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Persistence in real GDP: Evidence from Europe and the US

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Abstract

This note provides extensive evidence on the persistence properties of real GDP in 17 European countries and in the US over the period 1960-2023 using a fractional integration framework. The analysis suggests that in all cases shocks have permanent effects on the level of real GDP. This is consistent with the idea that it is the growth rate of output which is stationary and fluctuates around a long-run equilibrium level. Further, the degree of persistence varies across countries, with the US, Greece and Spain exhibiting the highest one and Sweden and Ireland the lowest. Policy makers should take such properties into account when formulating appropriate stabilisation policies.

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1. Introduction

Real GDP is a key measure of the wealth of a nation and of its economic performance. Therefore, understanding its stochastic behaviour is of crucial importance. In a well-known paper Nelson and Plosser (1982) modelled it (together with other real macroeconomic variables) as a non-stationary process including both a secular (non-stationary) component corresponding to the long run and a (stationary) cyclical one. Note that in Real Business Cycle (RBC) models (King et al., 1987; Shapiro and Watson, 1988) real GDP is not stationary in levels but exhibits a trend over time; however, it becomes stationary when detrended (normally by taking the log difference) and then fluctuates around a long-run growth equilibrium level.

Clearly, the effects of shocks are different in a trend stationary vis-à-vis a random walk or nonstationary framework. The early literature was based on a dichotomy between integrated of order 0, or I(0), stationary variables, and I(1), non-stationary ones, and unit root tests were carried out to distinguish between the two (Dickey and Fuller, ADF, 1979; Phillips and Perron, PP, 1988; Kwiatkowski et al., KPSS, 1992; Elliot et al., ERS, 1996; etc.). Studies following this approach to model real GDP include Perron and Phillips (1987), Schwert (1987), Campbell and Mankiw (1987), Perron (1988), Rudebusch (1993), and Diebold and Senhadji (1996) in the case of the US, and Zelhorst and De Haan (1995), Ben-David and Papell (1998), Ben-David et al. (2003), and Narayan (2007) for a wider set of countries.

However, a more general fractional integration framework was subsequently introduced (Granger, 1980; Granger and Joyeux, 1980) to allow for fractional degrees of integration. Gil-Alana and Robinson (1997) then showed that the fourteen US macroeconomic variables examined by Nelson and Plosser (1982) were in fact fractional processes with orders of integration in the interval (0, 1). Following the same approach various other studies provided additional evidence that real GDP is indeed characterised by long memory; these include Hosking (1981, 1984), Granger and Joyeux (1980), Beran (1992, 1994), Baillie (1996), Robinson (1995a, 1995b), Caporale and Gil-Alana (2009, 2013, 2022), and Caporale and Skare (2018).

The present note revisits this issue by using fractional integration techniques to examine the stochastic behaviour of real GDP in a wide set of 17 European countries as well as in the US over the period from 1960 to 2023. This approach sheds light on whether or not the series are mean reverting, whether the effects of shocks are transitory or permanent, and the speed of the dynamic adjustment process. Thus, the main innovation with respect to previous works is the use of an updated version of the series in a large group of European countries along with the US. We also examine the stability of the degree of persistence by estimating first the order of integration of the series with subsamples ending at 2000, and then adding one observation (year) at a time. Our results indicate that in Europe Greece and Spain exhibit the highest degree of persistence and Ireland and Sweden the lowest, the US series displaying a similar degree of persistence to the two former countries. This is essential information for policy makers to be able to decide on the appropriate course of action in response to shocks affecting real GDP. Therefore the economic relevance of fractional integration reflects the fact that it yields accurate measures of persistence with clear policy implications, since the policies required in response to shocks depend on their degree of persistence and thus whether they have permanent or transitory effects.

The note is structured as follows: Section 2 describes the data and the methodology and presents the empirical findings, Section 3 offers some concluding remarks.

2. Data, Methodology and Empirical Results

We analyse real GDP annual data for 17 European countries, namely Austria, Belgium, Denmark, Finland, France, Germany, Greece, Iceland, Ireland, Italy, Luxembourg, Netherlands, Portugal, Spain, Sweden, Switzerland and the UK, as well as for the US. The sample period goes from 1960 to 2023 in all cases. The data source is macrotrends.net.

The estimated regression model is the following:

$$y(t) = \alpha + \beta t + x(t), \quad (1-L)^d x(t) = u(t), \quad t = 1, 2, \ldots \tag{1}$$

where y(t) is the time series of interest (the logged value of real GDP in the present case), L is a lag operator, d the fractional differencing parameter, α the intercept, β the coefficient on a linear time trend *t*, and u(t) is assumed to be a white noise process.

Table I: Estimated coefficients based on the model given by Equ. (1)

Country	d (95% conf. band)	Intercept (t-value)	Time trend (t-value)
Austria	1.27 (1.10, 1.54)	6.7737 (75.99)	0.0671 (2.25)
Belgium	1.32 (1.12, 1.63)	7.0981 (80.05)	0.0619 (1.68)
Denmark	1.23 (1.07, 1.50)	7.1586 (82.55)	0.0642 (2.40)
Finland	1.26 (1.06, 1.63)	7.0051 (71.88)	0.0634 (1.94)
France	1.20 (1.05, 1.46)	7.1285 (70.30)	0.0575 (2.36)
Germany	1.23 (1.07, 1.50)	6.9754 (74.58)	0.0636 (2.27)
Greece	1.38 (1.21, 1.65)	6.1634 (74.69)	0.0739 (1.75)
Iceland	1.23 (0.95, 1.73)	7.2056 (54.74)	0.0653 (1.65)
Ireland	1.16 (0.96, 1.47)	6.4813 (76.93)	0.0782 (4.04)
Italy	1.21 (1.06, 1.46)	6.6499 (72.14)	0.0635 (2.48)
Luxembourg	1.26 (1.05, 1.58)	7.6765 (80.75)	0.0597 (1.87)
Netherlands	1.27 (1.10, 1.54)	6.9761 (79.48)	0.0674 (2.21)
Portugal	1.29 (1.10, 1.61)	5.8618 (79.33)	0.0671 (2.28)
Spain	1.34 (1.15, 1.65)	5.9088 (60.46)	0.0818 (1.87)
Sweden	1.16 (0.98, 1.45)	7.6023 (77.25)	0.0524 (2.32)

Switzerland	1.22 (1.06, 1.48)	7.5034 (83.99)	0.0644 (2.50)
United Kingdom	1.26 (1.07, 1.62)	7.1868 (8.16)	0.0556 (1.94)
United States	1.38 (1.27, 1.57)	7.9620 (360.2)	0.0517 (4.59)

Note: the values in brackets in column 2 are the 95% confidence intervals of the non-rejection values of d, while those in columns 3 and 4 are the t-values for the corresponding coefficients.

Table I reports the estimates of d (along with the 95% confidence bands), obtained using Robinson's (1994) LM approach, together with those of α and β. This approach is based on the likelihood function in the frequency domain and is preferred to others like Sowell's (1992) or some semiparametric methods because it has some features that makes it attractive. Among them we mention that it is valid for any real value d, and thus, it does not require preliminary differentiation if the series are nonstationary ; moreover, the limit distribution is standard normal and this behaviour holds independently of the inclusion or not of deterministic terms like an intercept and a linear time trend. This is also unusual compared with the classical unit root tests where numerical values have to be computed on a simulation basis for the critical values; finally, this method (Robinson, 1994) is the most efficient one in the Pitman sense agains local departures, a feaure that is also relevant in our fractional context. The specific functional form of the version used in this work can be found in Gil-Alana and Robinson (1997) and Gil-Alana (1998).[1]

It can be seen in Table 1 that all of them are statistically significant. In particular, in the case of d they are significantly higher than 1 except for the cases of Iceland, Ireland and Sweden, where the unit root null hypothesis cannot be rejected. The highest estimates of d are obtained in the case of Greece and the US (1.38), followed by Spain (1.34), whilst the lowest values are found for Sweden and Ireland (1.16). Note

[1] The use of alternative methods such as Sowell's (1992) maximum likelihood approach produced very similar results to those reported in this work; however, using semiparametric methods like those based on the log-periodogram regression (Geweke and Porter-Hudak, GPH, 1983; Robinson, 1995; Kim and Phllips, 2006) produced results that were very sensitive to the bandwidth parameters, probably due to the short sample sizes used in this application.

that the model has been estimated using the logged series in levels, therefore their first differences are the annual growth rates, and the corresponding values of d can be obtained in each case by subtracting 1 from the estimates for the logged levels – thus they would be 0.38 for Greece and the US and 0.34 for Spain (the highest values), and 0.16 (the lowest value) for Sweden and Ireland, with he effects of shocks disappearing at a much faster rate in the latter countries.

Further, the time trend coefficient displays the highest values in the cases of Spain (0.818), followed by Ireland (0.782) and Greece (0.739). This is not surprising, given the relatively low GDP level of those three countries at the beginning of the sample and the subsequent catch-up during the process of convergence.

Finally, we have included two figures in the Appendix. Figure A1 displays the impulse response functions for each country. They were obtained by using the infinite MA representation of the I(d-1) processes for the first differenced series and considering a 1-standard deviation shock. A decaying pattern is observed in the growth rate in all countries. It can be seen that for most of the countries the decay is fast, staying below 20% of its effect after the first two periods. The exceptions are Greece and the US where the adjustment is slower than in the rest of the countries. This might suggest that supply shocks resulting in technological changes over a long time span are more prominent in these countries. Figure A2 deals with the potential presence of breaks in the data; however, instead of including breaks that would produce subsamples with a very small number of observations, we adopted by a different strategy, examining if the differencing parameter has changed across time for each series. For this purpose we estimated d first for a subsample ending at 2000; then we re-estimate the order of integration adding one observation each time until the end of the sample. Results for each series are presented in Figure A2. In most cases a relatively stable pattern is found, with a slight reduction in the values of d around the 9th subsample that corresponds to data ending at 2008, the time of the financial crisis, and a subsequent increase in the following value (with data ending at 2009). This behaviour is particularly noticeable in the cases of Greece, Iceland, Ireland, UK and the US. For the latter country (the US) we also observe a sizable fall in the estimates of d with data ending at 2021, suggesting a potential additional break at this point in time for this country.

3. Conclusions

This note provides extensive evidence on the persistence properties of real GDP in 17 European countries and in the US over the period 1960-2023 using a fractional integration framework. The chosen approach is more general and flexible than standard models based on the I(0) versus I(1) dichotomy since it allows the differencing parameter to take any real values, including fractional ones, and thus allows for a wider range of stochastic processes and provides more thorough information about the degree of persistence of the series.

The analysis suggests that in all cases shocks have permanent effects on the level of real GDP. This is consistent with the idea that it is the growth rate of output which is stationary and fluctuates around a long-run equilibrium level. Interestingly, the degree of persistence varies across countries, with the US, Greece and Spain exhibiting the highest one and Sweden and Ireland the lowest. Policy makers should take such properties into account when formulating economic policies. In particular, they should pay attention to whether the degree of persistence is high, which would suggest that Real Business Cycle models driven by technology shocks are more relevant, or instead low persistence is observed, which would imply that stabilisation policies counteracting transitory demand shocks are more appropriate.

Future research should investigate the reasons for the differences between countries in Europe and the US in terms of the dynamic responses of real GDP to shocks. In addition, alternative models still based on fractional integration but allowing for breaks and/or incorporating non-linear structures could also be estimated. More specifically, the possible presence of structural breaks could be tested by using the Bai and Perron (2003) approach or performing the tests proposed by Gil-Alana (2008) and Hassler and Meller (2014), both specifically designed for the case of fractional integration. As for non-linearities, these could be modelled using methods based on Chebyshev's polynomials (Cuestas and Gil-Alana, 2016), Fourier transform functions (Gil-Alana and Yaya, 2021; Caporale et al., 2022) or neural networks (Yaya et al., 2021). Nonlinearities might result from financial crises, or other changes such as EU access and institutional changes.

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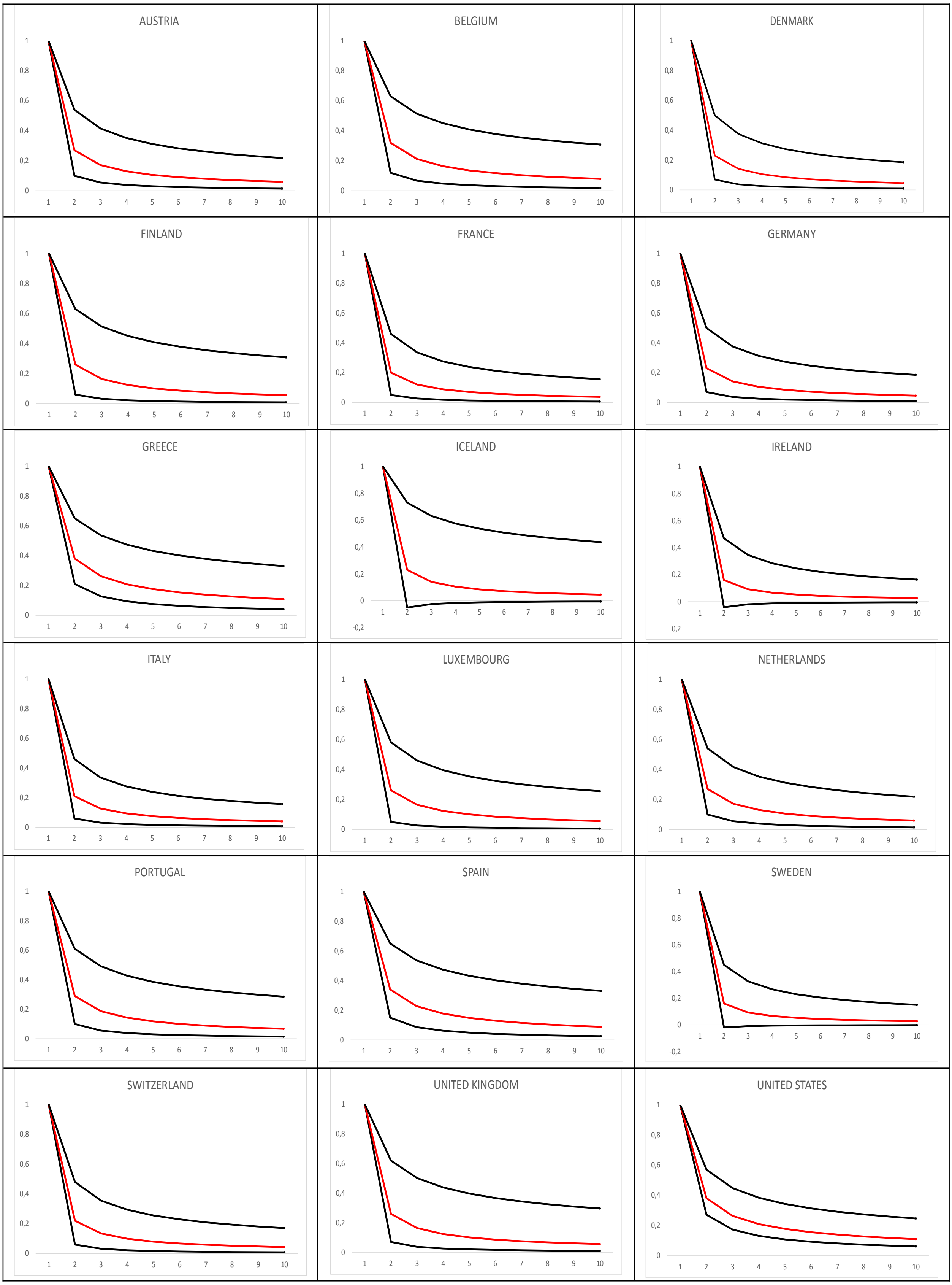
Figure A1: Impulse response functions. Time plots



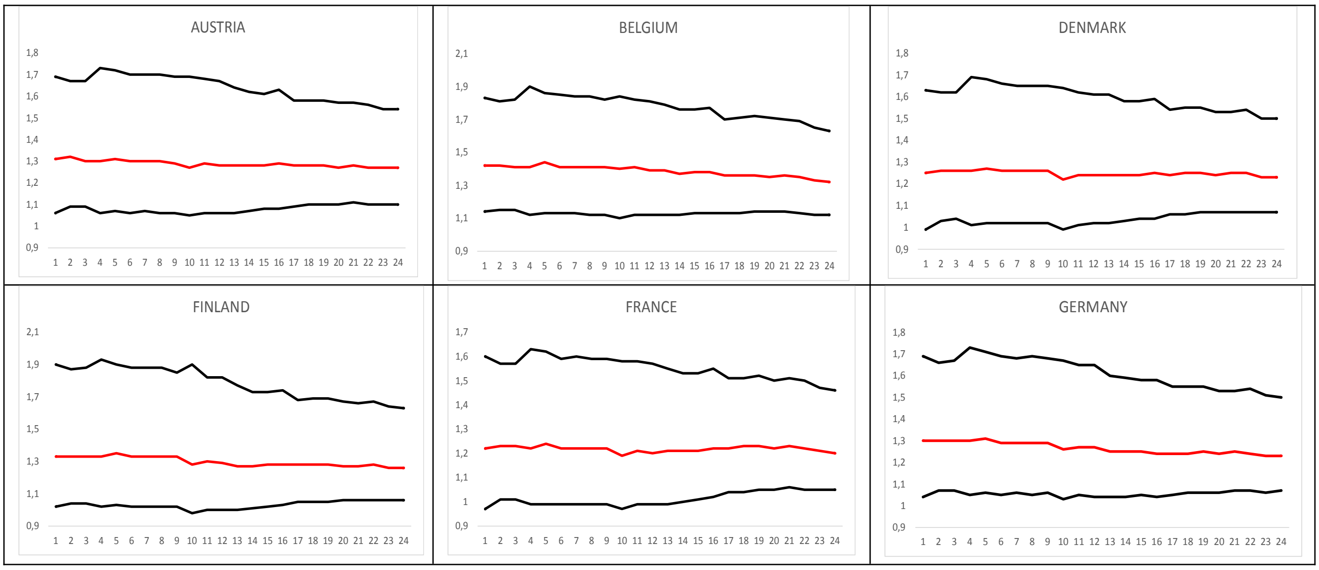
Note: The red lines are the impulse responses to a 1-standard deviation shock while the black ones are the 95% confidence bands.

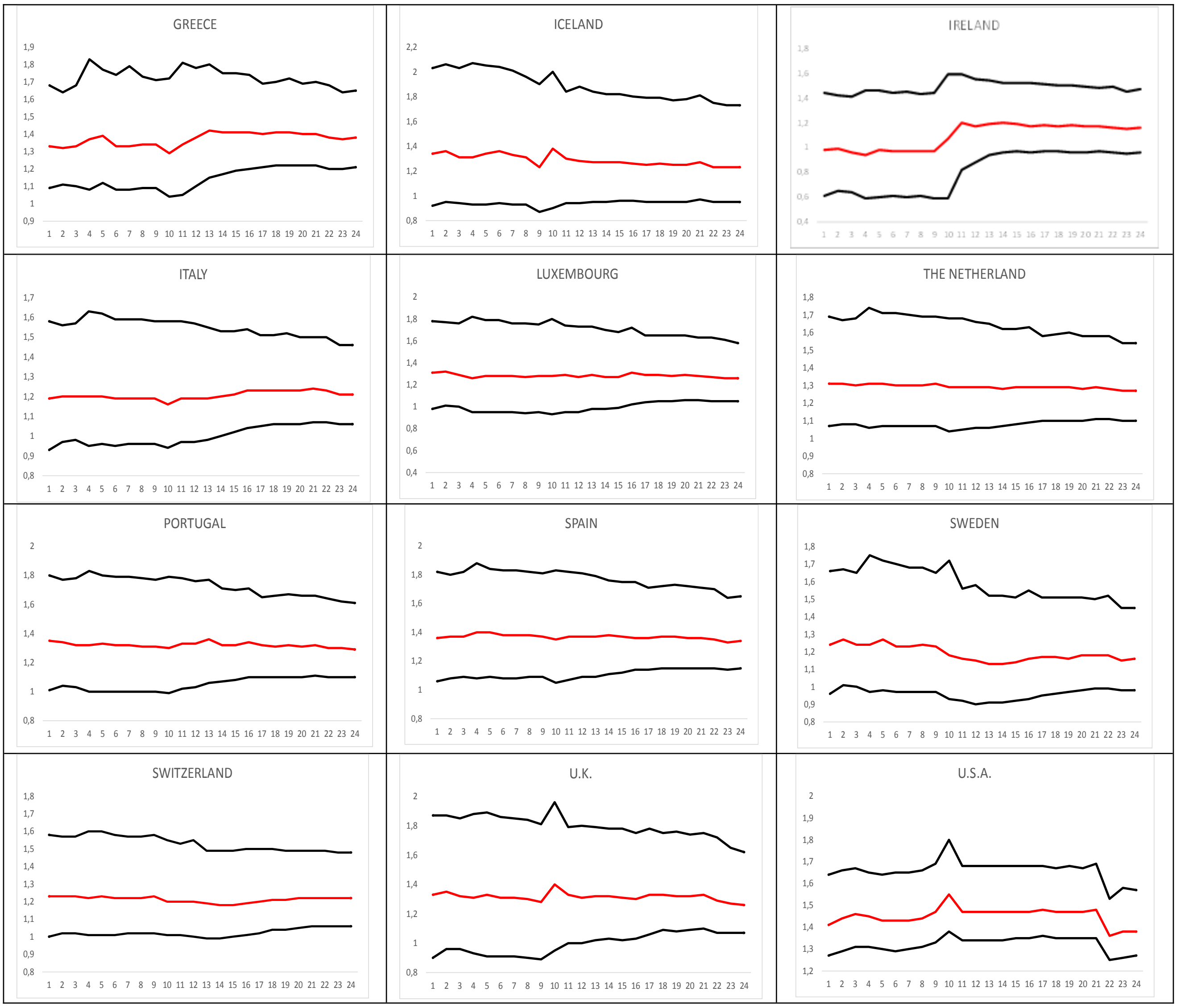
Table A1: Impulse response function. Numerical values: US vs. Ireland. Greece vs. Sweden

	d = 1.38		d = 1.16	
Value	United States	Greece	Ireland	Sweden
1	1.000	1.000	1.000	1.000
2	0.3800 (0.2700, 0.5700)	0.3800 (0.2100, 0.6500)	0.1600 (-0.0400, 0.4700)	0.1600 (-0.0200, 0.4500)
3	0.2622 (0.1714, 0.4474)	0.2622 (0.1270, 0.5362)	0.0928 (0.0192, 0.3453)	0.0928 (-0.0098, 0.3262)
4	0.2080 (0.1297, 0.3833)	0.2080 (0.0936, 0.4737)	0.0668 (0.0125, 0.2844)	0.0668 (0.0065, 0.2664
5	0.1757 (0.1060, 0.3421)	0.1757 (0.0751, 0.4322)	0.0528 (0.0093, 0.2467)	0.0528 (0.0048, 0.2298)
6	0.1540 (0.0906, 0.3127)	0.1540 (0.0632, 0.4020)	0.0439 (0.0073, 0.2206)	0.0439 (0.0038, 0.2045)
7	0.1380 (0.0795, 0.2903)	0.1380 (0.0549, 0.3785)	0.0377 (0.0061, 0.2011)	0.0377 (0.0032 0.1858)
8	0.1258 (0.0713, 0.2724)	0.1258 (0.0488, 0.3596)	0.0332 (0.0052, 0.1859)	0.0332 (0.0027, 0.1712)
9	0.1160 (0.0647, 0.2578)	0.1160 (0.0439, 0.3439)	0.0297 (0.0045, 0.1736)	0.0297 (0.0024, 0.594)
10	0.1081 (0.0595, 0.2455)	0.1081 (0.0400, 0.3305)	0.0270 (0.0040, 0.1633)	0.0270 (0.0021, 0.1497)

Note: The values in this table are the responses over the first 10 years to a 1-standard deviation shock. In parenthesis, the corresponding 95% confidence bands.

Figure A2: Impulse response functions. Time plots





Note: The red lines indicate the estimates of d, initially for a sample ending at 2000, and then adding recursiveely one observation (year) at a time. In red, the values corresponding to the 95% confidence bands.