Structural Breaks in Inflation Dynamics within the European Monetary Union

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Abstract

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1. Introduction

Ever since the beginning of the European Union, the topic of a common currency was a controversial issue. Although the Economic and Monetary Union (EMU) is now a fact, the discussion about the economic effects of the euro is far from being settled. The controversial topics range from the question of whether or not the eurozone is indeed an optimal currency area (OCA) all the way to the very survival of the euro in light of the budgetary problems of some of its member states. The effects of monetary unions on a number of macroeconomic indicators, with inflation being the most important one in this context, is in the center of an ongoing debate: this regards the question about short-run and steady-state inflation uncertainty (in context to inflation expectations) dealt with in Caporale and Kontonikas (2009), or the structural convergence of the inflation rates in EU countries, which is the topic of an investigation by Palomba, Sarno, and Zazzaro (2009).
The way towards the EMU basically consisted of 3 stages: stage I (1990-1994) as a phase of liberalization, stage II (1995-1998) a phase of convergence and stage III (1999-2001) the transition period, which ended with the introduction of the euro as legal tender. A possible fourth stage is the continuing integration of new member states, mostly former communist countries. Any new state that wants to join the EMU is first obliged to fulfill the Maastricht criteria, which, besides setting rules for government debts and interest rates, also requires the participation in the ERM II (Exchange Rate Mechanism II) for two years, while the exchange rates towards the euro are not allowed to cross the nominal band.
Emerson, Gros, Italianer, Pisani-Ferry, and Reichenbach (1992) put emphasize on the fact that a high inflation rate is also more variable and uncertain and thus causes more relative price variability, leading to a less efficient price mechanism. Failing to stabilize inflation leads to severe economic problems. Throughout the literature, there is still a considerable degree of uncertainty as to what extent the introduction of the euro, or monetary unions in general, affect the inflation rate. The EMU requires a country to possibly alter a number of economic policies and institutions, which may demand for significant short term adjustments to take place.

Given that a country experienced quite volatile inflation rates, its efforts to meet the convergence criteria are likely to lead to an alteration at least in the mean (since this was one of the requirements) and possibly in the volatility of their respective inflation rates as well.
However, a deeper inquiry into this problem is in need of a method able to capture changes in inflation dynamics. We contribute to this debate by developing a new method for testing for structural breaks in the dynamics of the inflation rates and thereby upon the effect of the EMU on a number of European countries both within and without the eurozone. The framework developed in this paper makes it possible to test for any significant changes of a countries inflation dynamics and thus helps interpreting the effects of the EMU. We focus on changes in the first three moments of these inflation rates, which we assume to be distributed according to a generalized logistic distribution (GL). The GL distribution is a suitable one due to its flexibility to capture nonsymmetry and because of its fatter tails. For this purpose, we develop a new estimation approach building on previous research by Zeileis and Hornik (2007).
This framework allows us to assess the evolution of mean, variance and skewness over time. The skewness parameter turns out to be quite important to measure changes in the long run inflation dynamics, as it changes considerably over the time of the analysis. This trend would be all but invisible using a normal distribution. Our estimation technique enables the simultaneous estimation of breaks in the mean (which is important in the context of the Maastricht criteria), the variance (which should be low for the price mechanism to be efficient) and the skewness (where a change in skewness is evidence of a shift of the distribution towards a new "center of gravity") of the inflation rates of interest. This than can be used to search for answers regarding the question about the effect of the EMU on the inflation dynamics of its member states or more general to investigate the effects of external shocks.
The remainder of this paper is structured as follows: Section 2 presents the data, Section 3 presents the model and the estimation techniques used, Section 3.5 illustrates our approach using Slovenia as an example, Section 4 presents the results and Section 5 concludes.

2. Data

We employ seasonally unadjusted HICP series of 21 countries ranging from January 1990 to December 2010 that are obtained from the Organisation for Economic Cooperation and Development (2010). Countries included are Austria, Belgium, Czech Republic, Denmark, Estonia, Finland, France, Germany, Greece, Hungary, Ireland, Italy, Luxembourg, Netherlands, Poland, Portugal, Slovakia, Slovenia, Spain, Sweden, and United Kingdom. Latvia and Lithuania as well as Bulgaria and Romania are excluded due to data scarcity. Cypria and Malta are not included since they are very small economies.

The countries in this sample can be divided into three different groups. First the euro countries (Austria, Belgium, Estonia, France, Finland, Germany, Greece, Ireland, Italy, Luxembourg, Netherlands, Portugal, Slovakia, Slovenia and Spain). Second, the EU members not participating in the ERM II: Czech Republic, Hungary, Poland, United Kingdom, and Sweden. Denmark stands on its own as a member of the EU and the ERM II, but not yet a member of the EMU.

As shown in a paper by Lünnemann and Mathä (2009), the harmonization of the treatment of sales prices led to considerable changes in some countries (Belgium, Italy, Luxembourg and Spain); failing to take this into account may lead to considerable distortions in the estimation. In these cases, we dealt with the inclusion of the sales periods effects[1] into the HICP series (January 2000 for Belgium, January 2001 for Italy, January 1999 for Luxembourg and January 2001 for Spain) by estimating dummies for these month[2] and subtracting them from the original time series.

[1] This was demanded by a Commission regulation, details may be found in: European Commission (2010) and in the Appendix B.

[2] There are two month with a significant decrease in prices (January and July), followed by significant price increases in the following months (February and August).

3. Model

To monitor changes in the dynamics of the inflation rates, we follow the usual strategy employed when testing for structural breaks. First, the model is fitted for the whole sample using the generalized logistic distribution. This then enables testing for the stability of the parameter estimates. If there is evidence of significant changes, the number and the timing of the break points are estimated.
This framework allows to trace the evolution of mean, variance and skewness of the inflation series and thus interpreting deviations in any of these parameters as possibly due to a change in monetary policy, is made feasible. Any huge intervention or change is captured by fluctuations in the parameter stability of the series and can be quantified and related to a distinct time frame.
The correlation dynamics of the series is of less importance to this framework, which concentrates on changes in mean, variance and skewness and does not focus on autocorrelation, the use of alternative tests, as described in Andreou and Ghysels (2002), is also hindered by the low number of observations, due to the monthly frequencies.[3]

3.1. Generalized Logistic Distribution

The data suggested that the use of a normal distribution might cause some estimation problems, since most of the time series exhibit asymmetric properties (with sometimes rather strong skewness) and – compared to the normal distribution – a fatter tail. To accommodate for this fact, the use of the generalized logistic distribution (GL) with its relative low degree of complexity, seemed appropriate. On the one hand, it is a simple distribution and on the other hand, it is flexible enough to capture essential properties of the data. For the inflation series $y_t = 100 \cdot log(HICP_t/HICP_{t-1})$ we therefore assume a GL distribution as defined in Johnson, Kotz, and Balakrishnan (1995):

$$f(y|\theta, \sigma, \delta) \quad = \quad \frac{\frac{\delta}{\sigma} \cdot \exp^{-\frac{y-\theta}{\sigma}}}{(1 + \exp^{-\frac{y-\theta}{\sigma}})^{(\delta+1)}} \tag{1}$$

with location θ, scale σ and shape δ. For $\delta = 1$ the distribution simplifies to the logistic distribution, for $\delta < 1$ it is skewed to the left and for $\delta>1$ it is skewed to the right (see Appendix A for more details).
In econometrics, the logistic distribution is often used in income distributions and growth models. This is due to its fatter tails, which fit these kind of data somewhat better. Wong and Bian (2005) develop a maximum likelihood estimator if the underlying distribution is a GL distribution. A similar GL distribution is also used in Tolikas, Koulakiotis, and Brown (2007) who analyze extreme risk and value–at–risk in the German stock market, although they do not use a shape parameter. Regarding inflation rates, the GL distribution is — to our best knowledge — only used in context with expected inflation.
Batchelor and Orr (1988) use a logistic distribution (not its generalization) to model the distribution of mean expected inflation rates.

[3] In some years the Billion Prices Project of the Massachusetts Institute of Technology (2010), which provides a much higher data frequency, could solve this problem.

3.2. Estimation

Under the assumption that y_t is distributed according to the generalized logistic distribution, we test for the hypothesis that the parameter vector $\hat{\phi}$ stays stable over time:

$$H_0 : \phi_t = \phi_0 \quad (t = 1, ..., n)$$

against the H_1 of changes in at least one of the parameters. We first estimate the parameters of the GL–distribution by means of maximum likelihood $\hat{\phi} \;=\; \underset{\phi}{argmax} \sum_{t=1}^{n} \log f(y_t|\phi)$. The parameter stability is assessed using the empirical scores $s(y_t|\hat{\phi})$ as measures of model deviation.[4] The empirical fluctuation process *efp*$(\cdot)$ captures deviations from a hypothesized zero mean over time and is given by:

$$efp(t) \;=\; \hat{V}^{-1/2} n^{-1/2} \sum_{i=1}^{\lfloor nt \rfloor} s(y_i|\hat{\theta}, \hat{\sigma}, \hat{\delta}) \quad (0 \leq t \leq 1),$$

where $\hat{V}$ is some consistent estimator of the variance of the scores, which is used to decorrelate the series as to avoid problems due to serial correlation. Then a functional central limit theorem (FCLT) for *efp*$(\cdot)$ is employed, which converges to a 3-dimensional Brownian bridge:

$$efp(\cdot) \;\xrightarrow{d}\; W^0(\cdot)$$

3.3. Test

After the *efp*$(\cdot)$ is computed, it conveys information about departures from the H_0 of parameter stability over time. For a precise quantification of the actual deviation from its hypothesized zero–mean path, the main idea of the test statistics presented here is to aggregate the *efp*$(\cdot)$ in such a way as to be able to construct a test statistics so that critical values and p values can be derived. In this paper, we employ the Supremum of LM (supLM) test by Andrews (1993). This test is well suited for single break alternatives and performs better if more than just one of the distribution parameters change.

[4] for details see Appendix A and Zeileis and Hornik (2007)

It is given by:

$$\sup_{t \in [0.1, 0.9]} \frac{\| efp(t) \|_2^2}{t(1-t)}$$

where the appropriate p–values can be computed using:

$$\sup_{t \in [0.1, 0.9]} \frac{\| W^0(t) \|_2^2}{t(1-t)}$$

The supLM test computes a test statistics for all possible change points in a given time frame and rejects the null of no change if the maximum is too high.

3.4. Breakpoint Estimation

The dating of structural breaks builds upon the work of Bai and Perron (2003). If a break is detected via the supLM test, the next step is estimating the number and the timing of these break points. Breakpoints $\tau_1, ..., \tau_B$ are estimated via maximization of the full segmented likelihood:

$$\sum_{b=1}^{B+1} \sum_{t=\tau_b+1}^{\tau_b} \log f(y_t | \phi^{(b)})$$

If the test statistics is significant (in this case the supLM test), at least one break is estimated. In principle, the optimal segmentation (i.e. the concrete dates of a break) can be computed once B (the number of breakpoints) is known. However, B needs to be based on the observed data as well. The solution is to compute the timing of the breaks for a number of different values for B and then choose B by means of an information criterion, in this case the LWZ (Liu, Wu and Zidek) as developed in Liu, Wu, and Zidek (1997). All parameters $\tau_1, \ldots, \tau_B, \phi^{(1)}, \ldots, \phi^{(B+1)}$ can be estimated jointly using dynamic programming.[5]

[5] More details for this estimation technique are provided in Zeileis, Shah, and Patnaik (2010).

3.5. An Example

To illustrate our method, we consider the case of Slovenia. After a period of relative instability following the independence from Yugoslavia, Slovenia was successfull in realigning its economy and introduced a number of reforms leading to a stable growth in the past years. The good economic performance made it possible for Slovenia to enter the ERM II in June 2004 and later on to introduce the euro in January 2007. The data collected contains the period from January 1996 onwards. In the first few years, Slovenia experienced very high inflation rates, but was successfull in reducing inflation to a much lower level.

A glance at the time series depicted in Figure 1 of the monthly inflation rates reveals a decline in the mean around the middle of the observed time period. The $efp(\cdot)$ starts to cross the border as early as 2001 and reaches the peak somewhere between 2003 and 2004. The p–value of the supLM test is 0.00014 and clearly indicates at least one break. The number of breaks is calculated by means of the LWZ criteria, which chooses one break in August 2003.[6] In this year, Slovenia introduced a number of financial reforms and severely decreased the year on year growth in the M3 monetary aggregate. Combined, these actions were enough to ensure the successfull participation in the ERM II a year later and the introduction of the euro two years later.

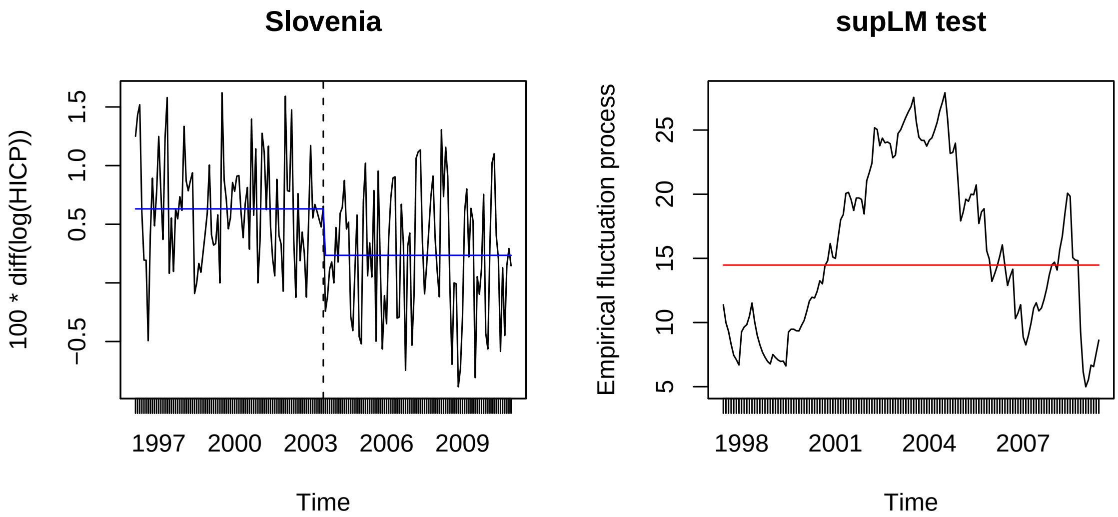

Figure 1: Slovenia: breakpoint estimate and supLM test

[6]For the changes in the parameters, see Table 1.

4. Results

Country	Segment	Mean	Variance	Skewness	ERM	ERM II	Euro
Austria and Belgium							
Austria	Feb 1990–Sep 2007	0.164	0.057	0.609	Jan 1995	–	Jan 1999
	Oct 2007–Dec 2010	0.167	0.149	−0.076			
Belgium	Feb 1991–Sep 2007	0.133	0.069	0.066	Mar 1979	–	Jan 1999
	Oct 2007–Dec 2010	0.159	0.181	−0.686			
The Eastern countries							
CzechRepublic	Feb 1995–Jul 1998	0.697	0.337	1.139	–	–	–
	Aug 1998–Dec 2010	0.177	0.206	0.979			
Estonia	Feb 1996–Mar 1998	0.865	0.420	0.404	–	Jun 2004	Jan 2011
	Apr 1998–Dec 2010	0.336	0.195	0.732			
Hungary	Feb 1995–May 1998	1.606	1.024	0.870	–	–	–
	Jun 1998–Dec 2010	0.491	0.310	0.708			
Poland	Feb 1996–May 2001	0.855	0.422	0.669	–	–	–
	Jun 2001–Dec 2010	0.202	0.120	−0.471			
Slovakia	Feb 1995–Feb 2004	0.563	0.338	1.139	–	Nov 2005	Jan 2009
	Mar 2004–Dec 2010	0.178	0.081	1.139			
Slovenia	Feb 1996–Jul 2003	0.631	0.211	0.588	–	Jun 2004	Jan 2007
	Aug 2003–Dec 2010	0.235	0.331	0.199			
Italy and Spain							
Italy	Feb 1990–May 1996	0.414	0.042	0.970	Mar 1979	–	Jan 1999
	Jun 1996–Dec 2000	0.168	0.020	0.726			
	Jan 2001–Dec 2010	0.466	0.143	0.620			
Spain	Feb 1992–Dec 2000	0.282	0.057	0.709	Jun 1986	–	Jan 1999
	Jan 2001–Dec 2010	0.476	0.136	0.425			
Ireland							
Ireland	Feb 1995–Jun 2008	0.257	0.202	−0.719	Mar 1979	–	Jan 1999
	Jul 2008–Dec 2010	−0.148	0.148	−1.891			
No change countries							
Greece	Feb 1995–Dec 2010	0.325	1.434	0.424	Mar 1998	Jan 1999	Jan 2001
Netherlands	Feb 1990–Dec 2010	0.178	0.292	0.606	Mar 1979	–	Jan 1999
France	Feb 1990–Dec 2010	0.155	0.073	−0.090	Mar 1979	–	Jan 1999
Finland	Feb 1990–Dec 2010	0.166	0.132	0.166	Oct 1996	–	Jan 1999
The 90s							
Portugal	Feb 1990–Jul 1992	0.853	0.173	1.139	Apr 1992	–	Jan 1999
	Aug 1992–Dec 2010	0.229	0.144	0.307			
Sweden	Feb 1990–Jan 1993	0.475	0.570	1.139	–	–	–
	Feb 1993–Dec 2010	0.155	0.183	0.542			
UK	Feb 1990–Apr 1992	0.569	0.385	1.138	–	–	–
	May 1992–Dec 2010	0.168	0.147	−1.175			
Former DM–zone							
Denmark	Feb 1990–Jun 2000	0.166	0.091	−0.745	Mar 1979	Jan 1999	–
	Jul 2000–Dec 2010	0.162	0.174	1.007			
Germany	Feb 1995–May 2000	0.088	0.060	0.924	Mar 1979	–	Jan 1999
	Jun 2000–Nov 2003	0.109	0.156	0.997			
	Dec 2003–Dec 2010	0.160	0.169	0.097			
Luxembourg	Feb 1995–Dec 1998	0.088	0.013	0.251	Mar 1979	–	Jan 1999
	Jan 1999–Dec 2010	0.240	0.177	−0.848			

Table 1: Break point, parameter estimates and ERM information

Table 1 presents the break dates and the parameter values of the segments as well as the dates of entry into the ERM/ERM II and the date of the euro introduction. We formed seven groups of countries according to the observed pattern of break points. These seven groups consist of Austria and Belgium, the Eastern European countries, the two big southern European countries, Italy and Spain, Ireland, the countries with no change in the inflation dynamics, the countries with breaks in the early 90s and the member countries of the former DM–zone (excluding Austria and Belgium).

4.1. The countries in detail

Regarding Austria and Belgium, the time after the break is characterized by an increase in the mean and a severe increase in the variance of the inflation rates. The skewness is now negative for both series, which indicates that the distribution has shifted to the right, implying that overall inflation is likely to increase. This change can be traced back to two major influences: the financial crisis and potentially some spillover effects of higher oil prices in late 2007.

The second group consists of the Eastern European countries: Czech Republic, Estonia, Hungary, Poland, Slovakia and Slovenia. In almost all of these countries, the mean and the variance of their inflation rates declined in the later part of the 90s (Slovenia is an exception with an increase in variance after the break). Most of them experienced a break in 1997/1998, with Poland, Slovakia and Slovenia somewhat later. The break during the late 90s is a result of these countries efforts to curb in inflation by decreasing money supply and indicates the time when these countrie's actually overcame the biggest transitionally shocks on their way towards free market economies. Slovakia is very similar to Slovenia, as its inflation dynamics too changed roughly one year prior to its entry into ERM II.

Another combination of countries with rather identical results are the two biggest southern European countries, Italy and Spain. Both of them experienced high rates of inflation in the early 90s. Afterwards, in the phase leading up to the monetary union, mean and variance were greatly reduced. Soon after the fixing of the exchange rates, this trend changed significantly, leading to a higher mean and a much higher variance. A possible reason for this is that the fiscal responsibility that was required for an entry into the eurozone was softened after this goal was achieved.

Ireland can be considered a rather special case. In Ireland, compared to other countries, the deflationary effect of the financial crises was more persistent. This deflationary trend, which for a time affected all countries, can be seen in Figure 2, which depicts a rolling mean estimation, where the mean is calculated using the first moment of a fitted GL distribution. There is evidence of a decline in mean inflation, which returns to its previous path roughly a year later. The break in 2008 can be attributed to the strong contraction triggered by the bursting of the Irish housing bubble and a beginning deflation in concert with the massive costs of the bailout of the countries banking system. Surprisingly, this deflationary pressure is not visible in Spain, which was affected by a housing bubble of roughly the same size as Ireland, however in Spain housing prices did not adjust so abruptly and in such a degree.

Another interesting group consists of four economically rather different countries. In none of them do we find any evidence of a change in inflation dynamics. Whereas France, Finland and Netherlands are on the lower side of the inflation margin, the Greek inflation rate is among the highest.
The next group consists of two countries that decided not to be part of any European Monetary System[7], the United Kingdom and Sweden. Portugal is included due to the same break point behavior. All three show a clear break in the early 90s. Two of them can be traced back to economic crisis, the currency crisis of the United Kingdom cumulating into the "Black Wednesday" in 1992 and the run on the Swedish currency in the same year. In the case of Portugal, the inflation rate after 1992 was much more stable, showing that the inflationary problems of Portugal were less severe thereafter.
The last group consists of the former DM–zone countries Denmark, Germany and Luxembourg. In the case of Denmark we observe an increase in variance after 2000 with a very distinct change in skewness. In Germany, the inflation rate increased both on average and in its variability. Luxembourg suffered a severe increase in variance, similar to Italy and Spain. All of these countries, which profited from the stability of the previous German currency, still have a very low inflation rate.

4.2. Overview over the general trend

The disparity of the national inflation rates is still an issue. Many researchers, like Hofmann and Remsperger (2005) find a considerable amount of inflation differentials. Caporale and Kontonikas (2009) estimate short-run and steady-state inflation uncertainty in 12 EMU countries and find a considerable degree of heterogeneity across EMU countries in terms of average inflation and its degree of persistence. In a paper examining structural convergence of inflation rates in EU countries, Palomba *et al.* (2009) try to answer the question if during the 1990s the inflation rate dynamics of EU countries become more similar. They find that convergence over time of inflation dynamics was only partly observable. In a paper studying core inflation and using an aggregated euro area inflation rate, Morana (2000) finds three regimes (roughly 1980—1984, 1984—1993 and 1993—2000) governing the core inflation rate. Concerning inflation differentials, Figure 2 depicts the inflation rates and a rolling mean estimation, where the mean is calculated using the first moment of a fitted GL distribution.[8] We do see that many of these countries follow a rather similar path. One event clearly visible is the financial crises and the short deflationary phase starting somewhere around the autumn of 2007 and lasting a few months. Over time, there are hints, that the inflation rates converged towards some common – albeit broad – mean level. For the inflation series of the whole EMU, as published by the Organisation for Economic Cooperation and Development (2010), we see an increase in mean from 0.12166 (1990(1)–1998(12)) to 0.1799 (1999(1)–2008(8)) with a strong decrease to 0.09364 (2008(9)–2010(12)) following the financial crises. Whereas variance has increased over time from 0.01838 to 0.09364. Regarding skewness, we see a clear change towards a negative skewness from 0.37783 to -0.77342.

[7]If we ignore the short participating time of both countries in the EMS.

[8]If more countries have the same inflation rate in a given month, the lines are overlapped and thus get darker. We wold see one black line if all countries would show the same month–on–month increase at any one time. The red lines depict the means.

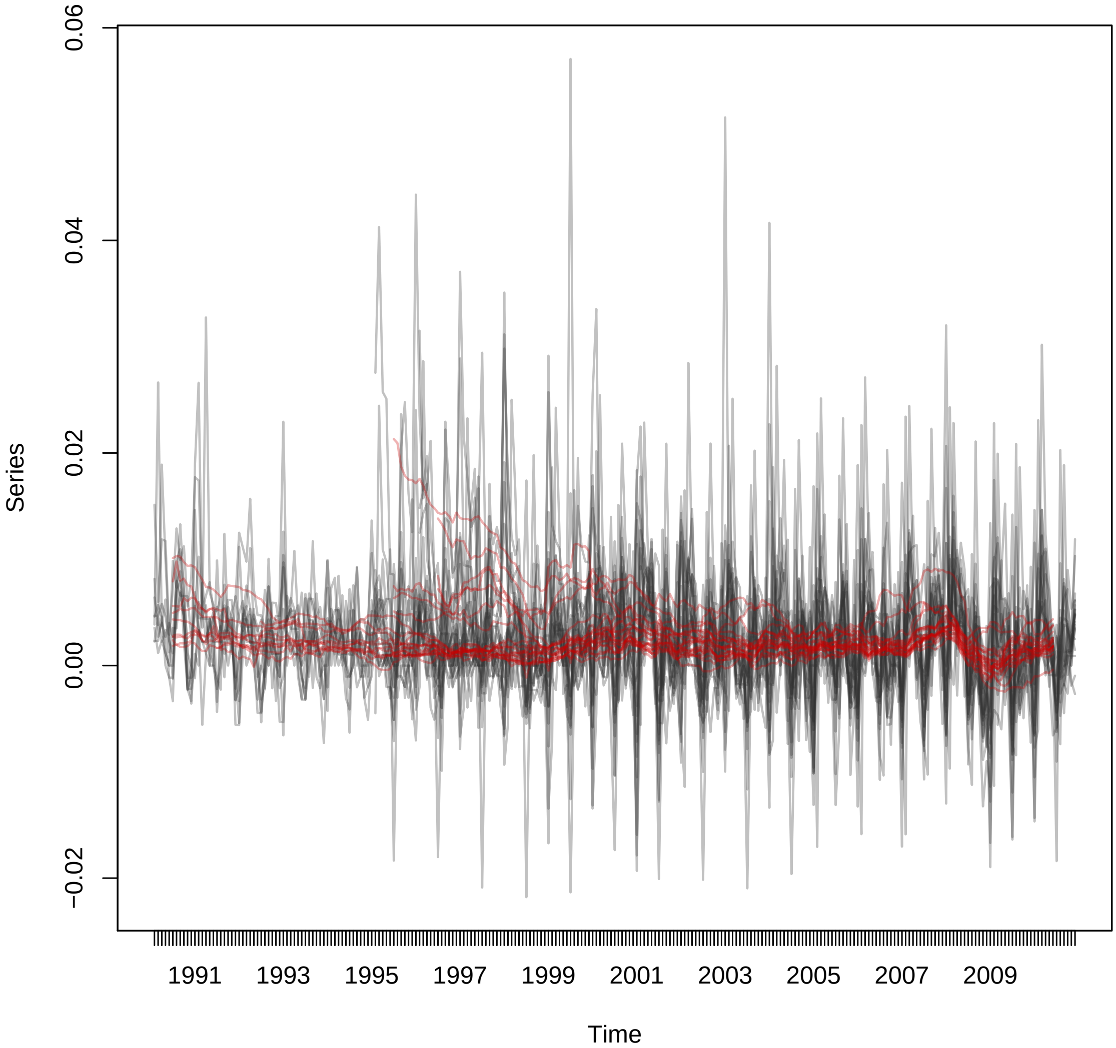

Figure 2: Inflation rates and estimated GL mean over time

Regarding the general trend, Table 2 presents lower, medium and upper (25, 50, 75) percentiles of the first three moments of the inflation series during the convergence phase (January 1990 until December 1998), the euro phase (January 1999 until August 2008) and the time after the financial crisis (since September 2008[9] until now), for the countries in the "core" of the eurozone (Austria, Belgium, Finland, France, Germany, Luxembourg and Netherlands), the PIIGS countries (Portugal, Italy, Ireland, Greece and Spain) and all the others.

Mean				
Zone	Percentile	1990(1)–1998(12)	1999(1)–2008(8)	2008(9)–2010(12)
Coreeuro	lower	0.113762	0.145448	0.065735
	medium	0.158311	0.150341	0.081703
	upper	0.172854	0.175015	0.112206
PIIGS	lower	0.282234	0.261603	0.035135
	medium	0.333546	0.271044	0.261645
	upper	0.445814	0.406342	0.332714
Others	lower	0.273841	0.174693	0.098347
	medium	0.626036	0.347267	0.125227
	upper	0.701675	0.455454	0.212070
Variance				
Zone	Quantil	1990(1)–1998(12)	1999(1)–2007(12)	2008(1)–2010(12)
Coreeuro	lower	0.055030	0.078013	0.130967
	medium	0.059632	0.131523	0.144926
	upper	0.080581	0.147817	0.208511
PIIGS	lower	0.064332	0.116015	0.165294
	medium	0.114927	0.156991	0.195874
	upper	0.184572	0.209940	0.229170
Others	lower	0.244471	0.182797	0.109261
	medium	0.357771	0.251990	0.148045
	upper	0.592824	0.344797	0.200331
Skewness				
Zone	Quantil	1990(1)–1998(12)	1999(1)–2007(12)	2008(1)–2010(12)
Coreeuro	lower	0.069759	0.005264	−0.596240
	medium	0.212504	0.237900	−0.299947
	upper	0.577825	0.260541	0.080041
PIIGS	lower	−0.158046	0.227354	−0.212959
	medium	0.541909	0.387669	0.273159
	upper	0.657313	0.888931	0.292769
Others	lower	0.734147	0.044526	−0.095666
	medium	0.982951	0.943455	0.168359
	upper	1.740923	1.453635	0.390828

Table 2: Development of first three moments

[9] After the insolvency of the Lehman Brothers in September 2008.

Regarding the mean, we see a clear evolution towards a lower mean in all groups, where the core countries show much lower values at the median. Almost the opposite is true concerning the variance of the inflation rates of the core countries, that has increased to a much higher value than at the start of the sample, with the upper percentile during the 90s now even smaller than the lower percentile for the time after the financial crisis. The same is true for the PIIGS countries, although the core country median variance is still lower; the variance for the other countries has declined significantly. Concerning skewness, we see a clear trend in the core countries for skewness to become negative, indicating that the bulk of the values now lies right to the mean. The same is true – although to a lesser extent – for the PIIGS and the other countries.

5. Conclusion

For the countries in the core of the eurozone, the most disturbing trend seems to be the rise in inflation volatility in almost half of the euro countries. A very volatile inflation rate will likely contribute to greater macroeconomic instability. Although Grier and Perry (2000) find no evidence that higher inflation uncertainty raises the average inflation rate – at least in the case of the USA – Jarociǹski (2010) finds a positive correlation between the level and the standard deviation of inflation. This finding might be still more problematic if we take into account that on average the skewness parameter has decreased substantially in most of the euro–zone countries indicating that in future higher inflation rates are more likely. This will put pressure on the ECB to achieve its inflation target of 2%.

Computational Details

The results shown are computed with a new R package **glogis**, which is an enhancement to the already existing R package **strucchange**, which currently does not support a GL distribution. The tests and the graphical illustration of the empirical fluctuation process were developed in Zeileis and Hornik (2007), the second part of the results - the dating procedure and the illustration of the densities fitted for the subsamples (divided by the breaks) - can be found in Zeileis *et al.* (2010).

A. GL Distribution

The moments of the GL distribution as given in equation 1 are:

$$E(y) = \theta + \sigma(\gamma(\delta) - \gamma(1)) \quad (2)$$

$$Var(y) = \sigma^2(\gamma'(\delta) + \gamma'(1)) \quad (3)$$

$$Skew(y) = \frac{\gamma''(\delta) - \gamma''(1)}{(\gamma'(\delta) + \gamma'(1))^{3/2}} \quad (4)$$

where $\gamma'(\cdot)$ and $\gamma''(\cdot)$ are the first and the second derivatives of the digamma function, respectively.
The log-likelihood is:

$$\begin{aligned} \log f(y|\theta,\sigma,\delta) &= \log(\delta) - \log(\sigma) - \frac{1}{\sigma}(y-\theta) - (\delta+1) \\ &\times \log(1 + \exp^{-\frac{y-\theta}{\sigma}}) \end{aligned}$$

The resulting score function, where the scores are the derivatives of the log–likelihood function, has three components $(s_\theta, s_\sigma, s_\delta)$, with $\tilde{y} = \exp^{-\frac{y-\theta}{\sigma}}$. These are employed in equation 2 for the construction of the test statistics and are a means of measuring the estimation error in a maximum likelihood framework.

$$\begin{aligned} s_\theta(y|\theta,\sigma,\delta) &= \frac{\delta \log f(y|\theta,\sigma,\delta)}{\delta\theta} \quad (5) \\ &= \frac{1}{\sigma} - (\delta+1) \cdot \frac{\frac{1}{\sigma}\tilde{y}}{(1+\tilde{y})} \end{aligned}$$

$$\begin{aligned} s_\sigma(y|\theta,\sigma,\delta) &= \frac{\delta \log f(y|\theta,\sigma,\delta)}{\delta\sigma} \quad (6) \\ &= -\frac{1}{\sigma} + \frac{1}{\sigma^2}(y-\theta) - (\delta+1) \times \frac{\frac{1}{\sigma^2}(y-\theta)\tilde{y}}{(1+\tilde{y})} \end{aligned}$$

$$\begin{aligned} s_\delta(y|\theta,\sigma,\delta) &= \frac{\delta \log f(y|\theta,\sigma,\delta)}{\delta\delta} \quad (7) \\ &= \frac{1}{\delta} - \log(1+\tilde{y}) \end{aligned}$$

The scores of the parameters $\phi = (\theta, \sigma, \delta)$ sum up to 0: $\sum_{t=1}^{n} s(y_t|\hat{\phi}) = 0$. The first order condition is given by:

$$\sum_{i=1}^{n} s(y_i|\hat{\theta}, \hat{\sigma}, \hat{\delta}) = 0 \tag{8}$$

B. Data and Sales Periods

In addition to Belgium, Italy, Luxembourg and Spain, there are a number of other countries, where the sales periods are regulated by law. France has an individual time span regulated by each department. This should pose no problem, since there is no global state level solution. In Greece there are also two set sales–periods in winter and summer. Portugal has two set sales periods as well, but for two month (January and February as well as Mid–July to Mid–September). The source for this information is European Consumer Centres (2010). According to information by their national statistical offices, the official dates of the inclusion of these sales periods are: January 2000 for Belgium and January 1999 for Luxembourg, while Italy and Spain introduced them in January 2001. Greece and Portugal were not estimated, since – as in the case of France – these periods either do not seem to be set for the whole country at once, or there does not seem to be additional seasonality as a result of sales periods.

C. Graphs

Here we present the actual data where the blue lines in the graphs correspond to the estimated mean of the series.

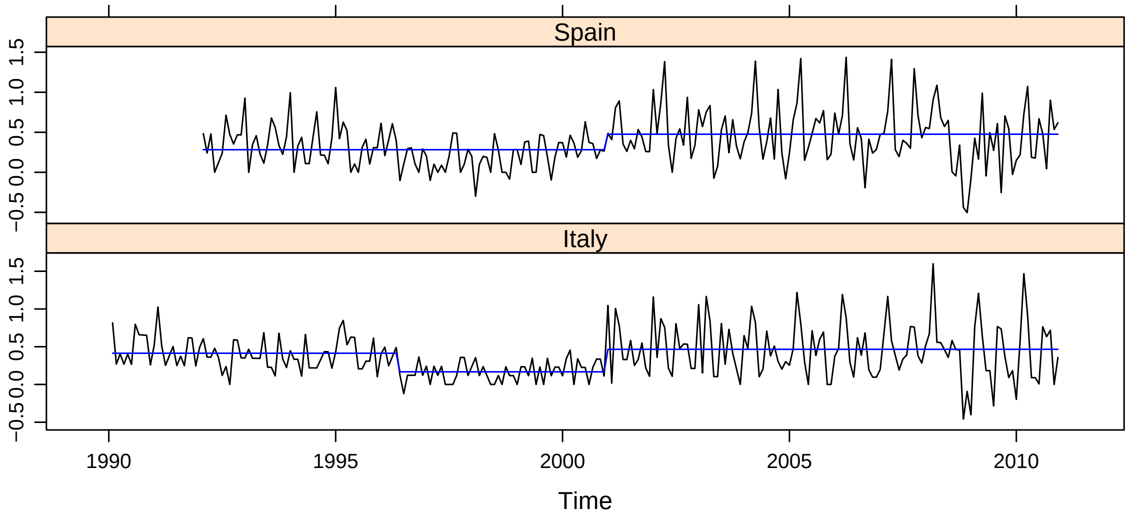

Figure 3: Italy and Spain

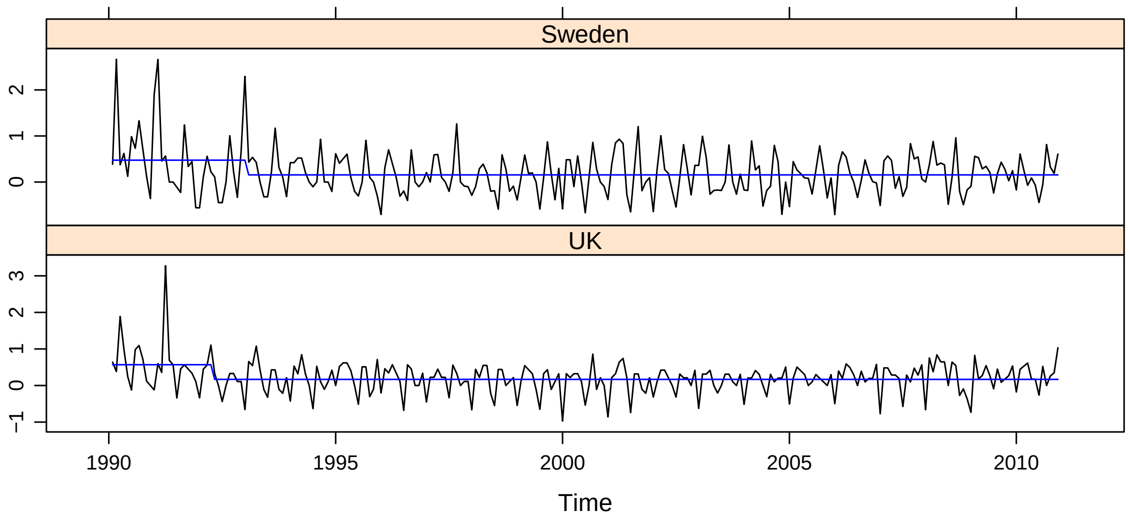

Figure 4: Sweden and United Kingdom

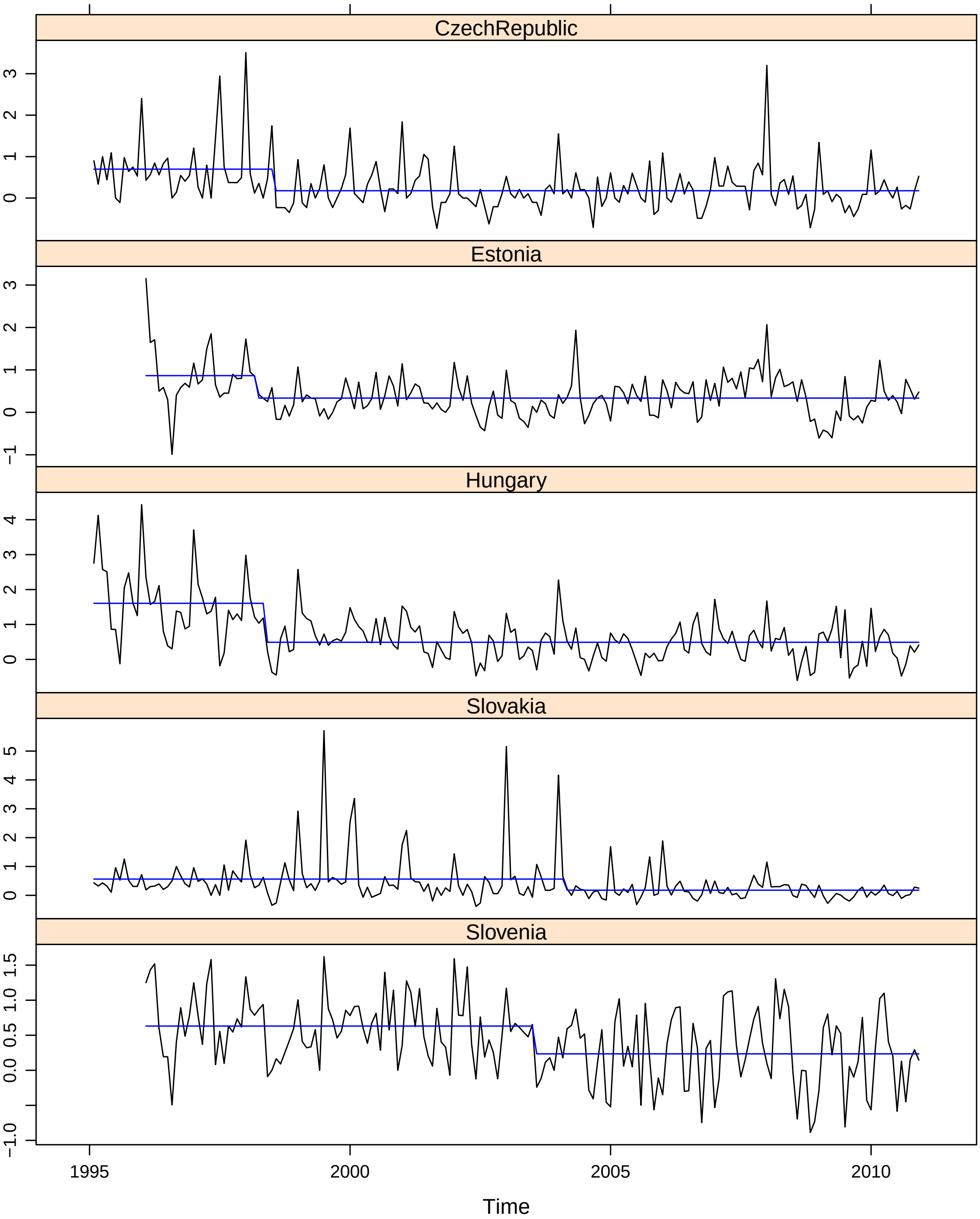

Figure 5: The Eastern Countries

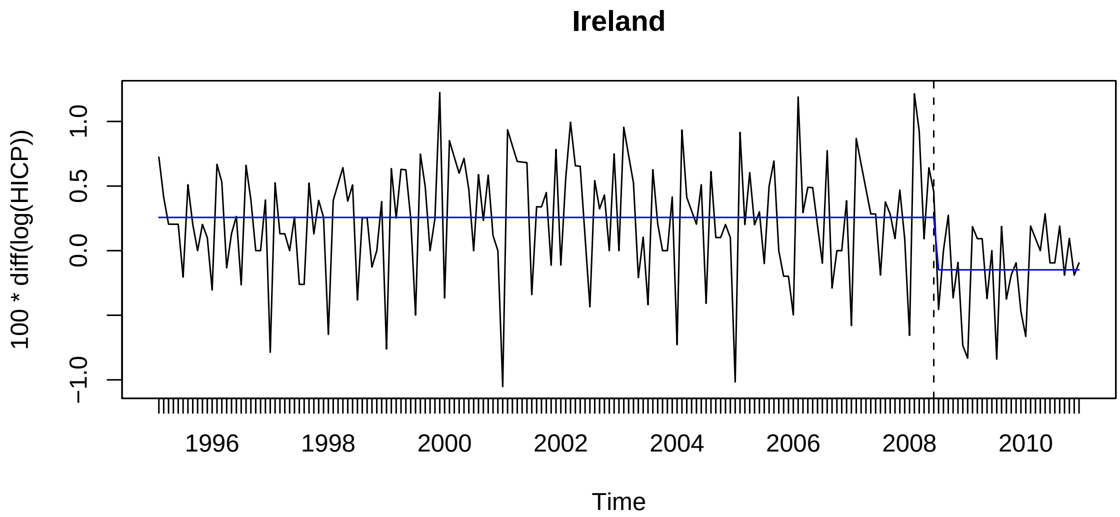

Figure 6: Ireland

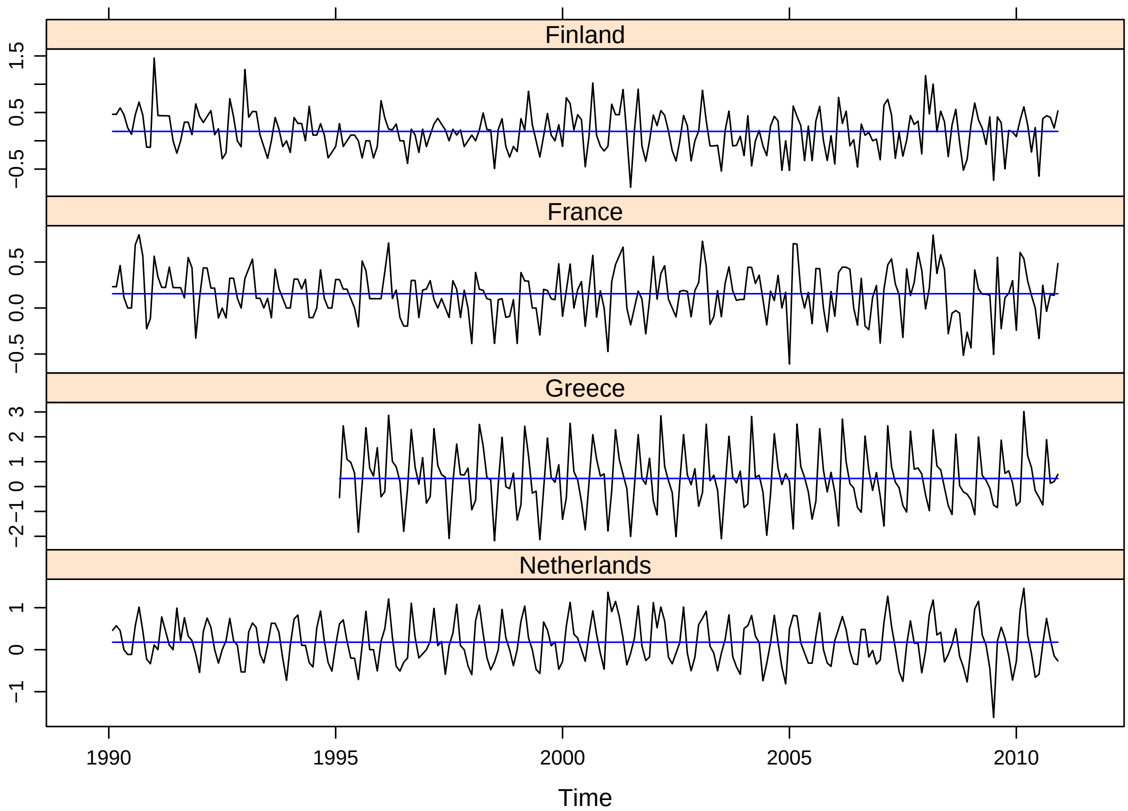

Figure 7: Nochange Countries

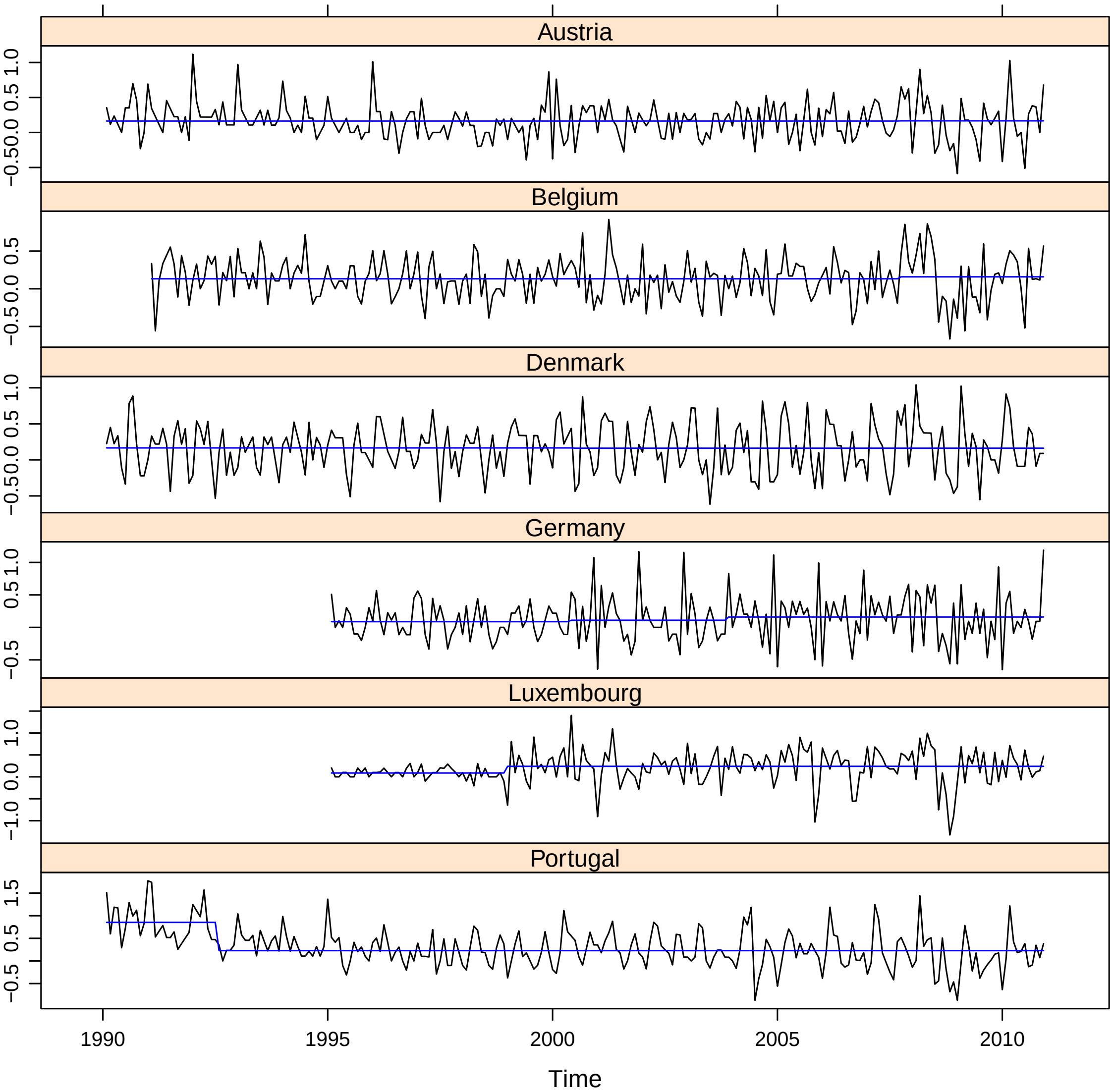

Figure 8: Countries without discernible pattern

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