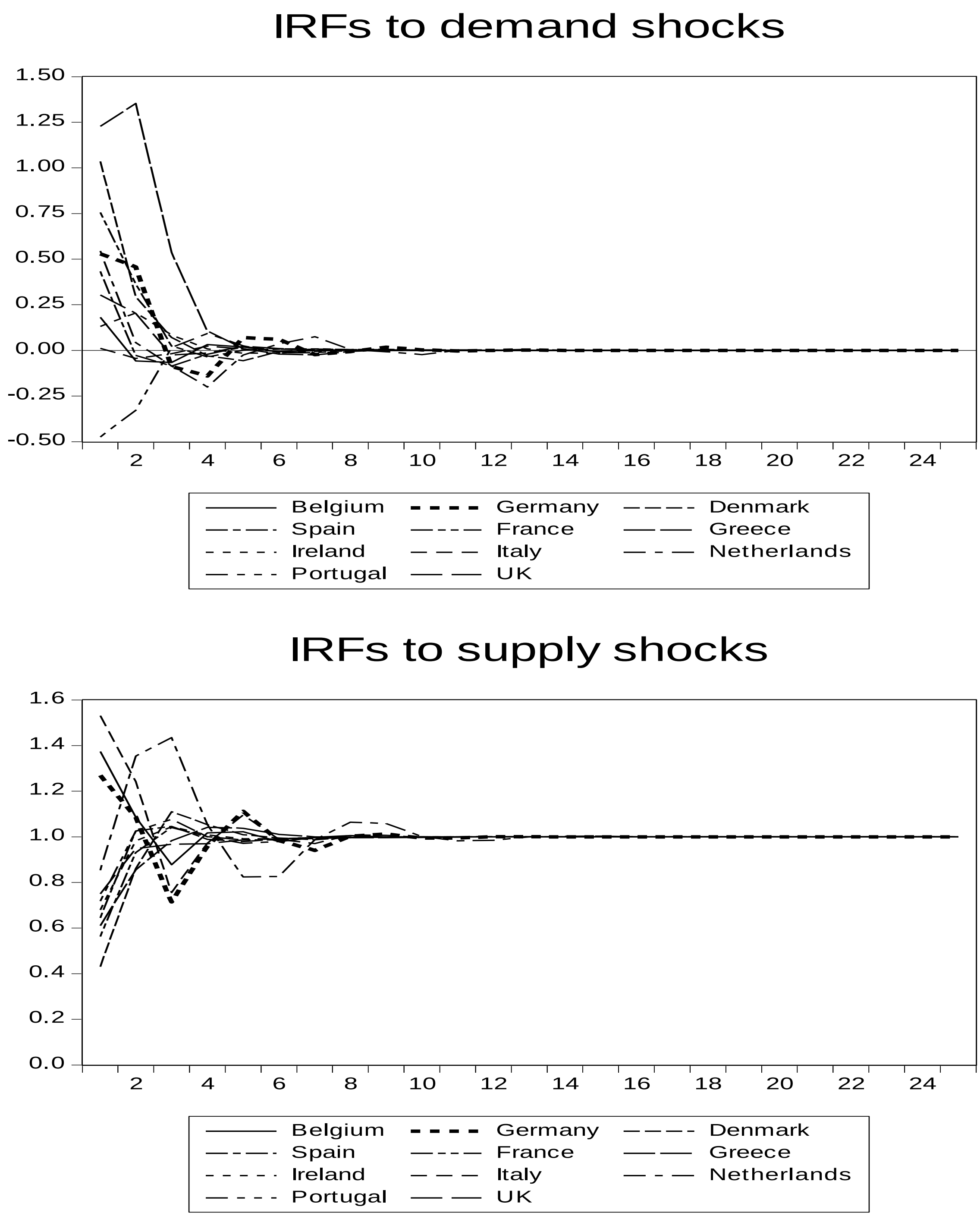
Note: IRFs report based on 10.000 VAR replications. The black line denotes the median IRF, whereas the dotted lines denote its 66% confidence interval. Structural residuals are retrieved from a SVAR where the over-identifying restriction described in Section 3 is imposed for all countries, with the exception of Ireland, Spain, Greece and Portugal. The sample is 1989 – 2015, with the SVAR being solved using 2 lags for all countries and no constant as in Bayoumi and Eichengreen (1993).

UK



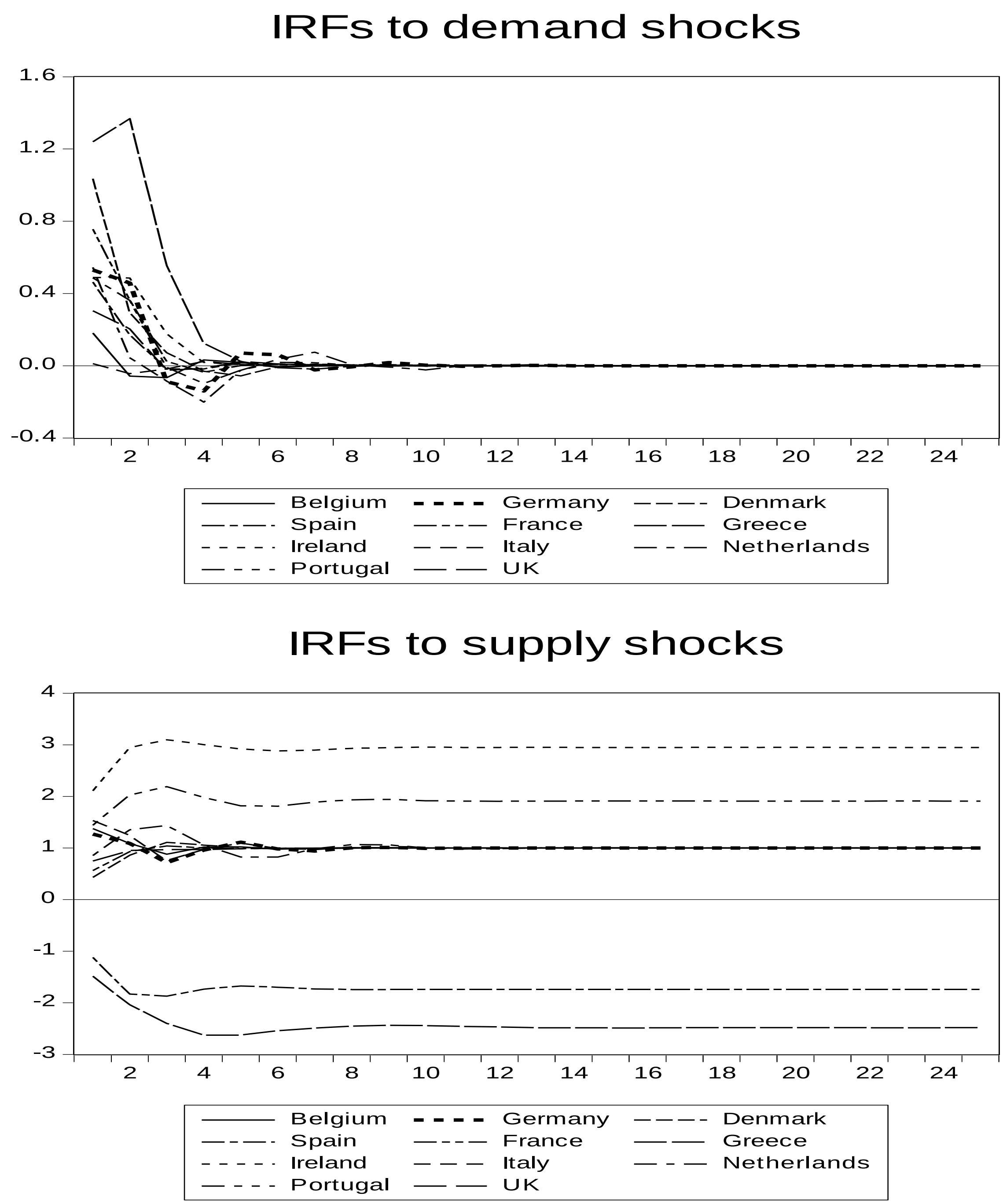
Note: IRFs report based on 10.000 VAR replications. The black line denotes the median IRF, whereas the dotted lines denote its 66% confidence interval. Structural residuals are retrieved from a SVAR where the over-identifying restriction described in Section 3 is imposed for all countries, with the exception of Ireland, Spain, Greece and Portugal. The sample is 1989 – 2015, with the SVAR being solved using 2 lags for all countries and no constant as in Bayoumi and Eichengreen (1993).

Figure 3A – SVAR IR Functions (demand) imposing the over-identifying restriction



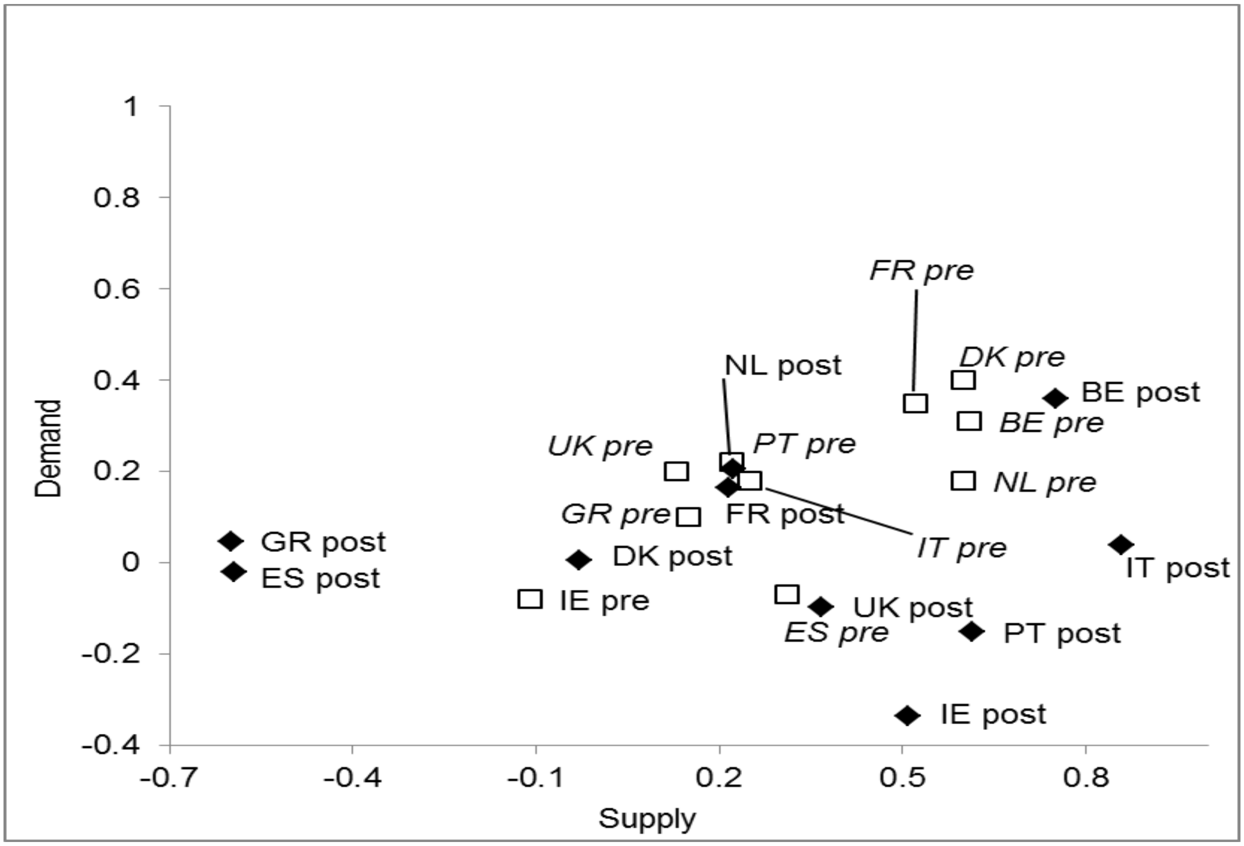
Note: IRFs report median values based on 10.000 VAR replications.

Figure 4A – SVAR IR Functions (demand) relaxing the over-identifying restriction for Ireland, Spain, Greece and Portugal



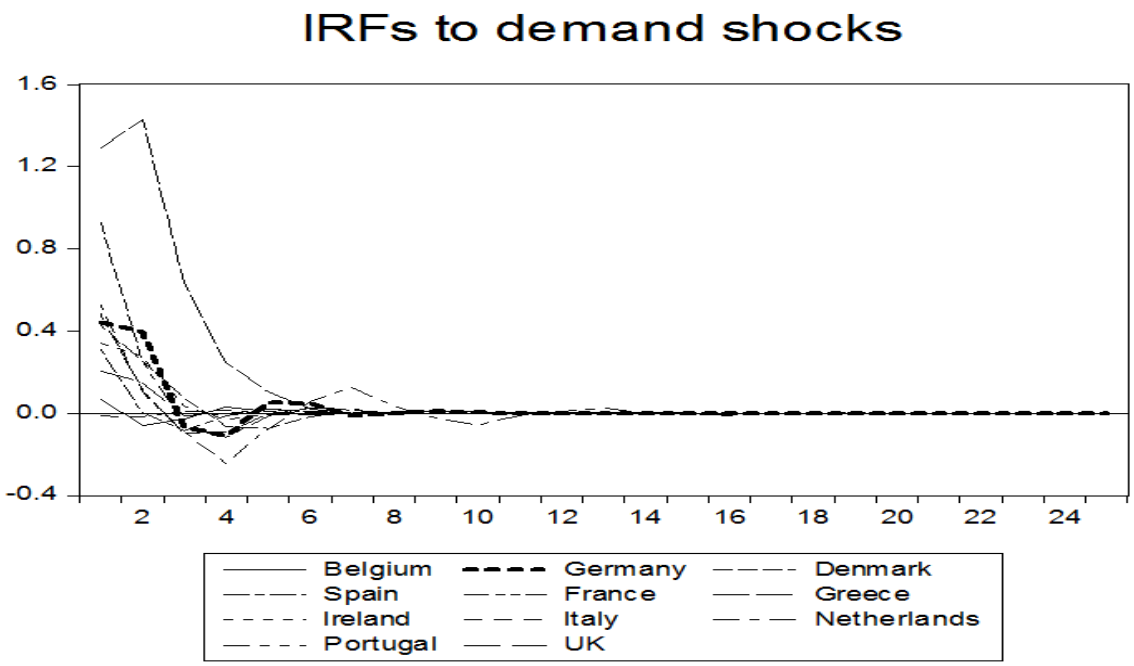
Note: IRFs report median values based on 10.000 replications.

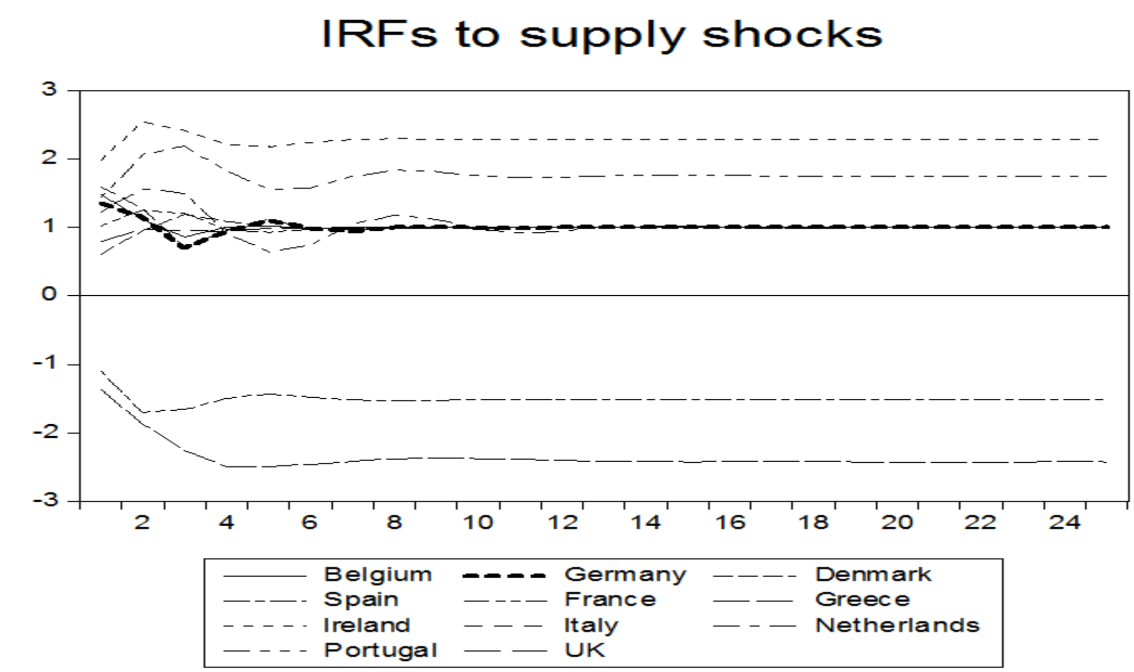
Figure 5A – Correlation of supply and demand disturbances *vis-à-vis* Germany, pre and post euro introduction



Note: The figure compares estimates from pre-Maastricht based on Bayoumi and Eichengreen (1993), covering the period 1963-1988, with our equivalent estimates for the period 1989-2015 ('post'). For each country, we estimate a bi-variate SVAR using (log) real GDP and the (log) deflator, both in first differences. The structural identification of the shocks for our sample relaxes the over-identifying restriction for Ireland, Spain, Greece and Portugal.

Figure 6A – SVAR IR Functions (demand) relaxing the over-identifying restriction for Ireland, Spain, Greece and Portugal and conditional on NOPI





Note: IRFs report median values based on 10.000 replications. The SVAR is conditional on NOPI dummies.