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The Macroeconomic Impact of the Euro

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Abstract

This paper examines whether the establishment of the euro caused structural breaks in the main macroeconomic relationships of member countries. It compares eight original members of the common currency with four European countries that did not join. The analysis constructs counterfactuals using both single equation models and a six equation vector autoregression with foreign exogenous variables, VARX*, explaining output, inflation, equity prices, exchange rates and short and long interest rates. It considers which equations changed the most and the most likely dates for any structural break.

Keywords: euro; structural-breaks; GVAR.

JEL classification: C5; E5; F4.

1. INTRODUCTION

This paper attempts to examine the effect of the establishment of the euro on eight economies that joined the common currency in comparison to four European economies that did not join the euro. There are very many ways to measure the effect of the euro and the most common way is to compare the outcome with some counterfactual, though constructing macroeconomic counterfactuals raises various difficult issues. Here we will use a very specific measure of the effect. We examine the extent to which joining the euro changed the main national macroeconomic relationships, that is whether there was a structural break in particular equations. We will ask: whether there was a significant break in 1999 for the economies that joined the euro and whether any break was bigger for those countries that joined than for countries that did not join. We also consider: in which equations the largest break occurred and whether the break was in 1999, or at another time.

The answers to these questions will always be conditional on other influences that we control for. We always control for foreign variables, the 2008 crisis, for instance, was global. Of course, to the extent that the foreign variables were also influenced by the

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formation of the euro we do not pick up that indirect effect. We also consider controlling for policy instruments, interest rates and exchange rates, which, once the common currency was established, might be regarded as exogenous to many, if not all, individual euro countries. There is also an issue as to whether to use a simple single equation approach of the sort discussed in Pesaran and Smith (2016) or a full system of the sort discussed in Pesaran and Smith (2018). Simple single equation models are parsimonious, there is a lot of evidence that parsimonious models forecast better and counterfactuals are conditional forecasts. On the other hand the systems allow for more feedbacks. We will use both.

Section 2 reviews the literature. Section 3 describes the data and provides some descriptive statistics and history. Section 4 uses single equation models and examines the timing of structural breaks. Section 5 describes the econometric approach adopted for the systems analysis. Section 6 provides results for systems. Section 7 contains some concluding comments.

2. LITERATURE

There is a large literature on the euro which considers such issues as whether the euro is an optimal currency area and the extent to which its members are subject to symmetric shocks. De Grauwe (2018) provides a text-book treatment of the economics of the monetary union. While a single currency means equality of nominal interest rates and exchange rates, it does not mean real equality. Real interest rates and real exchange rates diverged substantially among the euro economies. In addition, the differences in size, factor endowment productivity and political environment of the euro economies mean that a “one size fits all” setting of monetary policy is unlikely to be optimal for all members. The interest rate chosen by the ECB may be too low for booming countries and too high for those in recession.

Indeed, a common currency does not translate into equality among EMU members. Creditor countries tend to have more power and try to impose tight fiscal disciplines on debtor countries, De Grauwe (2016). During the Euro-crisis at the end of the 2000s, this pushed the Southern countries that were already suffering from liquidity crisis into a deeper recession, as they were forced to reduce wages and price level relative to the creditor members of the union. Greece is not one of the countries in our sample, but Alogoskoufis (2019) examines the economic history of Greece over the past 40 years and concludes that despite the significant constraints implied by membership, there are bigger risks in leaving the euro area.

Aksoy *et al.* (2002) noted that the determinants of the optimal interest rate are country-specific and this may raise tension within Euroland when it comes to choosing the optimal monetary policy for the system. Following Rudebusch and Svensson (1999) they developed a model which approximates an ECB optimal linear feedback decision rule. They found that the spread between nationally desired interest rate and the one decided by the ECB is wider for smaller countries in the union. Moreover, this difference tends to be larger when countries that desire to stabilize their output find it impossible to choose their optimal interest rate and hence left in frustration unable to achieve their objective. Nonetheless, the authors claimed that having the ECB to choose the optimal interest rate for the union based on economic conditions of all the countries in the system improves welfare by reducing losses from the volatility of inflation and variability of output. These tend to be higher when the choice of the interest rate is based on the nationalistic objectives.

There are also issues about the implications of a common monetary policy for fiscal policy and whether fiscal federalism is required, Farhi and Werning (2017). Single currency members are not able to issue debt in a national currency making them susceptible to self-fulfilling liquidity and solvency crises as investors are lacking a guarantee that cash will be available at the maturity date. This may push interest rates up and reduce liquidity available to the euro countries, De Grauwe (2013). Monetary union prevents members devaluing their currency, to remedy declining relative competitiveness. Instead, governments are forced to push their price level down by reducing the wages leading them to a deeper recession. Thus the economy's ability to defend itself against asymmetric shocks is tightly connected to the flexibility of its wages and price level.

De Grauwe and Ji (2013) provide evidence for the fragility hypothesis. They analysed the government bond markets of the EMU countries and used a control group of 14 'stand-alone' developed countries for comparison. 'Stand-alone' countries demonstrated much higher ability to sustain their sovereign debts as increase in their debt to GDP ratios was not perceived by financial market participants as a sign of increased fragility. As a result these countries overcame 2010-11 crisis without a noticeable increase in the interest rate spread. In contrast, the euro countries experienced a break in the spreads-debt to GDP ratio in 2010-11 as their financial vulnerability increased substantially when they accumulated public debt during times of financial distress. Thus, EMU countries are more prone to self-fulfilling liquidity crises.

Saka *et al.* (2015) provided further empirical support for the fragility hypothesis. Using a capital asset pricing model, they found evidence for the herding contagion that was effectively countered by the timely and reassuring ECB announcements that helped to reduce investors fear of losses and effectively addressed self-fulfilling nature of the crisis during late 2000s. Nevertheless, the analysis relies on the influence of Spain-specific news on market participants and might have been different if the estimations were based on the news of other EMU members.

Potjagailo (2017) noted that monetary shocks generated within the Eurozone tend to spillover on the 'stand-alone' European countries. The size of the spillover effects on each individual country depends on country-specific characteristics, such as its openness to trade. However, in case of majority of countries under consideration, financial variables, including short interest rates and real effective exchange rates, are significantly affected by the shocks originated in the Eurozone. As for common currency unions, when member economies are pushed 'out of sync' by shocks that are permanent, not only the nature, De Grauwe (2016), but also direction and extent of these shocks matter when estimating their effect on the flexibility-symmetry trade-off (Campos and Macchiarelli, 2018).

Those are mainly macroeconomic issues. The benefits of a monetary union may lie at the microeconomic level, De Grauwe (2018). A single currency aids free movement of goods and capital inside the EMU and reduced uncertainty about expected exchange rate. Hence, an increased economic efficiency for all the members in the monetary union.

Campos and Macchiarelli (2018) argue that creation of EMU improved the stability of the Euro Area. They distinguish between core, deep-rooted periphery and a mixed set of countries. The authors found that an economy is more likely to be a core country if it was a euro member and had strict product market regulations in place. The euro countries benefit from reduced transaction costs and uncertainty about exchange rate, as well as from price transparency, higher trade and competition. However, to fully enjoy the perks of being an EMU member, an economy needs to be able to achieve a minimum combination of symmetry, flexibility and openness.

There has been considerable dispute about the effect of monetary union on trade and Rose (2017) analyses the factors that cause the estimates to vary so much. It appears that estimates are sensitive to the sample size with the effects of a single currency on trade being stronger if the analysis includes more observations by country and time. The number of countries included seems to make especially substantial difference, which, as Rose suggested, might be explained by the bias in the estimation of the country-time fixed effects that arises if some smaller economies are omitted from the sample.

Besides development of the monetary union, various global disturbances could have affected welfare of the Euro members over the period covered in this paper. The Euro countries were not exempt from the harsh effects of the global crisis 2007-08. Caruso *et al.* (2018) analysed the effects of the financial crisis 2007-08 on the Euro countries. Using a multivariate VAR model they performed conditional and unconditional forecasts in order to examine a special nature of a financial crisis, that they argued, is different from a regular recession. They found major deviations in output, private and public debt ratios and other macroeconomic and financial indicators of the Euro countries, when the model was estimated over pre- and post-crisis periods. Some of these deviations, such as persistent decline in investment, are atypical to the extent that they, as the authors argued, cannot be explained by the business cycle regularities. The crisis was also characterised by the record high fiscal deficit-GDP ratios followed by an adaptation of extremely tight fiscal policies, all of which make this crisis unprecedented and likely to cause a structural change in the macroeconomic indicators of some Euro countries.

In studies that are more closely related to what we are doing Pesaran *et al.* (2007) estimate what would have happened to the UK and Sweden and the euro area if the UK or Sweden had joined the euro using a Global Vector Autoregression, GVAR. Smith (2009) examines whether the establishment of the euro caused a structural break. In both cases there was a relatively short sample of data after the establishment of the euro and the samples ended prior to the financial crisis. In both cases the effects were not large. Now there is more data and the large shocks associated with the financial crisis provides extra identifying information.

3. DATA

We use data from the GVAR toolbox 1979Q3-2016Q4, Mohaddes and Raissi (2018), to estimate VARX* for a number of European countries that did or did not join the euro.

For countries *i*=1,2,..,12 and *t*=1979Q2−2016Q4, the variables are:

- y_{it} natural logarithm of real GDP volume index
- Δp_{it} the rate of inflation, calculated by taking the difference of the natural logarithm of the consumer price index
- eq_{it} natural logarithm of the nominal equity price index deflated by CPI
- ep_{it} natural logarithm of the exchange rate of country *i* at time *t* expressed in units of foreign currency per US dollar deflated by country *i*'s CPI. We will refer to this as a real exchange rate, even though it is not adjusted for the US price level.
- r_{it} nominal short-term interest rate per quarter, in per cent; computed as $0.25 \times \ln(1 + R_{it}^{r})\,/100$ where R_{it}^{r} is the nominal short rate of interest per annum in percent.
- l_{it} nominal long-term interest rate per quarter, in per cent; computed as $0.25 \times \ln\left(1 + R_{it}^{lr}\right)/100$ where R_{it}^{lr} is the nominal long rate of interest per annum in percent.
- $poil_{it}$ natural logarithm of the nominal price of oil in US dollars

The euro-member countries we consider are: Austria, Belgium, Finland, France, Germany, Italy, Netherlands, Spain. The non-euro members are: Norway, Sweden, Switzerland, UK. The long interest rate is not available for Finland.

In addition to individual variables, say x_{it}, there are global equivalents x^*_{it} calculated as country specific trade weighted averages of the corresponding variables of the other countries:

$$x^*_{it} = \sum_{j=0}^{N} w_{ij} x_{jt}, \text{ with } w_{ii} = 0,$$

where w_{ij} is the share of country j in the trade (exports plus imports) of country i. So for instance if y_{it} is log real GDP, y^*_{it} is the weighted average of the log GDP of trading partners.

For each country, we break the data into three periods, as follows 1:1979Q4-1998Q4; 2:1999Q1-2008Q4; 3: 2009Q1-2016Q4. Table no. 1 provides means and standard deviations for the growth rate, the rate of inflation, the change in the real exchange rate, and the long rate (excluding Finland).

Table no. 1 – Mean and standard deviation, percent per annum (x400)

	Dy			Dp			Dep			Lr		
	1	**2**	**3**	**1**	**2**	**3**	**1**	**2**	**3**	**1**	**2**	**3**
Austria	2.28	2.00	1.08	3.14	1.99	1.70	-3.70	-2.28	0.93	7.19	4.10	2.15
	3.66	3.58	4.26	2.08	1.14	1.25	20.34	20.86	15.91	1.51	0.49	1.22
Belgium	1.94	1.91	1.14	3.49	2.23	1.55	-2.56	-2.51	1.07	8.72	4.38	2.47
	3.23	2.79	2.32	2.67	1.73	1.69	20.41	20.64	15.76	2.32	0.64	1.33
Finland	2.51	2.63	0.09	4.66	1.90	1.18	-3.13	-2.20	1.45	0.00	0.00	0.00
	5.65	4.42	6.01	3.55	1.60	1.71	20.63	20.62	15.80	0.00	0.00	0.00
France	1.84	1.82	0.77	4.52	1.77	0.92	-3.01	-2.05	1.70	9.20	4.32	2.20
	1.78	2.32	1.82	4.04	1.19	1.14	19.38	20.85	15.77	2.85	0.61	1.11
Germany	1.89	1.35	1.32	2.74	1.62	1.12	-3.10	-1.91	1.51	6.88	4.11	1.65
	3.78	2.67	3.79	2.18	1.17	1.03	20.21	20.74	15.79	1.43	0.46	1.07
Italy	1.89	0.97	-0.38	7.50	2.33	1.12	-3.79	-2.62	1.51	11.72	4.48	3.63
	2.65	2.31	3.12	5.50	0.83	1.37	20.11	20.74	15.72	3.66	0.49	1.46
Netherlands	2.31	1.90	0.45	2.68	2.18	1.44	-2.92	-2.48	1.19	7.42	4.35	1.97
	3.26	2.43	2.83	2.14	1.25	1.42	20.27	20.46	15.81	1.56	0.66	1.13
Norway	3.21	2.05	1.43	5.30	2.15	1.99	-3.12	-3.06	0.60	9.55	4.76	1.94
	5.33	3.78	5.17	3.71	3.17	2.05	18.06	22.40	18.05	2.69	1.11	0.85
Spain	2.51	2.82	0.43	6.99	3.12	0.66	-2.99	-3.41	1.96	11.54	4.17	3.66
	1.88	2.20	2.48	4.07	1.67	3.28	20.00	21.11	16.11	3.15	0.61	1.54
Sweden	1.91	2.30	2.63	5.51	1.61	0.65	-2.08	-1.89	1.17	10.10	4.40	1.91
	5.64	4.50	4.11	4.37	1.63	1.52	21.21	21.98	19.38	2.12	0.71	0.97
Switzerland	1.64	1.82	1.36	2.78	1.09	-0.30	-3.61	-3.04	-1.26	4.62	2.88	0.87
	2.82	3.94	2.79	2.33	1.12	1.13	22.94	16.45	15.47	0.90	0.44	0.83
UK	2.21	2.02	1.36	5.37	2.28	2.08	-3.80	-1.51	0.74	9.39	4.59	2.52
	2.86	2.86	2.01	3.74	1.31	1.68	21.36	17.42	16.13	2.10	0.24	0.85

Notes: Dy is the GDP growth rate, Dp is the rate of inflation, Dep is the change in the exchange rate, Lr is the long interest rate. The information are given for periods: 1: 1979Q4-1998Q4; 2: 1999Q1-2008Q4; 3: 2009Q1-2016Q4.

In the three year period after 1999 most of the countries suffered from lower growth, compared to the pre-euro period. The exceptions were Finland, Spain, Sweden, and Switzerland, who enjoyed a minor increase in the GDP growth rate. Moreover, growth is lower in the post crisis period, but this is a global phenomenon not just for euro countries. We control for this by including global variables x^*_{it} in our model.

The average inflation rate dropped during the period of euro formation. For most countries the inflation rate more than halved and in some cases, such as Italy it fell by more than two thirds. As for the euro countries, under the EMU, money supply and inflation were now closely coordinated by the the European Central Bank, ECB. However, it does not explain the decrease in inflation rates in the stand-alone countries. The inflation rate was maintained at a relatively stable rate after the financial crisis.

The average exchange rate growth over the first two sub-periods was negative, suggesting appreciation against US dollar. However, following the financial crisis, the dollar strengthened against national currencies of the countries in the sample. The long interest rate declined significantly during the years of euro formation, partially as a result of the strict anti-inflationary policy adopted by the ECB, and fell further in the third period, possibly due to low expected inflation and returns on investments. Average growth rates fell over the three periods both for the 8 euro countries: 2.15, to 1.92, 0.62, and for the 4 non-euro countries: 2.24, 2.05, 1.69. While the non-euro average is higher The differences are quite small for the first two periods, but in the post crisis period the four euro countries grew faster than any of the eight euro countries. Inflation fell over the three periods in both groups.

4. SINGLE EQUATION MODEL

In the comparisons of growth rates in the descriptive statistics above we did not control for other factors, particularly the crisis, we now do that. To begin we estimate a simple single equation model that makes the log of real GDP an ARDL function of foreign GDP and trend over the pre-euro period 1979Q2-1998Q4, say $t = 1,2, \dots, T_1$:

$$y_{it} = \alpha_0 + \alpha_1 y_{i,t-1} + \beta_0 y_{it}^* + \beta_1 y_{i,t-1}^* + \gamma t + u_t \tag{1}$$

This equation is then used to forecast GDP over the following T_2 quarters, $t = T_1 + 1, T_1 + 2, \dots, T$, with $T_2 = T - T_1$. The counterfactual here is the value of output that would be predicted from pre 1999 parameters using post 1999 information on foreign variables.

We use a number of tests for structural stability. Chow's first test is for equality of the k=5 parameters between the two periods, assuming the variances are the same. It uses the unrestricted residual sum of squares of the regressions over the two periods $(\hat{u}_1{}'\hat{u}_1 + \hat{u}_2{}'\hat{u}_2)$ with degrees of freedom $T-2k$ and restricted residual sum of squares from the regression over the whole period $\hat{u}'\hat{u}$ with degrees of freedom $T-k$. The null hypothesis that the parameters are equal in the two periods implies k restrictions and the test statistic is:

$$\frac{[\hat{u}'\hat{u} - (\hat{u}_1{}'\hat{u}_1 + \hat{u}_2{}'\hat{u}_2)]/k}{(\hat{u}_1{}'\hat{u}_1 + \hat{u}_2{}'\hat{u}_2)/(T - 2k)} \sim F(k, T - 2k)$$

Chow also suggested a second predictive failure test for the hypothesis that the first period predicts the second where the test statistic is:

$$\frac{[\hat{u}'\hat{u} - \hat{u}_1{}'\hat{u}_1]/T_2}{\hat{u}_1{}'\hat{u}_1/(T_1 - 2k)} \sim F(T_2, T_1 - k)$$

This tests the hypothesis that in:

$$\begin{bmatrix} y_1 \\ y_2 \end{bmatrix} = \begin{bmatrix} X_1 & 0 \\ X_2 & I \end{bmatrix} \begin{bmatrix} \beta_1 \\ \delta \end{bmatrix} + \begin{bmatrix} u_1 \\ 0 \end{bmatrix}$$

δ, the $T_1 \times 1$ vector of forecast errors, are not significantly different from zero. This has a dummy variable for each observation in the second period.

Table no. 2 gives the Chow Predictive Failure, PF, and Structural Stability, SS, tests p-values, mean and root mean square prediction errors for a break in 1999Q1. Plots of actual and predicted values are given in an Annex. These Chow tests assume a known break-point, but there are also Quandt-Andrews tests for an unknown break point, which searches over the possible dates for a single break. The trimming percentage is set to 15%. Table no. 2 also gives the Q-A break date, all are significant.

For Austria the predicted is close to the actual with a small mean error and RMSPE. Neither the PF nor SS test rejects the hypothesis of structural stability at the 5% level.

The countries which performed less well than expected from the pre-euro relationship are Italy, Netherlands, and Norway. Their mean errors and RMSPEs are on a higher end, comparing to the rest of the sample. Moreover, the Chow SS test statistics suggests possible structural instability in growth rates of Italy and Norway at 5% level and that of Netherlands at 10% level.

Table no. 2 – Structural Break in 1999Q1: log GDP

	PF F(72, 72)	SS F(5, 139)	Mean PE	RMSPE	Q-A Test
Austria	0.49	0.05	0.01	0.02	1988Q3***
Belgium	1.00	0.00	0.05	0.06	1987Q3***
Finland	0.99	0.08	0.10	0.11	1990Q2***
France	0.94	0.00	0.04	0.04	1998Q2***
Germany	1.00	0.04	0.00	0.08	1988Q3***
Italy	0.99	0.00	-0.07	0.09	2003Q1***
Netherlands	1.00	0.09	-0.06	0.08	1986Q1***
Norway*	0.91	0.03	-0.11	0.15	1987Q3***
Spain	0.14	0.02	-0.09	0.15	2008Q2***
Sweden*	1.00	0.02	0.15	0.17	1990Q1***
Switzerland*	0.01	0.00	0.13	0.16	2008Q2***
UK*	0.68	0.00	-0.59	0.83	2008Q2***

Notes: PF: Chow Predictive failure test; SS: Chow Structural stability test; Mean PE: Mean Prediction Errors; RMSPE: Root Mean Squared Prediction Error; Q-A Test: Quandt-Andrews Test (*** indicates a significance level of 1 percent, ** of 5 percent, and * of 10 percent); *: non-member countries.

The countries which performed better than expected from the pre-euro relationship are Belgium, Finland, France, Sweden, and Switzerland. Chow SS test suggests presence of structural instability in time series of all these countries, except Finland. In case of Switzerland this conclusion is supported by the Chow PF statistics as well. Sweden and Switzerland have the highest mean prediction errors and RMSPE in the sample, excluding the UK. The UK results are unreliable because the coefficient of the lagged dependent variable was greater than one on the pre-euro period.

The countries where the evidence is mixed are Germany and Spain. For Germany the predicted growth rate was higher than the actual one and was expected to decline around 2011. However, contrary to this prediction, ever since a sharp drop during the global crisis 2007-08, the actual growth rate has been slowly increasing. For Spain the actual was close to the predicted till the global crisis, after which the actual growth rate dropped below the predicted one and continued to increase at a much slower pace than was predicted.

Overall, for majority of countries in the sample their actual and predicted growth rates are not necessarily close to each other, but exhibit similar general patters. Only in case of Spain, Germany, Switzerland, and the UK the actual and predicted income growth rates diverge substantially from each other.

In Finland the break occurred in early 1990s (1990Q2), time of Finnish banking crisis and collapse of the Soviet Union, with which Finland had strong trading ties. In case of Sweden the Q-A test identified a break in 1990Q1, which might be connected to the Swedish banking crisis that erupted in 1992. For Spain, Switzerland, and the UK, the Q-A test statistics suggests a breakpoint in 2008Q2, year of the global financial crisis. As for the rest of the countries, the identified breakpoints are not characterised by any major economic or financial events.

The Bai-Perron test which allows for multiple breakpoints was also used, but did not suggest that Euro formation was a reason behind the breaks in the data for any of the countries in the sample.

We also estimate a growth rate relationship below:

$$\Delta y_{it} = \alpha_0' + \alpha_1' \Delta y_{i,t-1} + \beta_0' \Delta y_{it}^* + \beta_1' \Delta y_{i,t-1}^* + \epsilon_t \qquad (2)$$

where Δy_{it} and Δy_{it}^* are real domestic and foreign GDP growth rates, respectively, and ϵ is the error term.

This equation is then used to forecast GDP growth rate as with the levels relationship over the following T_2 quarters, $t = T_1 + 1, T_1 + 2, \dots, T$, with $T_2 = T - T_1$.

Compared to the levels relationship, the results for the growth equation are quite different. In Austria and Belgium overall foreign income growth effect is positive, but the lagged foreign GDP is only significant at 5% level in Belgium and insignificant in Austria. In addition, coefficient of the lagged national income is negative and in case of Belgium, insignificant. The equations, however, passed all the diagnostic tests, except the Belgium equation, which failed the serial correlation test at 10% level.

In case of Finland, France, Germany, Italy, and Netherlands only the foreign income growth coefficient is significant and positive. The equation for France passed all the diagnostic tests, while the ones for Finland and Germany failed serial correlation test. As for Italy, the equation failed serial correlation and normality tests. The equation for Netherlands failed normality and functional form tests, the latter only at 10% level.

In the Norway equation foreign income growth has significant positive effect on home GDP growth. However, lagged domestic income growth has negative and significant effect on the dependent variable. This equation passed all the diagnostic tests.

In case of Spain lagged domestic income growth and foreign income growth have positive effect on domestic GDP growth, but the foreign coefficient is only significant at 10% level. However, the equation failed serial correlation test and normality test, the latter only at 10% level. In Sweden, both foreign income growth and lagged domestic GDP growth coefficients are significant, although the former is positive, while the latter is negative. The equation passed all the diagnostic tests. In case of Switzerland and the UK, the foreign income growth and lagged domestic income growth have positive and significant effect on the dependent variable and the Switzerland equation passed all the diagnostic tests, while the UK one failed those for serial correlation at 5% level, as well as functional form and normality at 10% level.

The lagged foreign income is insignificant in all countries equations with exception of Belgium where it is significant at 10% level. Nonetheless, comparing to the levels specification model, the lagged foreign income growth coefficient is positive in all equations, except Finland, France, and Norway, with overall foreign income growth effect being positive for all countries. All equations passed functional form test (Netherlands and the UK failed it, but only at 10% level) and heteroskedasticity test, however Belgium (at 10% level), Finland, Germany, Italy, Spain, and the UK failed the serial correlation test and four failed the normality test (Italy and Netherlands at 5%, while Spain and the UK at 10% level).

Table no. 3 – Structural Break Analysis (Income Growth)

	PF F(72,73)	SS F(4,141)	Mean PE Annual %	RMSPE Annual %	Q-A Test
Austria	0.49	0.75	-0.28	3.22	1988Q2
Belgium	1.00	0.59	0.35	1.76	1986Q3**
Finland	1.00	0.05	-0.62	3.89	1988Q1***
France	0.94	0.34	-0.20	1.35	2001Q4
Germany	1.00	0.88	0.09	1.89	1988Q3*
Italy	1.00	0.01	-1.17	1.88	2003Q1***
Netherlands	1.00	0.14	-0.64	1.85	1986Q1**
Norway*	0.94	0.31	-1.27	4.36	1986Q2*
Spain	0.00	0.12	-0.59	2.14	2000Q2
Sweden*	0.99	0.03	1.22	3.30	1994Q2**
Switzerland*	0.06	0.08	0.62	2.95	2008Q4
UK*	1.00	0.69	-0.18	1.84	1986Q1

Notes: PF: Chow Predictive failure test; SS: Chow Structural stability test; Mean PE: Mean Prediction Errors; RMSPE: Root Mean Squared Prediction Error; Q-A Test: Quandt-Andrews Test (*** indicates a significance level of 1 percent, ** of 5 percent, and * of 10 percent); *: non-member countries.

Moving to the structural stability analysis, similarly to the levels equations, Table no. 3 summarises the p-values for the PF and SS tests, mean errors and RMSPE for a break in 1999Q1, as well as Q-A test breakpoints. Plots of actual and predicted values are given in an Annex.

For Germany the predicted is closest to the actual with the smallest mean prediction error in a sample, 0.09%. However, RMSPE is the smallest for France, 1.35% and its mean PE is also relatively small, -0.20%, suggesting that the country performed slightly worse than expected from the pre-euro relationship. In both cases, Germany and France, neither the PF nor SS test rejects the hypothesis of structural stability. However, Q-A test identified potential breakpoint for Germany in 1988Q3, but it is only significant at 10% level.

Similarly to France, the UK and Austria also performed a bit worse than expected, with the mean PE of -0.18% and -0.28%, respectively. The tests do not indicate either parameter instability or predictive failure. Spain, Finland, Netherlands, Italy, and Norway performed even less well with their negative mean errors and RMSPEs are on a higher end, comparing to the rest of the sample. Furthermore, in case of Spain both, Chow PF and SS tests, suggest possible structural instability. In Italy and Finland there is indication of parameter change, but not predictive failure (in case of Finland, at 10% level only). The Q-A test suggest a breakpoint in 1988Q1 in Finland and in 2003Q1 in Italy and in 1986Q1 in the Netherlands, as well as, in 1986Q2 in Norway (only significant at 10% level).

Moving to the countries which performed better than expected from the pre-euro relationship are Belgium, Switzerland, and most of all Sweden with the highest mean PE of 1.22% in the sample. While Chow PF and SS tests find no signs of structural instability in Belgium time series, the Q-A test suggest a possible breakpoint in 1986Q3. Both tests, Chow PF and SS, reject the null hypothesis of structural stability for Switzerland, but only at 10% level and the Q-A finds no structural breaks in this time series. Finally, Chow SS suggests structural instability in case of Sweden and the Q-A finds a break in 1994Q2.

Comparing the growth rate and levels relationship estimations, the Q-A test results are similar in case of Belgium and Norway, but very different for the Netherlands and Sweden. Interestingly the Q-A test does not identify a structural break around 1998-1999, third stage for the implementation of the EMU, neither when the relationship estimated in levels nor in first differences. However, for level models the Q-A finds significant breakpoints in all countries, while in growth rate models only in seven, even including breakpoints that are significant at 10% only.

Furthermore, if in the levels relationships the SS test suggested structural instability in all countries (three of them, Austria, Finland and the Netherlands, are only significant at 10% level), in the growth rate estimations it rejects the null only in four countries (in Finland and Switzerland only at 10% level). Overall, it appears the growth rate form fits the data better.

To consider how interest rate determination changed with the formation of the euro and the crisis, Taylor Rules were estimated for the pre-euro period, 1979Q4-1998Q4; the early pre-crisis euro period, 1999Q1-2008Q4; and the post crisis period 2009Q1-2016Q4. The estimated equation for each country made the short interest rate for each country, r_t a function of its lagged value, the lagged rate of inflation, π_{t-1}, and the lagged log output gap, $y_{t-1} - y^*_{t-1}$. This takes the form:

$$r_t = (1-\lambda)r_{t-1} + \lambda\left(\theta_* + \theta_\pi \pi_{t-1} + \theta_y(y_{t-1} - y^*_{t-1})\right) + u_t$$

Potential log output is approximated by a linear trend, so the estimated equation takes the form:

$$r_t = \alpha_0 + \alpha_1 r_{t-1} + \beta_\pi \pi_{t-1} + \beta_y y_{t-1} + \gamma t + u_t \tag{3}$$

In the pre-euro period both $\beta_y > 0$ in every country and significant in all but three, while $\beta_\pi > 0$ in 10, significantly so in 3 and insignificantly negative in the Netherlands and Spain. The average value of θ_π was 0.17 with a standard deviation of 0.26, ranging from -0.53 in the Netherlands to 0.47 in Belgium. The average value of θ_y was 0.15 with a standard deviation of 0.07, ranging from 0.06 in Sweden to 0.29 in the Netherlands In the second, early euro period 5 countries show negative β_π coefficients (Belgium, France, Italy, Netherlands and UK), none significantly negative, only Finland and Austria are significantly positive. β_y was significantly positive in all countries except Switzerland. The results for the third post-crisis period are subject to the fact that interest rates moved relatively little over this period. There are two negative β_π, Switzerland and the UK, neither significantly so, and 3 significantly positive

Sweden, Italy, and Austria. There are now six negative β_y, Belgium, Netherlands, Norway (significant), Spain, Sweden, Switzerland, and only the UK one is significantly positive.

For most countries the match between actual and predicted is fairly close until the crisis when the predicted interest rate falls sharply and the actual interest rate constrained by the zero lower bound cannot follow. With the exception of Sweden and Switzerland, predicted interest rates go sharply negative at the end of the sample, being below the actual. In Sweden and Switzerland, predicted is above actual for the whole period. In Germany the predicted is below the actual for the whole period. In Italy and Netherlands the predicted is below the actual from about 2002. The hypothesis of structural stability, no change in the parameters before and after the establishment of the euro is not rejected in Finland, Norway, Spain and Switzerland. Though the test is conditional on equality of variances, which is unlikely to be the case here. The standard error of regression for the pre-euro period is large relative to that for the two post euro periods.

There is little in these results that would suggest a big difference between the euro members and non-members in this group.

We proceed by estimating Taylor Rules using long-run interest rates with the estimated equation taking the following form:

$$lr_t = \alpha_0 + \alpha_1 lr_{t-1} + \beta_\pi \pi_{t-1} + \beta_y y_{t-1} + \gamma t + u_t \tag{4}$$

where lr_t is a nominal long-term interest rate per quarter.

Our sample is one country short, because the long-term interest rate data is not available for Finland.

In the pre-euro period β_y is positive in every country and significant in all but four. The average value of θ_y was 0.11 with a standard deviation of 0.13, ranging from 0.02 in the UK to 0.49 in Italy. As for inflation coefficients, β_π is positive in all but one country, significantly so in three and insignificantly negative in Norway. The average value of θ_π was 0.20 with a standard deviation of 0.30. The coefficients of θ_π ranged from -0.11 in the Norway, to almost unity (0.97) in Italy. In the second period β_π is negative for all countries except Germany, Italy, and Sweden, but none are significant. In contrast, β_y was positive for all countries, but significantly so only for Austria, Germany, the Netherlands, and Norway. As for the post-crisis period, β_π was positive for all countries except Norway, Switzerland, and the UK, but was insignificant for the whole sample. Moreover, β_y is now insignificant for all countries and is negative for Germany, the Netherlands, Sweden, Switzerland, and the UK.

As before, the results do not vary substantially between euro and non-euro members. The differences between actual and predicted are smaller for all countries when Taylor Rules are estimated using long-term interest rates instead of short-term ones.

5. SYSTEMS APPROACH

5.1 Theoretical model

Although we are not going to use a structural model, we will set out how our estimated model relates to a fully specified structural model. Consider the following rational expectations (RE) model for a small open economy in the $k \times 1$ vector $\boldsymbol{x}_t$ of endogenous variables, determined in terms of their expected future values, past values, a $k_* \times 1$ vector of corresponding foreign variables, $\boldsymbol{x}_t^*$ which are treated as exogenous, and a $k_d \times 1$ vector of deterministic elements like trend and intercept:

$$\boldsymbol{A}_0(\varphi)\boldsymbol{x}_t = \boldsymbol{A}_1(\varphi)E_t(\boldsymbol{x}_{t+1}) + \boldsymbol{A}_2(\varphi)\boldsymbol{x}_{t-1} + \boldsymbol{A}_3(\varphi)\boldsymbol{x}_t^* + \boldsymbol{A}_4(\varphi)\boldsymbol{d}_t + \boldsymbol{u}_t \tag{5}$$

For the expected future values, $E_t(\boldsymbol{x}_{t+1}) = E_t(\boldsymbol{x}_{t+1}|\mathcal{I}_t)$, the information set is $\mathcal{I}_t = (\boldsymbol{x}_t, \boldsymbol{x}_{t-1}, \dots; \boldsymbol{x}_t^*, \boldsymbol{x}_{t-1}^*, \dots)$. $\boldsymbol{A}_i(\varphi)$ are matrices of coefficients. For i=0,1,2, they are of dimension $k \times k$, for i=3 dimension $k \times k_*$, for i=4, dimension $k \times k_d$. $\boldsymbol{A}_0(\varphi)$ is non-singular, φ is a vector of deep parameters, and $\boldsymbol{u}_t$ is a $k \times 1$ vector of structural shocks. The exogenous variables are assumed to follow the VAR(1) model:

$$\boldsymbol{x}_t^* = \boldsymbol{a}(\rho) + \boldsymbol{R}(\rho)\boldsymbol{x}_{t-1}^* + \boldsymbol{\eta}_t \tag{6}$$

where $\boldsymbol{a}(\rho)$ is a $k_x \times 1$ vector of intercepts and $\boldsymbol{R}(\rho)$ is the $k_x \times k_x$ matrix of coefficients that depend on a vector of unknown coefficients, ρ. This marginal model is required because forecasts of $\boldsymbol{x}_t^*$ are required to construct the expectations $E_t(\boldsymbol{x}_{t+1})$. The errors, $\boldsymbol{u}_t$ and $\boldsymbol{\eta}_t$ are assumed to be serially and cross sectionally uncorrelated, with zero means and constant variances, $\boldsymbol{\Sigma}_u$, and $\boldsymbol{\Sigma}_\eta$, respectively.

If the quadratic matrix equation,

$$\boldsymbol{A}_1(\varphi)\boldsymbol{\Phi}^2(\varphi) - \boldsymbol{A}_0(\varphi)\boldsymbol{\Phi}(\varphi) + \boldsymbol{A}_2(\varphi) = \boldsymbol{0}$$

has a solution, $\boldsymbol{\Phi}(\varphi)$, with all its eigenvalues inside the unit circle, then, the RE model, (5) and (6), has the unique solution[1]:

$$\boldsymbol{x}_t = \boldsymbol{\Phi}(\varphi)\boldsymbol{x}_{t-1} + \boldsymbol{\Psi}(\varphi,\rho)\boldsymbol{x}_t^* + \boldsymbol{\mu}_a(\varphi,\rho)\boldsymbol{d}_t + \boldsymbol{\Gamma}(\varphi)\boldsymbol{u}_t \tag{7}$$

The variance matrix of the reduced form shocks, $\varepsilon_t = \boldsymbol{\Gamma}(\varphi)\boldsymbol{u}_t$ is

$$\boldsymbol{\Sigma}_\varepsilon(\varphi) = E({\varepsilon_t}'\varepsilon_t) = \boldsymbol{\Gamma}(\varphi)\boldsymbol{\Sigma}_u\boldsymbol{\Gamma}'(\varphi)$$

Equation (7) is labeled a VARX* in the GVAR literature. It corresponds to the reduced form of a standard simultaneous equations model, when $\boldsymbol{A}_1(\varphi) = 0$ and there are no forward looking terms. It corresponds to a vector autoregression when there are no exogenous variables, so $\boldsymbol{A}_3(\varphi) = \boldsymbol{\Psi}(\varphi,\rho) = 0$. What is relevant for the case of the euro is that the parameters of (7) may change either because the parameters of the process generating the endogenous variables, φ, change say from φ_1 to φ_2, or because the parameters of the process generating the exogenous variables, ρ, changes from ρ_1 to ρ_2. Changes in the process driving the exogenous variables could be important because they may change how people form their expectations, $E_t(\boldsymbol{x}_{t+1})$.

5.2 VARX*

The VARX* (7) was for a single country, now consider a set of countries i=0,1,2,...,N, with country 0, say the US, as the numeraire country: we use the exchange rate against the dollar. Suppressing the dependence on the deep parameters, a second-order country-specific VARX*(2,2) model with deterministic trends can be written as:

$$\boldsymbol{x}_{it} = \boldsymbol{B}_{id}\boldsymbol{d}_t + \boldsymbol{B}_{i1}\boldsymbol{x}_{i,t-1} + \boldsymbol{B}_{i2}\boldsymbol{x}_{i,t-2} + \boldsymbol{B}_{i0}^{*}\boldsymbol{x}_{it}^{*} + \boldsymbol{B}_{i1}^{*}\boldsymbol{x}_{i,t-1}^{*} + \boldsymbol{B}_{i2}^{*}\boldsymbol{x}_{i,t-2}^{*} + \boldsymbol{u}_{it} \quad (8)$$

where $\boldsymbol{x}_{it}$ is a $k_1 \times 1$ (usually six) vector of domestic variables, $\boldsymbol{x}_{it}^{*}$ is a $k_i^{*} \times 1$ vector of foreign variables specific to country i, and $\boldsymbol{d}_t$ is a $s \times 1$ vector of deterministic elements as well as observed common variables, oil prices in our case: $(1, t, p_t^o)$. The $\boldsymbol{x}_{it}^{*}$ are calculated as country specific trade weighted averages of the corresponding variables of the other countries.

$$\boldsymbol{x}_{it}^{*} = \sum_{j=0}^{N} w_{ij}\boldsymbol{x}_{jt}, \text{ with } w_{ii} = 0,$$

where w_{ij} is the share of country j in the trade (exports plus imports) of country i.

In the case of small open economies it is reasonable to assume that the $\boldsymbol{x}_{it}^{*}$ are "long run forcing" or $I(1)$ weakly exogenous, and then estimate the VARX* models separately for each country, allowing for cointegration both within $\boldsymbol{x}_{it}$ and across $\boldsymbol{x}_{it}$ and $\boldsymbol{x}_{it}^{*}$[2]. Tests for the weak exogeneity of $\boldsymbol{x}_{it}^{*}$ generally do not reject the hypothesis. The $\boldsymbol{x}_{it}^{*}$ would typically refer to the same variables as $\boldsymbol{x}_{it}$, thus there is a symmetric structure to the model.

The cointegrating VARX* can be written eq10as a VECM:

$$\Delta\boldsymbol{x}_{it} = \boldsymbol{B}_{id}\boldsymbol{d}_t + \boldsymbol{\Pi}_i\boldsymbol{z}_{i,t-1} + \boldsymbol{B}_{i0}^{*}\Delta\boldsymbol{x}_{it}^{*} + \boldsymbol{\Gamma}_i\Delta\boldsymbol{z}_{i,t-1} + \boldsymbol{u}_{it} \quad (9)$$

where $\boldsymbol{z}_{i,t} = (\boldsymbol{x}_{it}', \boldsymbol{x}_{it}^{*\prime})'$. Restricting the deterministic terms and assuming that $rank(\boldsymbol{\Pi}_i) = r_i < k_i + k_i^{*}$, we have $\boldsymbol{\Pi}_i = \alpha_i\beta_i'$, where β_i is the $(k_i + k_i^{*}) \times r_i$ matrix of the cointegrating coefficients and

$$\Delta\boldsymbol{x}_{it} = \alpha_i\beta_i'\left(\boldsymbol{z}_{i,t-1} - \boldsymbol{\Upsilon}_i\boldsymbol{d}_{t-1}\right) + \boldsymbol{B}_{i0}^{*}\Delta\boldsymbol{x}_{it}^{*} + \boldsymbol{\Gamma}_i\Delta\boldsymbol{z}_{i,t-1} + \boldsymbol{\Pi}_i\boldsymbol{\Upsilon}_i\Delta\boldsymbol{d}_t + \boldsymbol{u}_{it} \quad (10)$$

The r_i error correction terms of the model can now be written as:

$$\xi_{it} = \beta_i'\boldsymbol{z}_{i,t} - \beta_i'\boldsymbol{\Upsilon}_i\boldsymbol{d}_t = \beta_{ix}'\boldsymbol{x}_{it} + \beta_{ix*}'\boldsymbol{x}_{it}^{*} + \gamma_i'\boldsymbol{d}_t$$

The ξ_{it} are mean zero $r_i \times 1$ vectors of disequilibrium deviations from the long run relationships. Forecasts and counter-factuals are invariant to the just-identifying restrictions used to identify β_i'. To establish whether there are changes in the reduced form parameters, we do not need to identify either the structural shocks or the cointegrating vectors.

Notice that if $r_i = 0$ in (10), this gives a first difference model and if $r_i = k$ this gives an unrestricted levels VARX* (8).

5.3 Model Selection

We wish to examine whether there was a break at time T_0, the end of 1998Q4. To do this, we estimate (10) using the whole sample: 1979Q4-2016Q4, which we call period 0; then for period 1 (1979Q4-1998Q4) and period 2 (1999Q1-2016Q4) and examine the extent to which allowing for a structural break improves the fit. In estimating (10), for each country we have to (a) choose lag-length for endogenous and exogenous variables, (p_{ei}, p_{xi}), which are set to a maximum of (2,2), (b) choose the number of cointegrating vectors r_i and (c) judge the significance of any structural breaks. Although there are tests for each of these, some of which are non-standard, it seems better to make the choices for the various elements within a consistent framework. This can be done using information criteria, IC.

If model i has k_i estimated parameters and maximised log likelihood MLL_i the Akaike information criterion is $AIC_i = MLL_i - k_i$. The Schwarz Bayesian information criterion is $BIC_i = MLL_i - 0.4 \times k_i \times \ln T$, where T is the sample size[3]. Two models are estimated, one using (p_{ei}^A, p_{xi}^A) and r_i^A chosen on the basis of *AIC* and one using (p_{ei}^B, p_{xi}^B) and r_i^B chosen on the basis of *BIC*. We use period 1, pre-euro, data to make the choice.

In the case of nested models, which is what we will be concerned with, the information criteria can be interpreted as likelihood ratio tests. Suppose that the unrestricted model with MLL_U has one more parameter than the restricted model with MLL_R. Then a standard likelihood ratio test with probability of type I error, α=5%, would choose the unrestricted model if $LR = 2(MLL_U - MLL_R) > 3.84$. The AIC would choose the unrestricted model if $2(MLL_U - MLL_R) > 2$, corresponding to roughly to α=15%. The *BIC* would choose the unrestricted model if $2(MLL_U - MLL_R) > \ln T$, which for T=100 is 4.6, which corresponds to roughly α=3%. The AIC and LR keep α constant and use any extra information to increase the power. They will reject any deviation from the null, however small, for a sufficiently large sample size. The BIC reduces α with the sample size, so the probabilities of both type I and type II errors fall with sample size. The BIC is consistent in that it will choose the true model as the sample size gets large, if the true model is in the set being considered. If the true model is not in the set being considered the AIC, including more parameters may provide a better approximation to it. By using both we can judge how robust our results are to possible over-fitting or under-fitting.

We will use the difference between the sum of the IC for periods 1 and 2 and the IC for the whole period as an indication of the extent of the structural break. For the AIC, the difference is $D_A = AIC_1 + AIC_2 - AIC_0$ and $2D_A = LR - K$, where K is the number of parameters. So the AIC choice corresponds to an LR test with a critical value of K. Similarly, $D_B = BIC_1 + BIC_2 - BIC_0$, which corresponds to an LR test with a critical value of K $(\ln T_1 + \ln T_2 - \ln T)$. The dimensions are $T_1 = 77$, $T_2 = 72$, T=149, and with $(p_{ei}, p_{xi}) = (2{,}2)$ and $r_i = 6$ there are 25 parameters in each equation.

5.4 Conditioning

In order to identify where the structural changes originate, we will condition on some elements of $\boldsymbol{x}_{it}$, say $\boldsymbol{x}_{2,it}$, in particular interest rates and exchange rates, and treat them as exogenous, in explaining $\boldsymbol{x}_{1,it}$. The interpretation of this process follows Pesaran and Smith (1998). For clarity of exposition we abstract from the country identifier, i, the deterministic terms $\mathbf{d}_t$ and the other exogenous foreign variables, $\boldsymbol{x}_{it}^*$. Then (9) can be written as:

$$\Delta\boldsymbol{x}_t = \boldsymbol{\Pi}\boldsymbol{x}_{t-1} + \boldsymbol{\Gamma}\Delta\boldsymbol{x}_{t-1} + \boldsymbol{u}_t \tag{11}$$

We now partition $\boldsymbol{x}_t = (\boldsymbol{x}'_{1,t}, \boldsymbol{x}'_{2,t})'$ to give:

$$\Delta\boldsymbol{x}_{1t} = \boldsymbol{\Pi}_{11}\boldsymbol{x}_{1,t-1} + \boldsymbol{\Pi}_{12}\boldsymbol{x}_{2,t-1} + \boldsymbol{\Gamma}_{11}\Delta\boldsymbol{x}_{1,t-1} + \boldsymbol{\Gamma}_{12}\Delta\boldsymbol{x}_{2,t-1} + \boldsymbol{u}_{1t}$$

$$\Delta \boldsymbol{x}_{2t} = \boldsymbol{\Pi}_{21}\boldsymbol{x}_{1,t-1} + \boldsymbol{\Pi}_{22}\boldsymbol{x}_{2,t-1} + \boldsymbol{\Gamma}_{21}\Delta \boldsymbol{x}_{1,t-1} + \boldsymbol{\Gamma}_{22}\Delta \boldsymbol{x}_{2,t-1} + \boldsymbol{u}_{2t}$$

where the covariance matrix of the reduced form disturbances is given by:

$$\boldsymbol{\Sigma} = \begin{pmatrix} \boldsymbol{\Sigma}_{11} & \boldsymbol{\Sigma}_{12} \\ \boldsymbol{\Sigma}_{21} & \boldsymbol{\Sigma}_{22} \end{pmatrix}$$

This partition does not impose any restrictions in itself, but provides a framework for examining how exogenous variables relate to the structure of the *VAR*.

To condition $\boldsymbol{x}_{1t}$ on current values of $\boldsymbol{x}_{2t}$, define $E(\boldsymbol{u}_{1t}|\boldsymbol{u}_{2t}) = \boldsymbol{\Sigma}_{12}\boldsymbol{\Sigma}_{22}^{-1}\boldsymbol{u}_{2t} = \boldsymbol{\Theta}\boldsymbol{u}_{2t}$ with $\boldsymbol{u}_{1t} = \boldsymbol{\Theta}\boldsymbol{u}_{2t} + \eta_t$. The system for $\boldsymbol{x}_{1t}$, can then be written:

$$\begin{aligned} \Delta \boldsymbol{x}_{1t} &= (\boldsymbol{\Pi}_{11} - \boldsymbol{\Theta}\boldsymbol{\Pi}_{21})\boldsymbol{x}_{1,t-1} + (\boldsymbol{\Pi}_{12} - \boldsymbol{\Theta}\boldsymbol{\Pi}_{22})\boldsymbol{x}_{2,t-1} + \boldsymbol{\Theta}\Delta \boldsymbol{x}_{2,t} \\ &\quad + (\boldsymbol{\Gamma}_{11} - \boldsymbol{\Theta}\boldsymbol{\Gamma}_{21})\Delta \boldsymbol{x}_{1,t-1} + (\boldsymbol{\Gamma}_{12} - \boldsymbol{\Theta}\boldsymbol{\Gamma}_{22})\Delta \boldsymbol{x}_{2,t-1} + \eta_t \\ \Delta \boldsymbol{x}_{1t} &= \boldsymbol{B}_1\boldsymbol{x}_{1,t-1} + \boldsymbol{B}_2\boldsymbol{x}_{2,t-1} + \boldsymbol{C}_{20}\Delta \boldsymbol{x}_{2,t} + \boldsymbol{C}_{21}\Delta \boldsymbol{x}_{1,t-1} + \boldsymbol{C}_{22}\Delta \boldsymbol{x}_{2,t-1} + \eta_t \end{aligned} \tag{12}$$

By construction $E(\eta_t|\Delta \boldsymbol{x}_t) = \boldsymbol{0}$, and the parameters of (12) can be estimated efficiently by OLS. Also denoting the (conditional) variance of η_t by $\boldsymbol{\Sigma}_{\eta\eta}$ it is easily seen that

$$\boldsymbol{\Sigma}_{\eta\eta} - \boldsymbol{\Sigma}_{11} = -\boldsymbol{\Sigma}_{12}\boldsymbol{\Sigma}_{22}^{-1}\boldsymbol{\Sigma}_{21} \leq \boldsymbol{0}$$

The variance of η_t will generally be smaller than that of $\boldsymbol{u}_{1t}$, so the parameters in the conditional model, (12), are likely to be estimated more precisely than the parameters of the unconditional model (11). Whether this is an advantage depends on what the economic parameters of interest are. If the parameters of interest are $\boldsymbol{\Pi} = (\boldsymbol{\Pi}_{11}, \boldsymbol{\Pi}_{12})$, it is clear from equation (12) that $\Delta \boldsymbol{x}_{2,t}$ will be weakly exogenous for $\boldsymbol{\Pi}_{11}$ only if either $\boldsymbol{\Sigma}_{12} = \boldsymbol{0}$, so that $\boldsymbol{\Theta} = \boldsymbol{0}$, or if $\boldsymbol{\Pi}_2 = (\boldsymbol{\Pi}_{21}, \boldsymbol{\Pi}_{22}) = \boldsymbol{0}$[4]. In either of these cases the coefficient matrix on $(\boldsymbol{x}_{1,t-1}, \boldsymbol{x}_{2,t-1})$ in the conditional model (12) will provide an estimate of $\boldsymbol{\Pi}$, otherwise it will not. In other cases, the economic parameters of interest may be the long-run effects of $\boldsymbol{x}_{2,t}$ on $\boldsymbol{x}_{1,t}$ so one might be interested in $(\boldsymbol{\Pi}_{12} - \boldsymbol{\Theta}\boldsymbol{\Pi}_{22})$ directly, in which case the model conditional on $\boldsymbol{x}_t$ is appropriate whether or not $\boldsymbol{\Pi}_2 = \boldsymbol{0}$.

For some purposes we are interested in the complete system but for other purposes we are interested in the responses to particular policy variables and how these responses changed with the introduction of the euro. In this case, the parameters of interest are the parameters of the conditional model (12). Of particular interest is the case where the parameters of the marginal model, the process generating the policy variables, interest rates and exchange rates, changed, shifting $\boldsymbol{\Pi}_{21}$, $\boldsymbol{\Gamma}_{22}$ and $\boldsymbol{\Theta}$,but the parameters of the conditional model, $\mathbf{B}_i$, $\mathbf{C}_{ij}$, did not change[5].

6. SYSTEMS RESULTS

The lag orders for the endogenous and exogenous variables (p_e, p_x) and the number of cointegrating vectors *r* could be chosen from the pre-euro sample; the post-euro sample or the whole sample. We determined them on the basis of the pre-euro sample, since any subsequent change will appear as a structural break. Thus for each country we estimate a cointegrating VARX* for period 1: 1979Q4-1988Q4, and use the AIC to determine

(p_{ei}^A, p_{xi}^A) and r_i^A and the BIC to determine (p_{ei}^B, p_{xi}^B) and r_i^B. Using these values we estimate the AIC model and the BIC model for period 2: 1999Q1-2016Q4 and for period 0, with no break: 1979Q4-2016Q4. This gives the results summarised in the Table no. 4.

As one would expect the AIC model tends to have larger values of (p_{ei}, p_{xi}) and r_i. In some cases such as Sweden, $r_i^A = 6$, indicating that all the variables are $I(0)$, while $r_i^B = 0$, indicating that all the variables are $I(1)$ and not cointegrated. The differences of AIC and BIC are all positive, indicating that the model with a break at the time of the euro formation is preferred. The differences are always smaller for the BIC than for the AIC because the BIC imposes a heavier penalty for the extra parameters in the break model. If we rank AIC by differences, the break seems smaller for the non-euro countries: Sweden had the 8th smallest difference, Norway 10th, UK 11th and Switzerland 12th. Among the euro countries, Finland ranked 9th had the smallest difference. The ranking by BIC is similar, the difference in ranks is small except for Sweden which goes from 8th by AIC to 4th by BIC. The table indicates that there is evidence for a break and it seems larger in the euro countries than the non-euro countries.

Table no. 4 – Akaike and Bayesian Information Criteria for a euro structural break, m=6

	Model	*r*	Order *e,x*	IC by Period 0	1	2	Difference	Rank
Austria	A	5	1,1	3362.7	1799	1842.6	278.9	1
	B	1	1,1	3238.5	1683	1727.3	171.8	1
Belgium	A	6	2,2	3377.8	1728.6	1866.4	217.2	6
	B	1	1,1	3254.3	1625.6	1746.4	117.7	7
Finland	A	4	2,1	2240.2	1131.4	1256	147.2	9
	B	0	1,1	2153.4	1054.6	1145.8	47	11
France	A	4	2,1	3524.1	1746.5	2024.3	246.7	3
	B	0	1,1	3397.9	1655.3	1899.1	156.5	2
Germany	A	6	2,2	3508.3	1775.7	1956.5	223.9	4
	B	3	1,1	3365.5	1658	1831.9	124.4	5
Italy	A	6	2,2	3182.2	1580.4	1800.7	198.9	7
	B	1	1,1	3034.6	1460.4	1678.5	104.4	8
Netherlands	A	6	2,1	3547.7	1792.9	1975.4	220.6	5
	B	1	1,1	3435.2	1698.8	1856.6	120.2	6
Norway	A	5	2,1	3041.2	1569.7	1615.6	144.1	10
	B	2	1,1	2871.6	1451.8	1507.2	87.4	9
Spain	A	6	2,1	3141.1	1620.1	1770.2	249.2	2
	B	1	1,1	2973.7	1474.7	1647.7	148.7	3
Sweden	A	6	2,1	3093.8	1507.1	1768.9	182.2	8
	B	0	1,1	2948.8	1404.6	1670.3	126.1	4
Switzerland	A	4	2,1	3423.9	1765.5	1740.2	81.8	12
	B	0	1,1	3294.3	1649.3	1660.2	15.2	12
UK	A	5	1,1	3409.9	1711.7	1839.1	140.9	11
	B	2	1,1	3283.1	1596.6	1749.1	62.6	10

Notes: Model A is chosen by AIC, Model B by BIC. *r* is the number of cointegrating vectors. *e,x* gives the lag orders on endogenous and exogenous variables. The information criteria are given for periods: 0: 1979Q4-2016Q4; 1: 1979Q4-1998Q4; 2: 1999Q1-2016Q4. Difference gives the value for IC0-(IC1+IC2), rank gives rank of the difference. Endogenous variables are *y*, *dp*, *eq*, *ep*, *r*, *lr*; except Finland where *lr* is not available. Exogenous variables are *ys*, *dps*, *eqs*, *rs*, *lrs*, *poil*.

We next repeat the exercise treating r_{it}, the short interest rate as exogenous, since it is controlled externally by the ECB for the euro countries in the second period (Table no. 5).

Table no. 5 – Akaike and Bayesian Information Criteria for a euro structural break, m=6

	Model Criterion	*k=6*	*k=5*	*k=4*
Austria	A	278.90	95.90	75.10
	B	171.80	-2.80	-10.10
Belgium	A	217.20	106.70	82.20
	B	117.70	-1.70	-14.30
Finland	A	147.20	56.17	39.50
	B	47.00	-27.67	-13.50
France	A	246.70	132.10	110.70
	B	156.50	51.60	51.10
Germany	A	223.90	152.90	120.80
	B	124.40	59.90	33.30
Italy	A	198.90	152.70	123.60
	B	104.40	69.90	72.10
Netherlands	A	220.60	113.00	66.20
	B	120.20	30.20	21.30
Norway	A	144.10	92.50	73.10
	B	87.40	14.60	13.00
Spain	A	249.20	98.90	71.50
	B	148.70	-16.80	-7.60
Sweden	A	182.20	97.60	78.60
	B	126.10	45.00	15.40
Switzerland	A	81.80	63.70	36.60
	B	15.20	-13.50	-16.40
UK	A	140.90	72.50	50.40
	B	62.60	5.60	-6.10

Notes: Difference in AIC, A, and BIC, B, between whole period and two sub-periods. For 3 models. *k=5*: interest rates exogenous; *k=4*: interest rates and exchange rates exogenous. Negative value suggests no structural break.

We thus see how large the structural break is controlling for interest rates. Since the size of the system changes from k=6 to k=5 (except for Finland where it changes from k=5 to k=4), we again need to choose (p_{ei}^A, p_{xi}^A) and r_i^A and (p_{ei}^B, p_{xi}^B) and r_i^B on the first period data. Assuming the real exchange rate, ep_{it}, is controlled by the ECB, to an extent after the formation of the euro, we continue the analysis treating the exchange rate as exogenous. Hence, further reducing the number of endogenous variables in the system, from k=5 to k=4, except for Finland (in which case k=4 is reduced to k=3)**.** The table gives the differences in the AIC and BIC between the whole period and the sum of the two sub-periods for the three cases. For k=6, the first column, the differences are the same as in the previous table. When one controls for the short interest rate and exchange rate, there is clearly less evidence for a structural break, suggesting that the main breaks came in interest rate and exchange rate equations. For half of the sample, namely for Austria, Belgium, Finland, Spain, Switzerland, and the UK, the difference of BIC is negative, meaning the model estimated over the whole period is preferred. The big reduction appears to come from the interest rate equation.

Table no. 6 – Mean Error of Forecast 1999Q1-2016Q4, GDP

	Level		Growth	
	Single	**System**	**Single**	**System**
Austria	0.84	-0.05	-0.07	-0.05
Belgium	5.20	1.32	0.09	0.04
Finland	10.33	-1.96	-0.16	-0.15
France	3.88	0.22	-0.05	-0.06
Germany	-0.07	-2.61	0.02	0.05
Italy	-6.79	-8.94	0.35	-0.31
Netherlands	-5.95	-4.79	-0.16	-0.14
Norway	-11.02	-9.32	-0.32	-0.28
Spain	-8.80	-1.67	-0.15	-0.17
Sweden	14.65	10.94	0.31	0.37
Switzerland	13.10	3.05	0.15	0.08
UK	-58.78	0.74	-0.04	-0.13

Notes: Level in percent; Growth in percentage points at annual rates.

We compared the system and single equation results (Table no. 6). The level equations for log GDP seem unreliable, the series seem to be difference stationary rather than trend stationary. The mean errors of forecast for growth rates from systems and single equation estimates always the same sign and very similar in magnitude, difference less than 0.05 except for Sweden, Switzerland and UK. Of the non-euro countries 2 out of 4 did worse than expected. The outliers among the 12 are non-euro: Sweden had the largest positive difference of actual over expected; Norway the largest negative difference. Of the euro countries 6 out of 8 did worse than expected. Germany and Belgium did slightly better, Italy much worse, Spain and Netherlands quite a lot worse.

7. CONCLUSION

From this analysis we can only really conclude that it is difficult to draw strong conclusions. The results are sensitive to a range of specification choices and the confidence intervals around counterfactuals are large. The main structural break seems to be in the interest and exchange rate equations, where there was a clear institutional change in their determination with the introduction of the euro. The date of the formation of euro is not identified as the most likely date for a structural break in the GDP equations and the GDP growth rate equation shows no structural break for many countries. There do not appear to be obvious differences in the patterns of structural breaks between euro and non-euro country equations. The effects of the euro are apparent everywhere except in the main macroeconomic equations. This may be significant. The formation of the euro was a major break which required a change in the patterns of economic relationships to provide alternative methods of economic adjustment to changes in interest and exchange rates. The fact that the economic relationships did not seem to have changed may have been a source of tensions for the euro.

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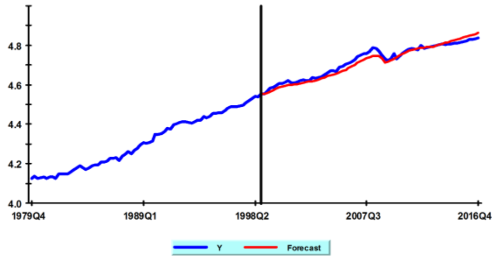
ANNEX

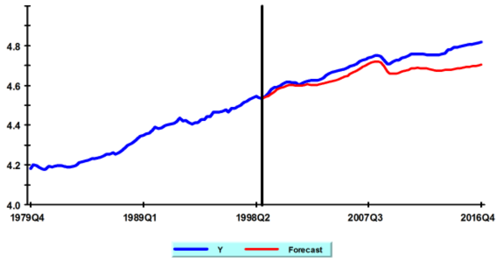
Forecasts (the level of log GDP)

Plots of Actual and Single Equation Dynamic Forecast(s) of Y

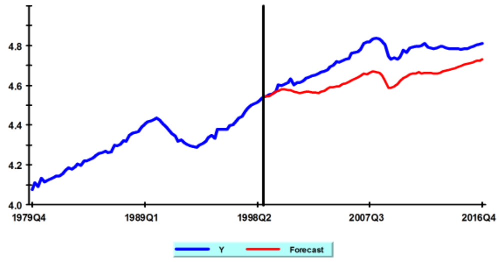
Austria



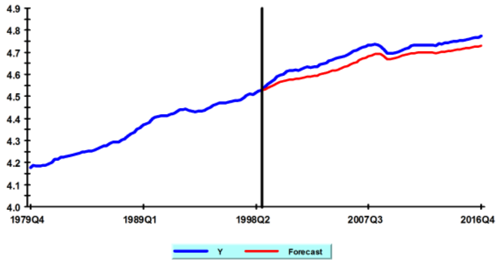
Belgium



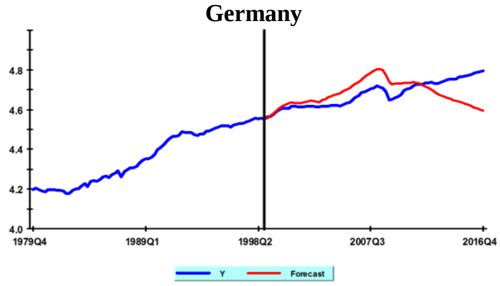
Finland



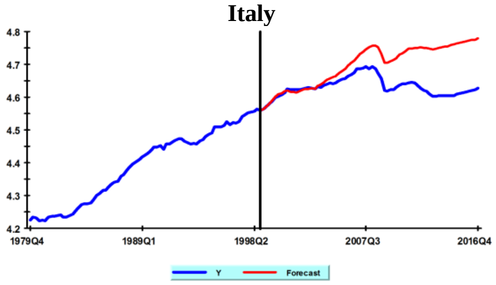
France



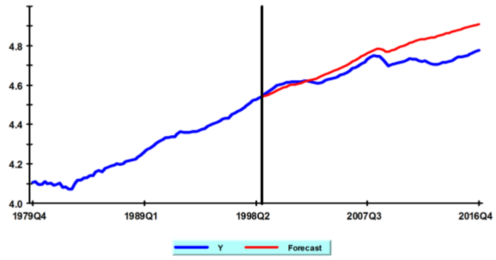
Germany



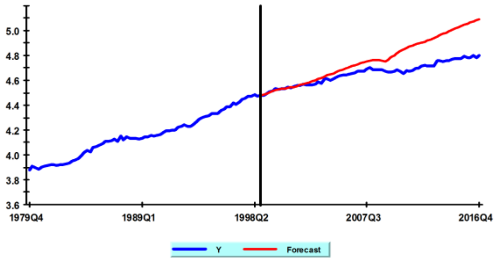
Italy

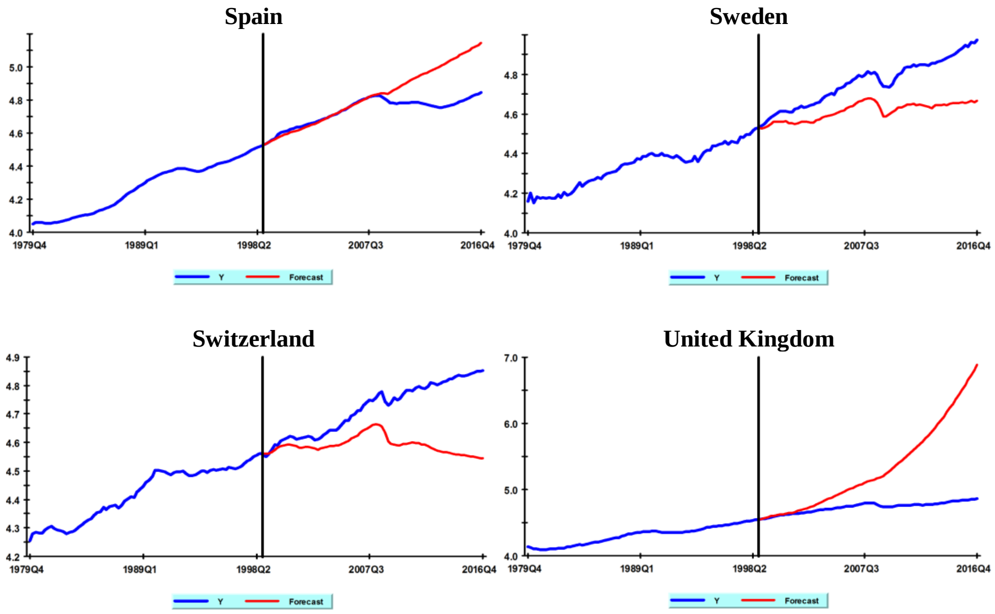


Netherlands



Norway





Notes

[1] See, for instance, Chapter 20 of Pesaran (2015).

[2] This is unlikely to apply to a large economy like the US which may influence world interest rates. But it seems reasonable for the European countries we consider.

[3] Some programs report -2 times these numbers.

[4] When the restrictions $\mathbf{\Pi}_2 = \mathbf{0}$ hold, $\boldsymbol{x}_{2t}$ is referred to as "long-run forcing" for $\boldsymbol{x}_{1t}$. This is different from Granger non-causality, GNC. $\boldsymbol{x}_{2t}$ is said to be GNC for $\boldsymbol{x}_{1t}$ if $\mathbf{\Pi}_{12} = \mathbf{0}$ and $\mathbf{\Gamma}_{12} = \mathbf{0}$; $\boldsymbol{x}_{2t}$ does not predict $\boldsymbol{x}_{1t}$. If $\mathbf{\Pi}_2 = \mathbf{0}$, $\boldsymbol{x}_{2t}$ cannot themselves be cointegrated..

[5] As is clear from (7) the Lucas critique says that any change in the marginal model determining policy will change the conditional model.