Structural Breaks in Inflation Dynamics within the European Monetary Union

Thomas Windberger
Universität Innsbruck

Achim Zeileis
Universität Innsbruck

Abstract

The aim of this paper is to investigate how a countries decision to join one or all of the steps leading towards the common European currency changed its inflation rate dynamics. Our attempt is to model the inflation rates via a Generalized Logistic Distribution (GL) and then test for breaks in the parameters of the distribution.

Keywords: inflation rate, structural break, EMU, Generalized Logistic Distribution, EURO.

1. Introduction

The European Central Bank (ECB) defines price stability "as a year-on-year increase in the Harmonized Index of Consumer Prices (HICP) for the Euro area of below 2%"[1]. Emerson, Gros, Italianer, Pisani-Ferry, and Reichenbach (1992) emphasizes 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.

The question of interest centers around the way in which a countries decision to join the Economic and Monetary Union (EMU) changed its inflation rate dynamics. Theory unfortunately does not provide any clear foundation for a proper answer whether or not the creation of a monetary union between two or more states is likely to reduce or increase the variability or even the level of the inflation rate. Cooper and Kempf (2003) points out, that "a central bank under a monetary union will internalize the interdependence between countries and optimally choose a lower inflation rate" and he argues that a "central Bank governing the growth of money supply will optimally choose zero inflation." This is not the case with the ECB which targets at 2% so as to avoid the alleged risk of deflation. It is thus not quite clear how a monetary union will affect the volatility and the level of the inflation rate.

An interesting approach to this question is taken by Holtemöller (2007). What he finds out via simulations of different interest rate rules is that the standard deviation of the home CPI inflation rate can be substantially reduced by joining a monetary union.

The effects of joining a monetary union on inflation rate variability depends on structural parameters like risk aversion, price flexibility, export demand elasticity, openness and shock correlations. Due to the fact that not all of these parameters are known and that their interaction as well has to be estimated, the whole model is very dependend on assumptions.

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

[1] http://www.ecb.europa.eu/mopo/strategy/pricestab/html/index.en.html

in terms of average inflation and its degree of persistence. In a paper examining structural convergence of the inflation rates in EU countries, Palomba, Sarno, and Zazzaro (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.

Inquiring into the convergence properties of inflation rates among countries of the EMU, Busetti, Forni, Harvey, and Venditti (2007) find that from 1980—1997 there was convergence of inflation rates, but afterwards there is some diverging behavior. This does not quite go along with our findings, that do not indicate stronger divergence after 1997 by means of the detection of break points.

In section 2 we will present the data, in section 3 we will present our approach, in section 4 we will present the results, in section 5 we will discuss our findings and try to bring them into context with the overall economic condition, section 6 concludes.

2. Data

Inflation is measured as the monthly change in the HICP rates from 1.1990—3.2010. Countries included are Austria, Belgium, the Czech Republic, Denmark, Estonia, Finland, France, Germany, Greece, Hungary, Ireland, Italy, Luxembourg, the Netherlands, Poland, Portugal, Slovakia, Slovenia, Spain, Sweden, and the United Kingdom.[2] The data is obtained from the OECD Statistics.[3]

The countries can be divided into three different groups: the EURO countries (Austria, Belgium, Estonia[4], France, Finland, Germany, Greece, Ireland, Italy, Luxembourg, the Netherlands, Portugal, Slovakia, Slovenia and Spain)[5], EU members not participating in the ERM II (Exchange Rate Mechanism): the Czech Republic, Hungary, Poland, the 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.

[2] Latvia and Lithuania as well as Bulgaria and Romania are excluded due to data scarcity.
[3] http://www.oecd-ilibrary.org/content/data/data-00285-en
[4] although Estonia has not yet introduced the EURO, but in 2011.
[5] Cypria, Malta, Andorra, Monaco, San Marino and the Vatican are left out due their minor importance.

3. Model

Our main question of interest is whether or not a countries decision to join one or all of the steps leading towards a common European currency changed its inflation rate dynamics. The economic reason why this should be so is rather clear cut for more than half of the countries in our sample. We would expect changes in the dynamics of all Eastern European countries due to their transition from previously centrally planned economies towards free market economies. All of these countries experienced some problems with very high inflation rates at the beginning and they all were successfull in eliminating excess inflation afterwards. It can be safely assumed that the former constituent republics of Yugoslavia follow the same pattern. In the case of the Southern European countries, with a history of high inflation and less pronounced fiscal discipline, we would also expect inflation dynamics to change towards a regime of lower inflation. The case for the Central European countries of course is different. There is no reason why the European Central Bank (ECB) should actually follow the rather stable monetary policies of the former German central bank in a one–to–one fashion.

The approach we use follows one simple idea: i.e. we want to track the evolution of mean, variance and skewness of our series over time. Thereby we hope to be able to find and estimate possible breakpoints in the data. The correlation and the correlation structure of the time series themselves are of no interest and are treated as a nuisance term.

To do this, we assume that our data is a draw from a Generalized Logistic Distribution (GL) and we proceed by estimating the parameters of the GL–distribution via maximum likelihood. We then test for structural change and estimate them using dynamic programming.

3.1. Generalized Logistic Distribution

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. The application of the generalized logistic distribution as we will use it in this paper is rather rare. Wong and Bian (2005) use a GL distribution in a regression model with autocorrelated errors. They use this type of distribution rather than a Student's t-distribution, as to model the fact that these are oftentimes non symmetric and severely left or right skewed. 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 don't use a shape parameter. Regarding inflation rates, the GL distribution is — to our best knowledge — only used in context to expected inflation. Batchelor and Orr (1988) use a logistic distribution (not its generalization) to model the distribution of mean expected inflation rates.

We apply the general framework, as developed in Zeileis and Hornik (2007), to a more specific model, in this case by means of a GL distribution. Prior to that, we would like to give a short justification for the usage of the GL distribution. Regarding the data at hand, it was not possible to utilize the already existing method developed in Zeileis and Hornik (2007), since almost all inflation rates, with the notable exception of Greece, are not normally distributed and clearly exhibit asymmetric properties.

Therefore, a somewhat more flexible distribution had to be used. We needed a distribution exhibiting rather fatter tails and the property to be both left and right skewed. To serve this purpose, we use a generalization of the logistic distribution as defined in Johnson, Kotz, and Balakrishnan (1995).

For our inflation series $y_t = 100 \cdot log(HICP_t/HICP_{t-1})$ we assume a GL distribution given by:

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

with location (θ), scale (σ) and shape (δ). For δ=1 the distribution simplifies to the logistic distribution, for δ<1 it is skewed to the left and for δ>1 it is skewed to the right. The moments are:

$$\begin{aligned} E(y) &= \theta + \sigma(\gamma(\delta) - \gamma(1)) && (1)\\ Var(y) &= \sigma^2(\gamma'(\delta) + \gamma'(1)) && (2)\\ Skew(y) &= \frac{\gamma''(\delta) - \gamma''(1)}{(\gamma'(\delta) + \gamma'(1))^{3/2}} && (3) \end{aligned}$$

where $\gamma(\cdot)$ is the digamma function and its derivatives.
The log-likelihood is:

$$\begin{aligned} \log f(y|\theta,\sigma,\delta) &= \log(\delta) - \log(\sigma) && (4)\\ &- \frac{1}{\sigma}(y-\theta) - (\delta+1)\\ &\times \log(1+\exp^{-\frac{y-\theta}{\sigma}}) \end{aligned}$$

The resulting score function, where the scores are the derivatives of the log–likelihood function, has 3 components $(s_\theta, s_\sigma, s_\delta)$, with $\tilde{y} = \exp^{-\frac{y-\theta}{\sigma}}$. These are used later on for the construction of the test statistics and are a means of measuring the estimation error in a maximum likelihood framework.

$$\begin{aligned} s_\theta(y|\theta,\sigma,\delta) &= \frac{\delta \log f(y|\theta,\sigma,\delta)}{\delta\theta} && (5)\\ &= \frac{1}{\sigma} - (\delta+1)\cdot\frac{\frac{1}{\sigma}\tilde{y}}{(1+\tilde{y})}\\ s_\sigma(y|\theta,\sigma,\delta) &= \frac{\delta \log f(y|\theta,\sigma,\delta)}{\delta\sigma} && (6)\\ &= -\frac{1}{\sigma} + \frac{1}{\sigma^2}(y-\theta) - (\delta+1)\times\frac{\frac{1}{\sigma^2}(y-\theta)\tilde{y}}{(1+\tilde{y})}\\ s_\delta(y|\theta,\sigma,\delta) &= \frac{\delta \log f(y|\theta,\sigma,\delta)}{\delta\delta} && (7)\\ &= \frac{1}{\delta} - \log(1+\tilde{y}) \end{aligned}$$

An enhancement to the already existing R package **strucchange**, which currently does not support a GL distribution is provided with this paper. The asymptotic testing theory still

holds for this generalization.
The first part of the results we are going to present here, i.e., the tests and the graphical illustration of the empirical fluctuation process can be backtracked to 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, Shah, and Patnaik (2010).

3.2. Estimation Framework

In order to analyze our data, we first estimate the parameters of the GL–distribution by means of maximum likelihood, which is possible, if we assume that $\phi = (\theta, \sigma, \delta)$ is stable over time t:

$$\begin{aligned} \hat{\phi} &= \underset{\phi}{argmax} \textstyle\sum_{t=1}^{n} \log f(y_t|\phi), \\ \textstyle\sum_{t=1}^{n} s(y_t|\hat{\phi}) &= 0, \end{aligned}$$

Our problem now simplifies to the following question: Is the assumption valid or do the ϕ_t change?

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

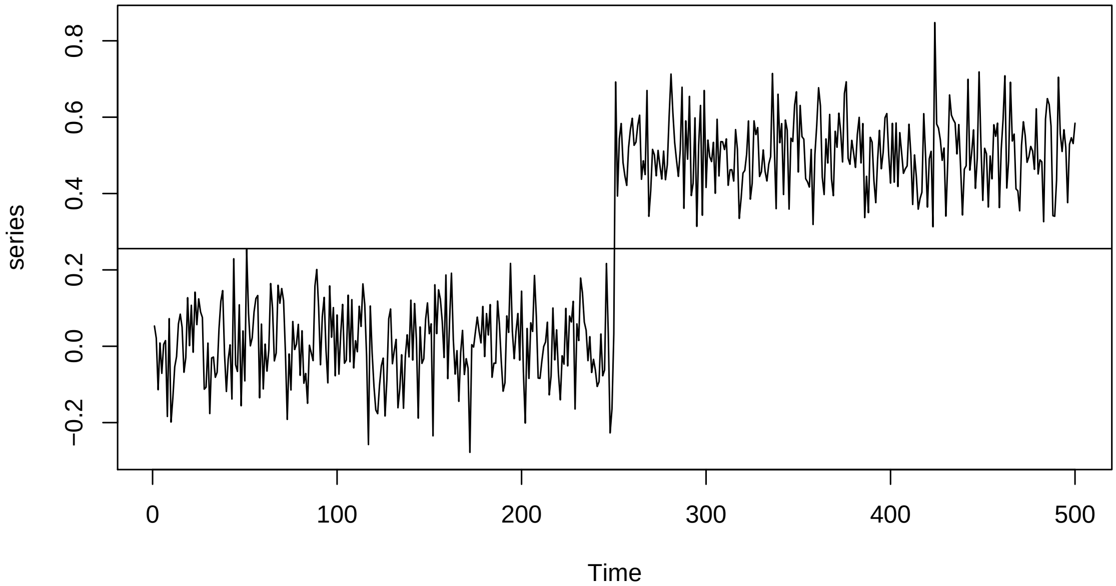
This can be assessed using the empirical scores $s(y_t|\hat{\phi})$, as presented above, as measures of model deviation. The empirical scores process $efp(\cdot)$ captures deviations from zero over time and is given by:

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

where $\hat{V}^{-1/2}$ is some consistent estimator for the variance of the scores, which is used to decorrelate the series as to avoid problems due to serial correlation. We then employ a Functional central limit theorem (FCLT) for $efp(\cdot)$ which converges to a 3-dimensional Brownian bridge:

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

Lets elaborate on this basic idea. Consider some time series given by the following graph. We would not need any test, since we see a clear break at the middle of the sample.



If we estimate a simple least squares regression we get a straight regression line at the mean of the series. What we do now is to sum up the scores (which can be viewed as similar to the residuals) by means of the efp (i.e. something like a cumulated sum of residuals). This summing up of the residuals would be constructed by first aggregating all the negative residuals down to the lowest point, then we proceed by accruing the other (positive) residuals. The bottom of this cumulated sum will be our estimation of the breakpoint. If this peak crosses some test statistics, then it can be viewed as a significant deviation from the null hypothesis.

3.3. Test

Actually the $efp(\cdot)$ could be used for aggregated test statistics in various ways. In our paper we employ the Supremum of LM (supLM) test by Andrews (1993). This test statistic is given by:

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

where the appropriate p–values can be computed using:

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

The supLM test is the most powerful test for a single abrupt change in unknown time and has the highest power if more than one parameter of the distribution changes over time. Throughout the paper we will use the supLM test.

3.4. Breakpoint Estimation

After a break is detected via the supLM test, we proceed by estimating the timing of the break. We estimate B breakpoints $\tau_1, ..., \tau_B$ via maximization of the full segmented likelihood:

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

All parameters $\tau_1, \ldots, \tau_B, \phi^{(1)}, \ldots, \phi^{(B+1)}$ can be estimated jointly using dynamic programming. More details for this estimation technique are provided in Zeileis *et al.* (2010).

4. Results

To help with the interpretation, we give a short overview over the history of the EMU. It 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 (1999-2002) the transition period, which ended with the introduction of the Euro as legal tender and started with the Exchange Rate Mechanism II (ERM II) for Non–members and the irrevocable fixing of the exchange rates of the Member States.
It should be pointed out however, that these monthly series do still contain all the seasonality[6]. If we correct for this by using the lag of order 12 instead of order 1, we find somewhat less pronounced changes with the exception of the latest observations, with the financial crises and the overall lowering in inflation rates, which of course leads to a considerable increase in variance for more than half of our sample. 5 countries change their inflation dynamics after October 2007, coinciding with the ongoing financial crises.

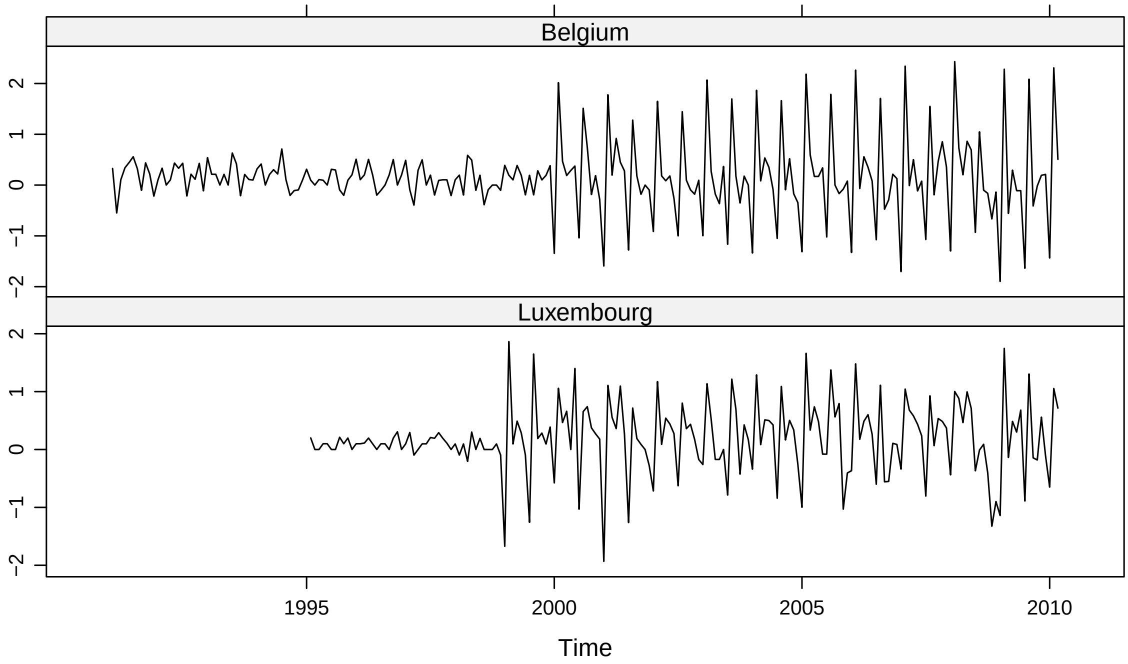
4.1. Belgium and Luxembourg

On the basis of these results, we are able to trace out different groups. The first of this groups consists of Belgium and Luxembourg. This may come as no surprise since both countries formed a monetary union prior to the start of the Exchange Rate Mechanism (ERM). While both countries enjoyed very stable inflation rates with low values of both mean and variance, this changed around the timing of the irrevocable fixing of the exchange rates. But in the case of Belgium and Luxembourg, what we actually observe is a change due to the definition of the HICP. According to an official of the OECD, these changes can be traced back to the inclusion of sales (winter and summer) in January 1999 for Belgium and January 2000 for Luxembourg which was demanded by a Commission regulation.[7] This lead to a 6 month seasonality pattern (January and July) afterwards (meaning after the breakpoints induced by sales inclusion), since sales periods in Belgium and Luxembourg are regulated by law and have to occur in the first two weeks of these months and therefor we observe this sharp price decreases in both month. Interestingly enough, both these price decreases are immediately followed by increases in February and August of even higher magnitude.

Country	Segment	Mean	Variance	Skewness	ERM	ERM II	EURO
Belgium	Feb 1991–Dec 1999	0.146	0.064	−0.037	Mar 1979	–	Jan 1999
	Jan 2000–Mar 2010	0.177	0.954	0.504			
Luxembourg	Feb 1995–Dec 1998	0.088	0.013	0.261	Mar 1979	–	Jan 1999
	Jan 1999–Mar 2010	0.224	0.531	−0.484			

[6] We would like to mention that in the case of Belgium and Luxembourg, the seasonality patterns can be traced back to changes in HICP definition and legal proquirements. In addition to Belgium and Luxembourg, there are a number of other states, 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. Given the probelems concerning the Greek data, this does not pose any addional difficulty. In Italy the sales periods are fixed each year by the Chambers of Commerce. Portugal has two set sales periods as well, but for two month (January and February as well das Mid–July to Mid–September. The source for this information is: *http://www.konsumenteuropa.se/en/Shopping−in−Europe*

[7] for technical details see: http://eur-lex.europa.eu/LexUriServ/LexUriServ.do?uri=OJ:L:2000:300:0016:0017:EN:PDF


Belgium
Luxembourg
2
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1995
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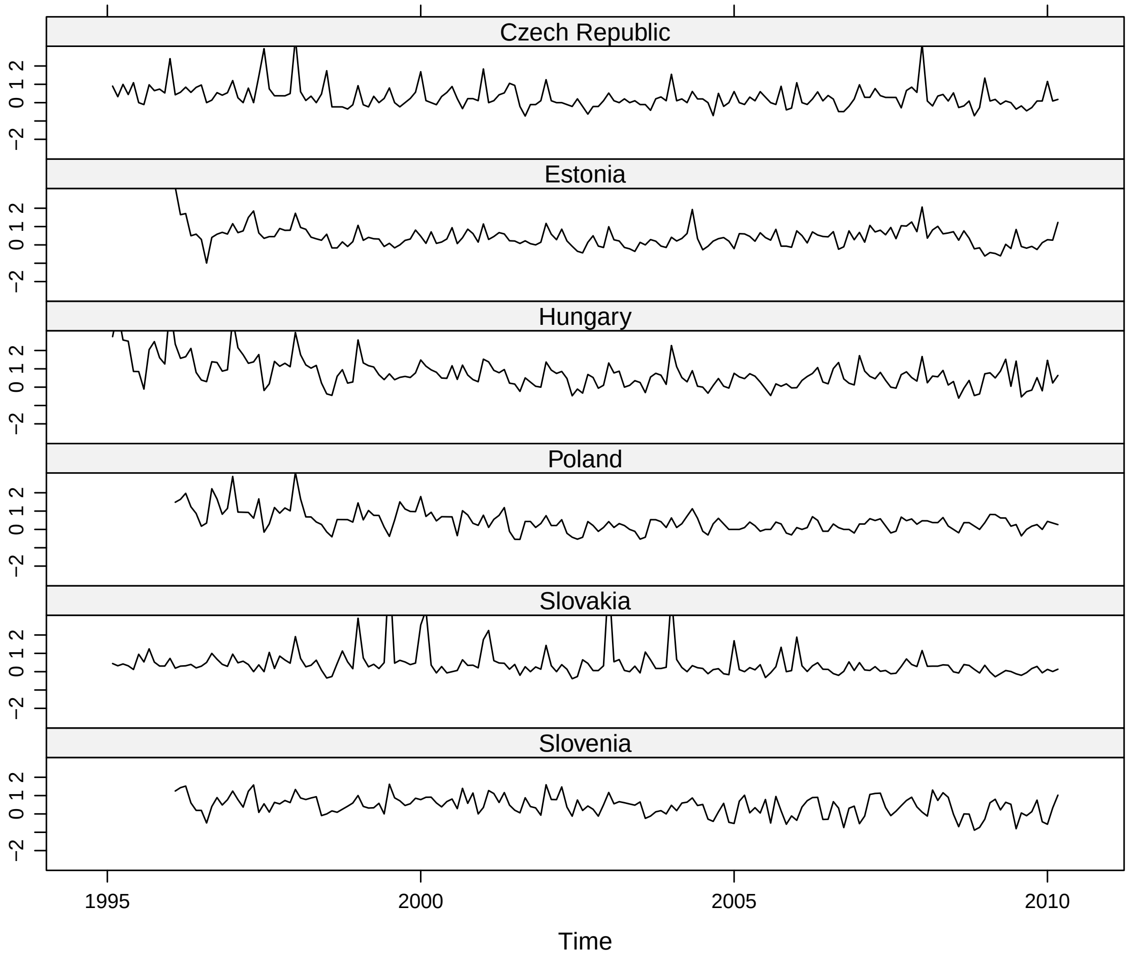
Time

4.2. The Eastern Countries

The second of these groups are the Eastern European countries, the Czech Republic, Estonia, Hungary, Poland, Slovakia and Slovenia. In all of these countries, the mean and the variance of their inflation rates declined in the later part of the 90ies. Most of them experienced a break in 1997/1998, with Poland, Slovakia and Slovenia somewhat later.

Country	Segment	Mean	Variance	Skewness	ERM	ERM II	EURO
CzechRepublic	Feb 1995–Jul 1998	0.697	0.336	1.139	–	–	–
	Aug 1998–Mar 2010	0.182	0.216	0.990			
Estonia	Feb 1996–Mar 1998	0.865	0.420	0.404	–	Jun 2004	Jan 2011
	Apr 1998–Mar 2010	0.333	0.206	0.802			
Hungary	Feb 1995–May 1998	1.606	1.024	0.878	–	–	–
	Jun 1998–Mar 2010	0.507	0.316	0.709			
Poland	Feb 1996–May 2001	0.855	0.421	0.668	–	–	–
	Jun 2001–Mar 2010	0.202	0.123	−0.315			
Slovakia	Feb 1995–Apr 1997	0.480	0.058	1.139	–	Nov 2005	Jan 2009
	May 1997–Feb 2004	0.587	0.442	1.140			
	Mar 2004–Mar 2010	0.186	0.089	1.139			
Slovenia	Feb 1996–Jul 2003	0.631	0.211	0.588	–	Jun 2004	Jan 2007
	Aug 2003–Mar 2010	0.244	0.344	0.143			



Czech Republic
Estonia
Hungary
Poland
Slovakia
Slovenia
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0 1 2
1995
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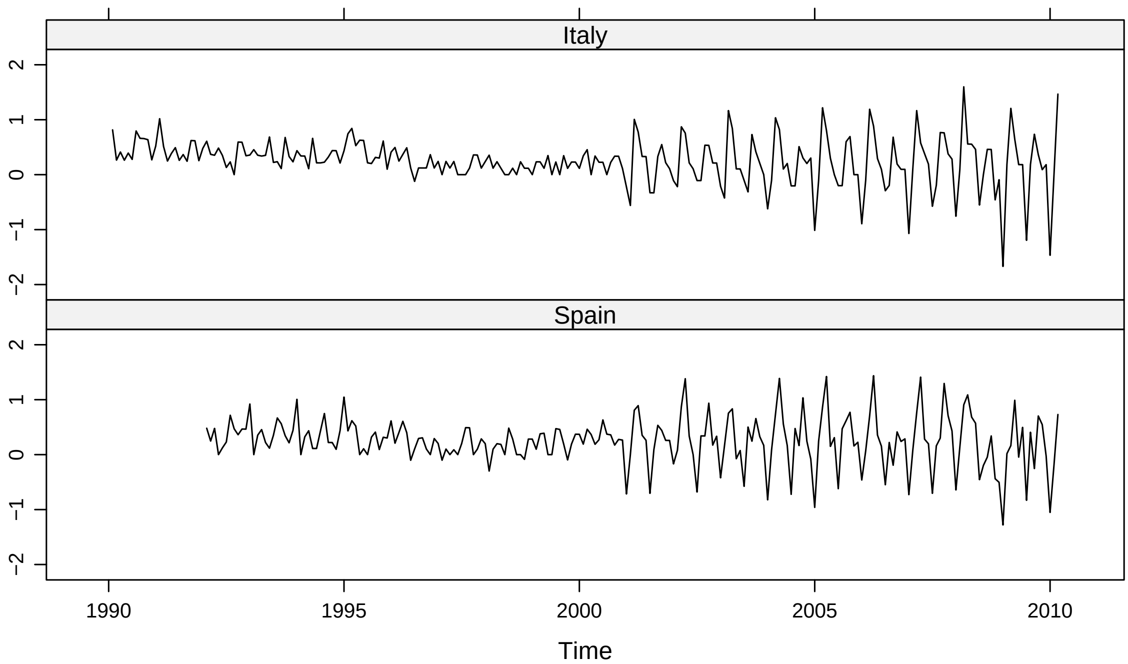
2005
2010
Time

4.3. Italy and Spain

Another combination of countries with – this time – identical results are the two biggest southern European countries Italy and Spain. Both of them experienced high rates of inflation in the early 90ies. Afterwards, in the phase leading up to the monetary union, mean and variance were significantly reduced. This trend changed drastically soon after the fixing of the exchange rates leading to higher mean and much higher variances then ever before since the beginning of the HICP series.
Italy is quite a nice example to explain the usefullness of the skewness parameter. If we focus on the first two periods, with skewness lower than 1 (which indicates that the distribution is skewed to the right) we find that positive shocks are not normally neutralized by negative shocks. This changes in the last phase with skewness negative, indicating that the GL is now skewed to the left and that positive shocks will more likely be followed by negative shocks.

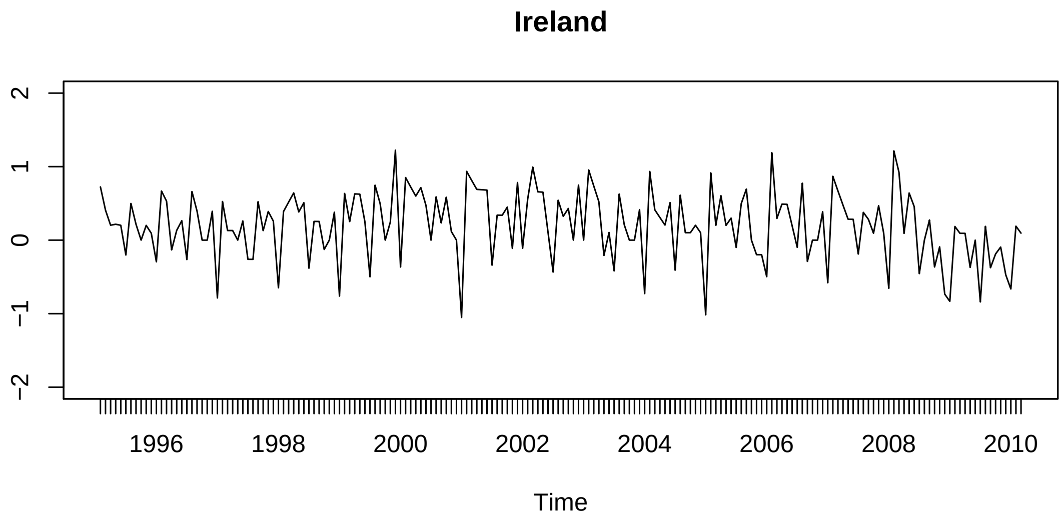
Country	Segment	Mean	Variance	Skewness	ERM	ERM II	EURO
Italy	Feb 1990–May 1996	0.414	0.041	0.963	Mar 1979	–	Jan 1999
	Jun 1996–Dec 2000	0.168	0.020	0.726			
	Jan 2001–Mar 2010	0.182	0.321	−0.261			
Spain	Feb 1992–May 1996	0.372	0.070	1.139	Jun 1986	–	Jan 1999
	Jun 1996–Dec 2000	0.200	0.040	0.019			
	Jan 2001–Mar 2010	0.223	0.342	−0.362			



4.4. Ireland

Ireland can be considered a rather special case. It is the only economy where the effects of the ongoing financial crisis of 2008 are visible. We find a structural break in 2008 following the strong contraction triggerd by the bursting of the Irish housing bubble and a beginning deflation.

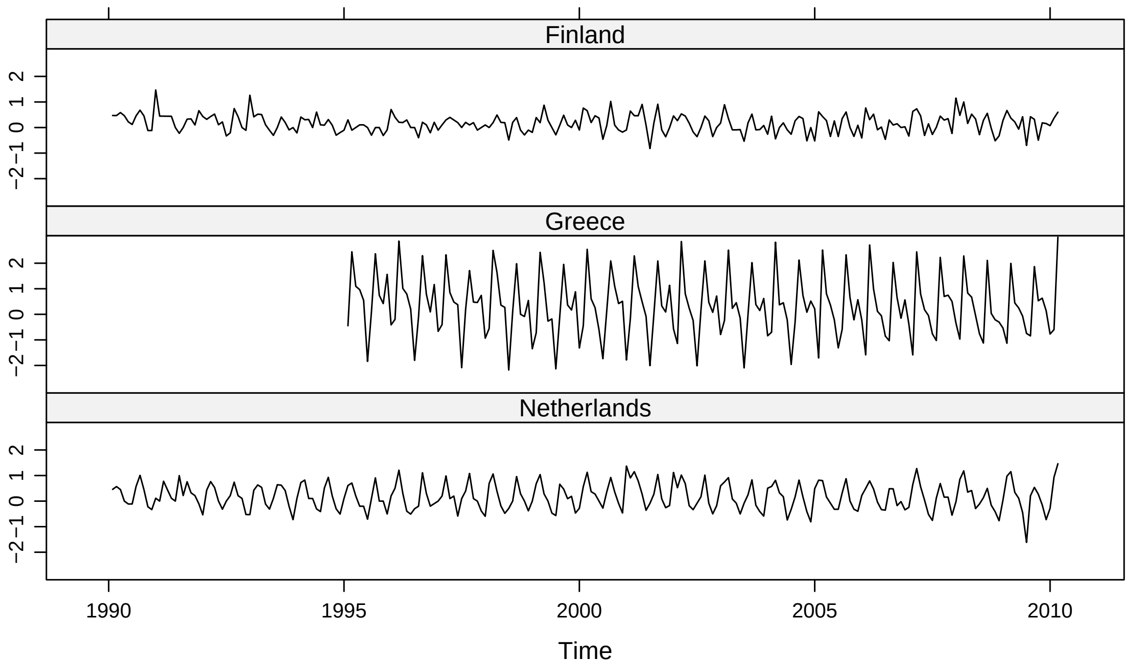
Country	Segment	Mean	Variance	Skewness	ERM	ERM II	EURO
Ireland	Feb 1995–Mar 2008	0.255	0.205	−0.696	Mar 1979	–	Jan 1999
	Apr 2008–Mar 2010	−0.131	0.184	−0.995			



4.5. No change countries

Another interesting group consists of 3 economically rather different countries. In all of them we find no evidence of a structural change. Whereas the Netherlands and Finland are both on the lower side of the inflation margin, Greece exhibits some very strange behaviour. It has by far the highest variance of all series and seems to be completely different from all the others.

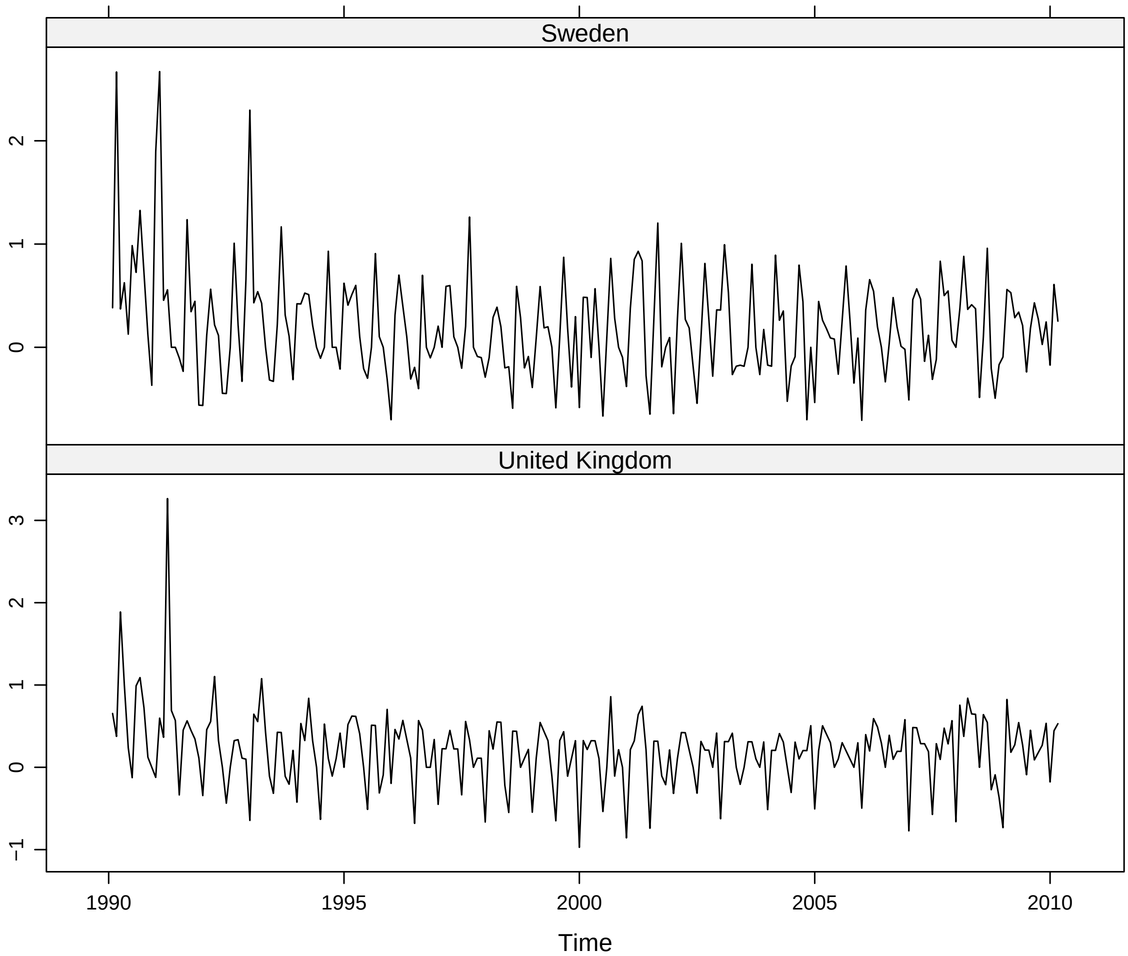
Country	Segment	Mean	Variance	Skewness	ERM	ERM II	EURO
Greece	Feb 1995–Mar 2010	0.323	1.480	0.431	Mar 1998	Jan 1999	Jan 2001
Netherlands	Feb 1990–Mar 2010	0.185	0.293	0.598	Mar 1979	–	Jan 1999
Finland	Feb 1990–Mar 2010	0.165	0.132	0.280	Oct 1996	–	Jan 1999



4.6. The Non–participating Countries

The last group with clear patterns can be seen with the two countries that decided not to be part of any European Monetary System[8], the United Kingdom and Sweden. Both have a clear break in the early 90ies, and both of these breaks can be traced back to 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.

Country	Segment	Mean	Variance	Skewness	ERM	ERM II	EURO
Sweden	Feb 1990–Jan 1993	0.475	0.572	1.139	–	–	–
	Feb 1993–Mar 2010	0.155	0.185	0.534			
UnitedKingdom	Feb 1990–Apr 1992	0.570	0.387	1.139	–	–	–
	May 1992–Mar 2010	0.162	0.149	−1.265			

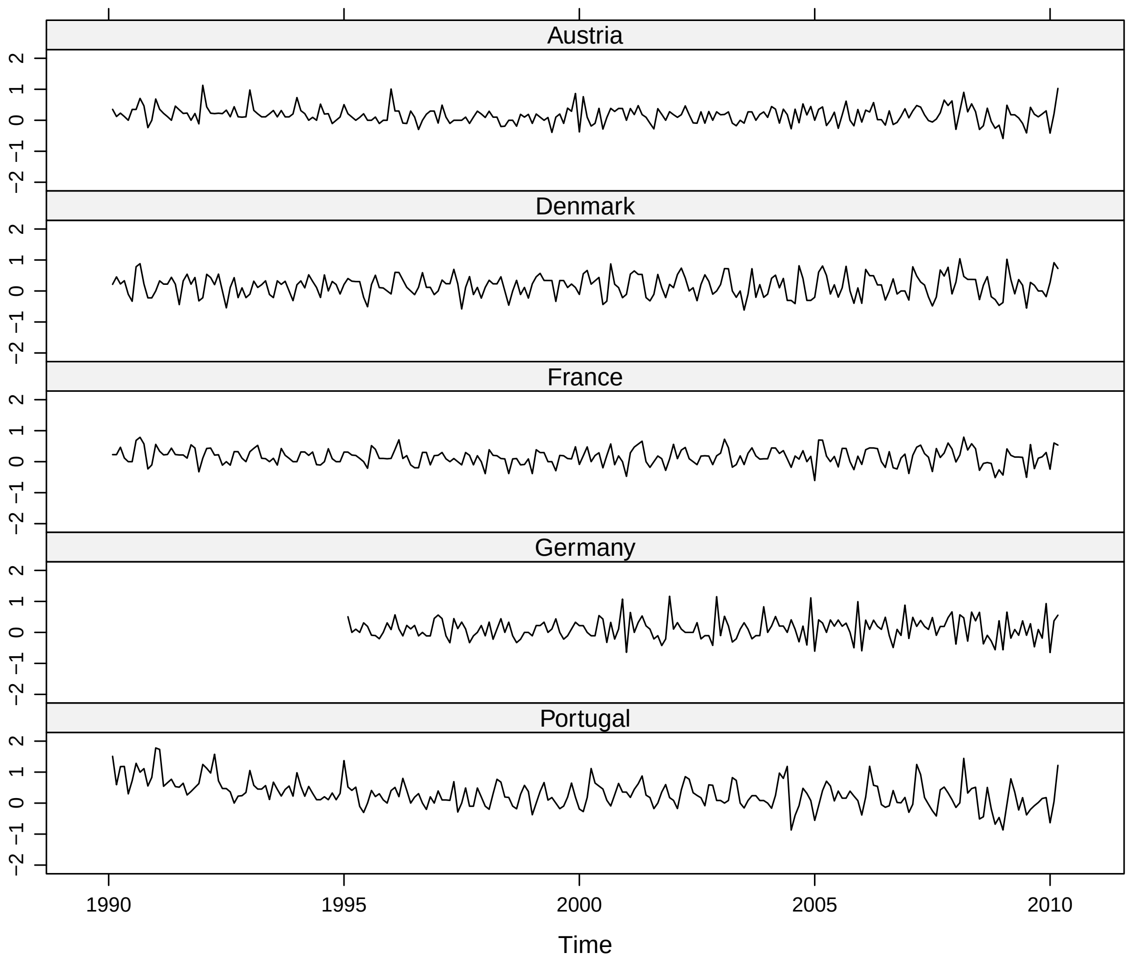


[8]If we ignore the short participating time of the United Kingdom in the EMS

4.7. The Other Countries

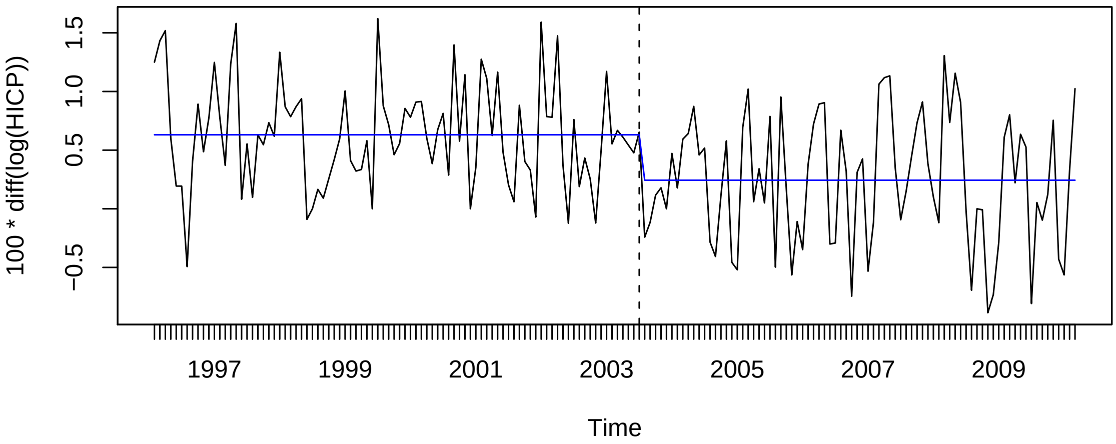
No clear interpretation is available for: Austria, Denmark, France, Germany and Portugal.

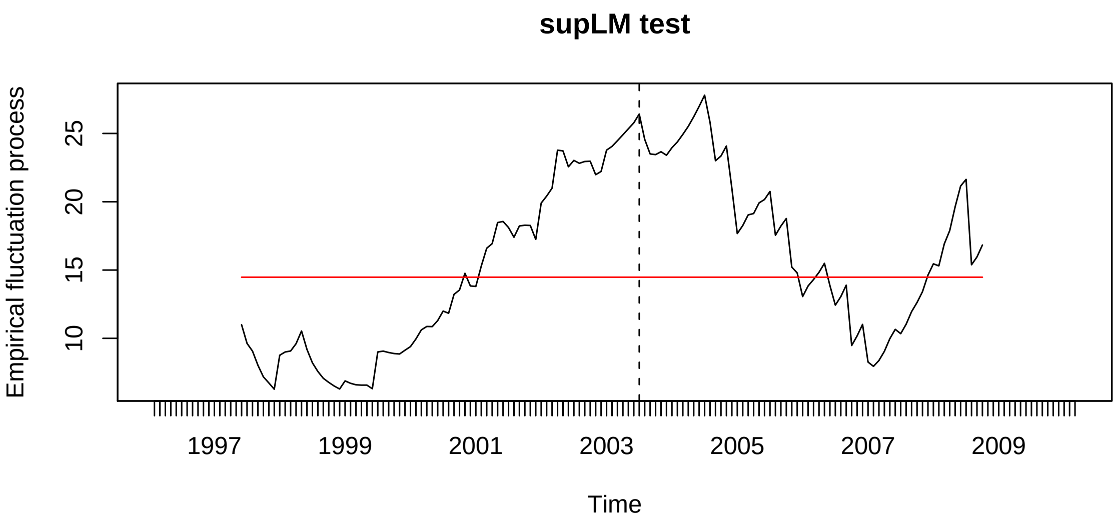
Country	Segment	Mean	Variance	Skewness	ERM	ERM II	EURO
Austria	Feb 1990–Sep 2007	0.163	0.057	0.606	Jan 1995	–	Jan 1999
	Oct 2007–Mar 2010	0.173	0.163	0.169			
Denmark	Feb 1990–Jun 2000	0.166	0.091	−0.742	Mar 1979	Jan 1999	–
	Jul 2000–Mar 2010	0.168	0.188	1.047			
France	Feb 1990–Dec 2004	0.159	0.058	0.196	Mar 1979	–	Jan 1999
	Jan 2005–Mar 2010	0.150	0.131	−0.794			
Germany	Feb 1995–May 2000	0.088	0.060	0.922	Mar 1979	–	Jan 1999
	Jun 2000–Dec 2004	0.140	0.164	0.992			
	Jan 2005–Mar 2010	0.142	0.184	−0.662			
Portugal	Feb 1990–Jul 1992	0.852	0.172	1.139	Apr 1992	–	Jan 1999
	Aug 1992–Mar 2004	0.270	0.105	0.865			
	Apr 2004–Mar 2010	0.161	0.256	0.569			



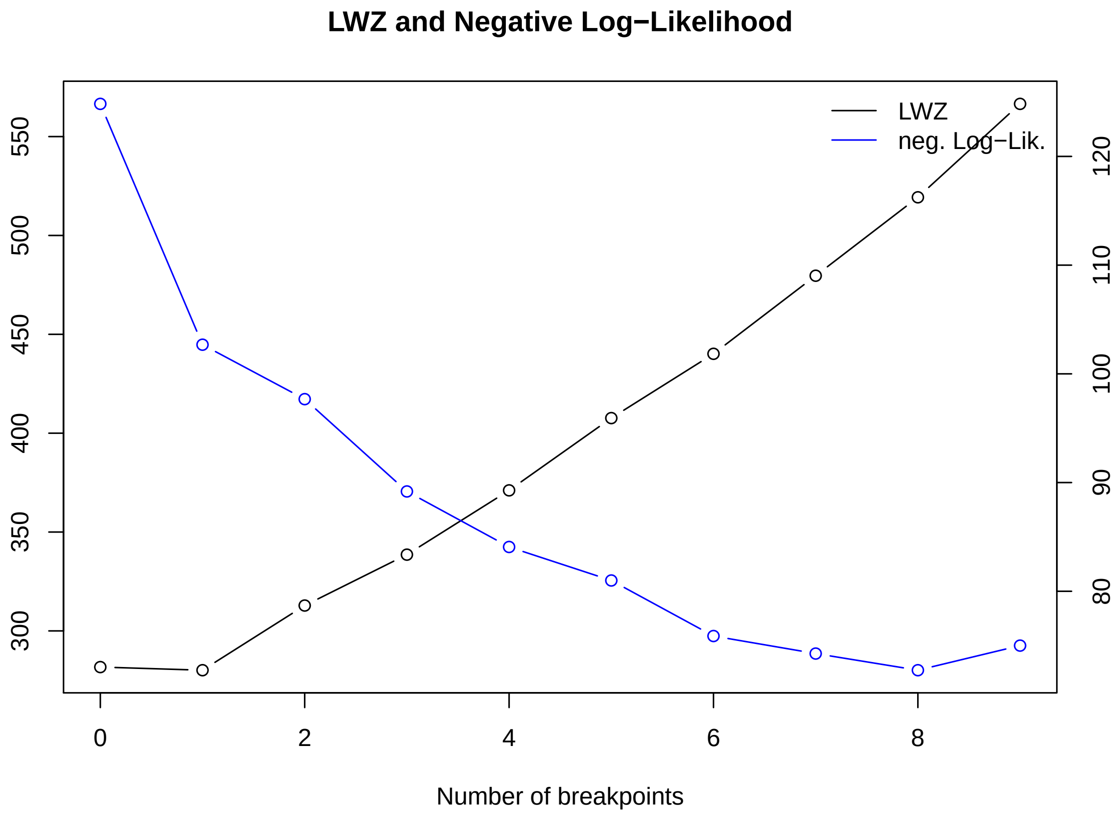
4.8. An example–Slovenia

Let us go through all the steps of our analysis by means of an example, for this we have chosen the case of Slovenia. The economic reason why we suppose to find changing dynamics are clear. In the first phase of the consolidation after the break–up of Yugoslavia and after the short war of 1991, Slovenia experienced very high inflation rates, with yearly growth rates as high as 90% in August 1990. But gradually Slovenia was able to reduce the inflation rate towards much lower values. This first phase is not captured in our data since it starts as late as January 1996. For the period up to today we would expect one or two breaks from an economic point of view. The efforts to reach the Maastricht criteria as soon as possible are likely to lead to a further lowering of the mean and possibly also of the variance of the inflation rate. Another break point might possibly be found towards the introduction of the EURO coins themselves in 2007.
Actually when we look at the data, this is precisely what we get:



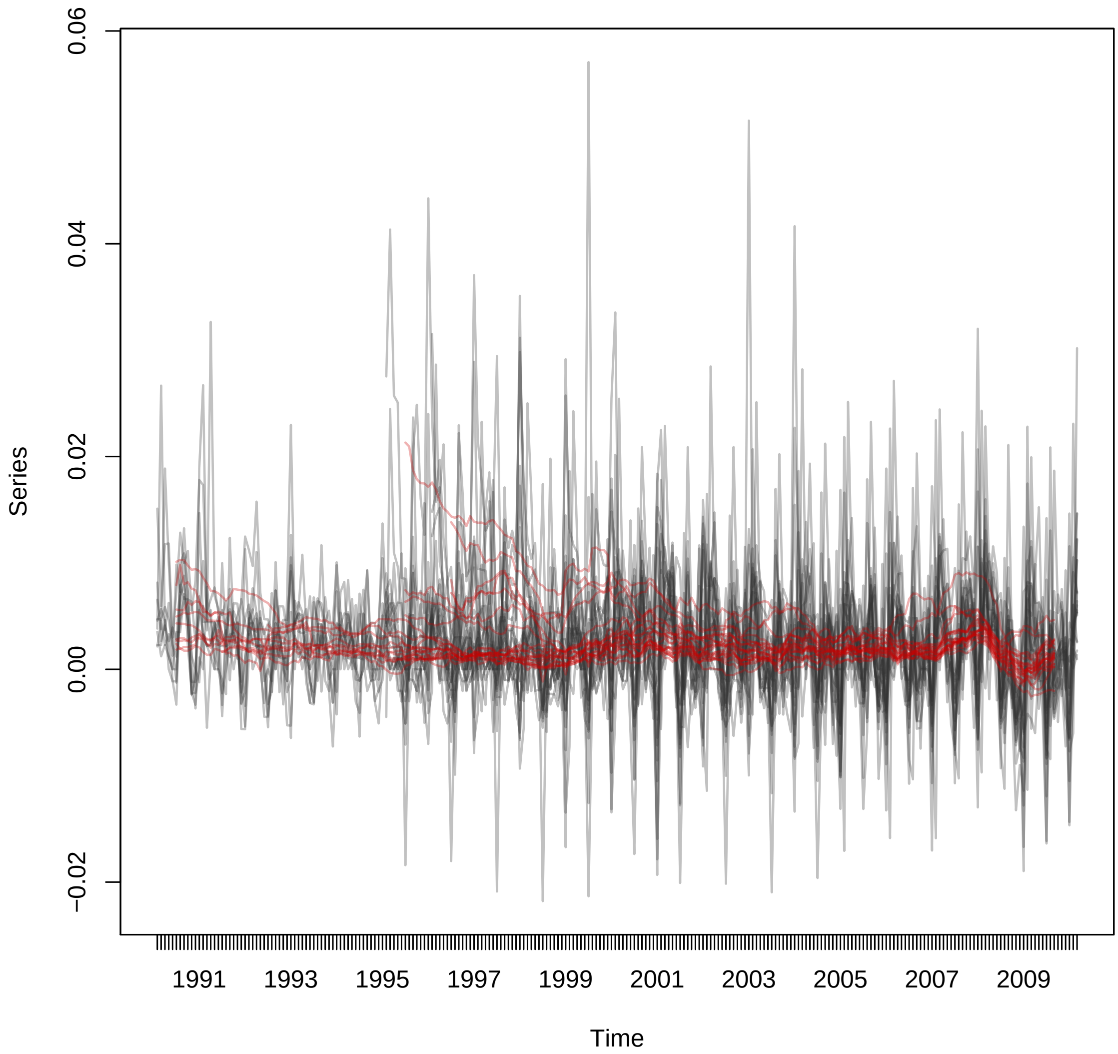


The supLM test graph suggests one break around 2003 and possibly another towards the end of the sample. This second break point is not estimated, since we restricted a regime towards a length of at least 24 month. Thus we find one break in August 2003 with a much lower mean value (only a third now of its previous value) and interestingly enough, a higher volatility.
To be as precise as possible with the detection of break points, we also include a graph showing the values of the LWZ (Liu, Wu and Zidek) criteria as developed in Liu, Wu, and Zidek (1997). Whereas the supLM test suggest at least one break, the LWZ criterion favors no break. But the decline in the negative log–likelihood from 0 breaks to 1 is the strongest, we concluded that one break point actually would be the best alternative.



Slovenia did achieve this by two very important measures. First it liberalized the financial markets in 2003 and – most important – it restricted the year–on–year growth of the monetary base M1 severely starting in 2003. Combined, these actions were enough to ensure Slovenian participation in the EURO–zone.

5. Discussion



To clarify some issues, lets take a look at the data aggregated in just one graph. If more countries have the same inflation rate in a given month, the lines are overlayed and thus get darker.[9] 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. However, this was not enough for most of the countries – with the exception of Ireland – to lead towards a structural break in inflation rate dynamics.

One interesting aspect we would like to focus on is the clear trend in means. What we have

[9]We wold seee one black line if all countries would show the same month–on–month increase at any one time.

plotted here actually is a rolling mean, where the mean is calculated using the first moment of a fitted GL distribution. It is clearly visible that these rates converged towards some common mean (graphically emphasized by the redder lines) which, from 2005 onwards, comes rather close to the ECB target rate (which would translate to a monthly increase of roughly 2 per mill). So concerning the mean of the inflation rates, we see a clear convergence pattern. However, this is not the case for variance where we find clear evidence of an increase. This of course does make some economic sense, since the convergence criteria and the ECB focus on mean inflation only and do not take volatility into account.

6. Conlcusion

We were able to outline changes in the dynamics of the inflation rates in various countries across Europe. However, we fail to find any clear pattern relating to the EURO introduction, but find clear evidence that the zeal of the Eastern European countries to join the European currency at the earliest possible date, clearly helped in reducing mean and variance of their inflation rates.

Considering the severe increase in overall volatility of the European inflation rates, it should be pointed out however, that most of these increases can be traced back to the strong seasonality in our time series.

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Affiliation:

Thomas Windberger, Achim Zeileis
Department of Statistics
Universität Innsbruck
Universitätsstraße 15
6020 Innsbruck, Austria
E-mail: Thomas.Windberger@student.uibk.ac.at, Achim.Zeileis@R-project.org