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Changes in the winter precipitation in Romania and its relation to the large-scale circulation

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ABSTRACT

The variability of winter mean precipitation as observed at 14 Romanian rain gauge stations from 1901–1988 is examined. Pettitt's statistic is used to detect changes of regimes in the time series. Almost all stations exhibit a systematic decrease ("downward shift") at about 1969. Furthermore, upward shifts are identified for the southwestern stations at about 1933, and a downward shift in the mid 1920's in the northwest. An upward shift at about 1919 for the Bucharest station is likely determined by the urbanisation effect. These systematic changes are shown to be real and not an artifact due to inhomogeneities in the precipitation data in a two-step procedure. First, the precipitation field and the European-scale sea-level air-pressure field are related to each other through a Canonical Correlation Analysis (CCA). Two relevant pairs of characteristic patterns are found. In a second step, the CCA-coefficients of these two pairs are studied with Pettitt's statistic. In both pairs of time series, simultaneous change points are found in the precipitation and in the pressure-related coefficients. The 1933 and 1969 change points are related to a change of the southwesterly flow which brings moist Mediterranean air to Romania. The mid-1920s change point is triggered by changes in the frequency or intensity of the northwesterly circulation. As a byproduct, we found that Pettitt's statistic is sensitive to the presence of trends and serial correlation so that its use for statistical hypothesis testing is limited. Therefore, we have used Pettitt's statistic only as an explanatory tool.

1. Introduction

The study of climate variability is important for many reasons. The users from agriculture and hydrology require information about the range of the natural variability. The time scale of the variability varies from months to years or decades. Most presentations of the observed climate variability focus on the surface variability of importance to man, in particular temperature and precipitation and the results depend on the period and the region considered.

Recently, Schönwiese et al. (1994) mapped the trend of seasonal temperature and precipitation in Europe during the past 100 years. They found considerably different trends in different seasons and regions. To understand why the changes occur, it is essential to consider the atmospheric dynamics, as well as the local physical processes. The atmospheric circulation, which is the main forcing for the regional variability of wind, temperature, precipitation and other climatic variables exhibits variability not only on year-to-year time scales but also on decadal scale (Trenberth, 1990; Xu, 1993).

A traditionally used approach for the determination of different "regimes" in time series is to

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apply the concept of "change points" (Sneyers, 1975) which are times of abrupt changes of the statistics of a time series. For the determination of these change points, usually a technique called Pettitt-test (Pettitt, 1979) is used (Sneyers et al., 1994; Boroneant and Râmbu, 1992; Busuioc and Bojariu, 1993). However, the applicability of the Pettitt-test is fundamentally limited by the two assumptions of stationarity and of the lack of serial correlation. We demonstrate in the Appendix that the test's performance is highly sensitive to violations of any of the two conditions. In climate applications, the zero autocorrelation conditions is only rarely satisfied; this problem may be solved to some extent by "prewhitening" (Katz, 1988; Zwiers and Von Storch, 1994; Kulkarni and Von Storch, 1995). The violation of the stationarity condition can hardly be solved. If the time series exhibits a monotonic trend, or if it is composed of several piecewise linear trends, then the Pettitt-test indicates more often than permitted by the significance level the presence of "change points" in cases of no abrupt changes [see Appendix, and Sen and Srivastava (1975), Solow (1987)]. Indeed, the null hypothesis to be rejected by the Pettitt-test is "the time series is instationary", which may be interpreted as "presence of an abrupt change of the mean" if all types of non-abrupt instationarities, such as a piecewise linear trend, can be excluded as culprits because of some additional knowledge unrelated to the data. In general, such knowledge is unavailable, and the possible cure, to subtract all non-abrupt instationarities from the data, can not be used. Another problem with the interpretation of the result of a Pettitt-test is that it is unclear whether a once determined "change point" in the data is due to a change of the dynamical regime or to inhomogeneities in observing, reporting and analyzing the data.

Because of these methodological problems with the traditional "change point"-analysis we use the Pettitt-test in a different set-up. Firstly, we use Pettitt's statistic not as a *confirmatory* tool but only as a *exploratory* tool for the determination of possible regime changes. We make no probability statements (such as the risk of rejecting the null hypothesis of no change point). Secondly, we make no attempts to objectively differentiate between abrupt changes in the mean or to changes in trends. Thus, in this article the expression "change point" refers to changes in the mean or to changes in local trends (with the implicit assumption that the non-stationary component is made up of piecewise linear trends).

Thirdly, to overcome the inhomogeneity problem we search for simultaneous change points in dynamically related time series. For that purpose, the relationship between Romanian precipitation and large-scale circulation is studied with the canonical correlation analysis (CCA). When the regional change may be traced back to a change in a large-scale forcing mechanism, then consistent changes will likely be detectable in other areas. As a parameter to represent the large-scale circulation we use the seasonal mean sea level pressure (SLP) field on the European scale. Then we analyse the time series associated to the most significant CCA pairs of precipitation and SLP to detect "change-points" in both time series. We find that systematic and physically plausible changes of the mean state happen simultaneously in both parameters. Thus the changes in the mean precipitation are real and not due to inhomogeneities such as changes of instrumentation.

The combination Pettitt/CCA is not the only approach we could have pursued. Indeed, an alternative to CAA would have been what is called misleadingly singular value decomposition (Bretherton et al., 1992), and an alternative to the Pettitt-approach would have been Solow's (1987) idea to deal explicitly with piecewise linear trends. However, we are convinced that our results are insensitive to such methodical details.

The paper is organised in the following way. Section 2 presents the data used in this study. A brief description of the Pettitt's procedure and CCA are presented in Section 3. The results obtained from the analysis of the Romanian precipitation and sea level pressure, separately as well as from the simultaneous variation of both the variables are shown in Section 4. The conclusions and discussions are presented in Section 5.

For the sake of brevity, this paper deals exclusively with winter conditions; the analysis for summer reveals similar results (Busuioc and Von Storch, 1995).

2. Data

The data used in this paper are the time series of the winter seasonal precipitation amount at

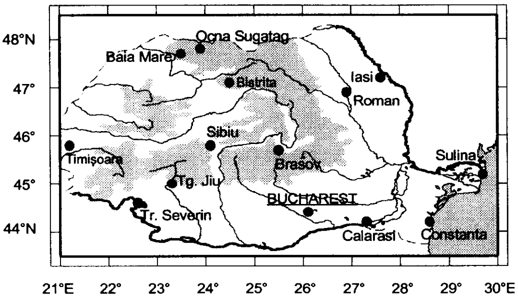

Fig. 1. Position of the 14 stations used in this study. The contours of the Carpathian mountains are marked. The Carpathian mountains as well as the Black Sea are marked by stippling.

the 14 Romanian stations and seasonal mean sea level pressure (SLP) in the 1901–1988 interval during the winter season (December to February). The position and names of the Romanian stations used in this study are shown in Fig. 1. For the SLP the area between 30°N–55°N and 5°E–50°E was selected. The monthly SLP data have been taken from the National Center for Atmospheric Research (USA) with a resolution of 5° × 5° (Trenberth and Paolino, 1980). For both parameters the anomalies have been computed by subtracting the long term seasonal mean from the original values.

3. Methods

3.1. *Pettitt's approach to the change point problem*

Only a brief description about the non-parametric technique to approach the change-point is presented in this section. More details are given by Pettitt (1979) and Sneyers (1975) where the general concept of "change-points", in the sense of abrupt changes of the mean, and their detection are presented.

A sequence of random variables $X_1 \ldots X_T$ is said to have a change point at τ if for $t = 1 \ldots \tau$ all X_t have a common distribution function $F_1(x)$ and for $t = \tau + 1 \ldots T$ all X_t have a common distribution function $F_2(x)$ with $F_1(x) \neq F_2(x)$. We want to test the null hypothesis of "no change of the mean" (if the considered variable is not symmetric then the mean is replaced by the expectation). Therefore we assume $F_1(x) = F_2(x + \Delta)$ and use the null hypothesis

$$H_0 : \Delta = 0$$

or, equivalently

$$H_0 : \tau = T$$

against the alternative of "change"

$$H_A : 1 \leqslant \tau < T.$$

For the testing of H_0 against H_A, the non-parametric statistic

$$K_T = \max_{1 \leqslant t < T} |U_{t,T}| \tag{1}$$

is used; for change in one direction (i.e., $\Delta > 0$ or $\Delta < 0$) the statistics are

$$K_T^+ = \max_{1 \leqslant t < T} U_{t,T}, \tag{2}$$

$$K_T^- = -\min_{1 \leqslant t < T} U_{t,T}. \tag{3}$$

The numbers $U_{t,T}$ are calculated iteratively by

$$U_{t,T} = U_{t-1,T} + V_{t,T}$$

for $t = 1 \ldots T$, where

$$V_{t,T} = \sum_{j=1}^{T} \operatorname{sgn}(X_t - X_j), \tag{4}$$

and

$$\operatorname{sgn}(x) = \begin{cases} 1 & \text{if } x > 0 \\ 0 & \text{if } x = 0 \\ -1 & \text{if } x < 0 \end{cases}$$

$U_{T,T} \equiv U_{0,T} = 0$. A large K_T^+ is indicative for a "downward shift" ($\Delta < 0$) in the level from the beginning of the series and a large K_T^- is indicative for a "upward shift" ($\Delta > 0$) in the level of the series.

Provided that the random variables $X_1 \ldots X_T$ are independent, the critical values k^+ (or k^-) of K_T^+ (or K_T^-) and critical value k of K_T for a risk (significance level) p

$$p = \text{prob}(K_T^+ \geqslant k^+ | H_0)$$

is approximated by

$$(k^+)^2 = -\left(\frac{\ln p}{6}\right)(T^2 + T^3) \tag{5}$$

for K_T^+, and

$$k^2 = -2\left(\frac{\ln p}{6}\right)(T^2 + T^3) \tag{6}$$

for K_T, where the approximation holds good, accurate to 2 decimal places for $p < 0.5$ (for details see Pettitt, 1979). The time τ is equal with t for which $U_{t,T} = K_T^+$ (or K_T^-).

Pettitt's procedure operates as a test with a significance level as specified only if the considered time series is formed from *independent* data. It is meant to be used when *one* change point is present, i.e., *one* abrupt change of the mean. To find out whether *several* change points are present the pragmatic approach is used to first find the most visible one, to split the time series into two sub time-series at the change point and to repeat the analysis for the sub time-series.

To ensure that the basic assumptions of stationarity and independence are satisfied, the considered time series is often subjected to two other null hypothesis tests. First a test of the null hypothesis of no serial correlation (for instance, Wald and Wolfowitz, 1943) and then a test of the null hypothesis of no trend (Mann, 1945; Sneyers et al., 1994; Kulkarni and Von Storch, 1995). However, such tests can never positively prove that the data are free of serial correlation and a (linear) trend, simply because no statistical test can lead to the rejection of the alternative hypothesis. A non-rejection of the null hypothesis is merely indicative of the available data not being sufficient to reject the null hypothesis. If more data would be available, it could very well happen that the same null hypothesis is rejected at a high significance level (see the discussion of this matter in Von Storch and Zwiers (1988)).

The effects of serial correlation and of the linear trend on the Pettitt-test are examined by means of Monte-Carlo experiments in the Appendix.

The performance of the test depends sensitively on these assumptions. A possibly existing serial correlation can be filtered by a "prewhitening" (Katz, 1988; eq. (8) in the Appendix). All Pettitt statistics used below are calculated from such prewhitened time series.

A linear trend has a disastrous effect on the test since it makes the test markedly liberal (i.e., the null hypothesis is too often incorrectly rejected). If only a linear trend is present the subtraction of the sample trend reduces the rejection rate of the null hypothesis to the nominal one. The problem becomes more complicated if a linear trend and one or more abrupt changes are present because every abrupt change point induces an artificial trend. A subtraction of the sample trend (which includes both real and artificial trend), however, reduces the power of the test such that it becomes sometimes useless. In this unsatisfactory situation we have chosen the Pettitt-test not to use in the conventional manner. Firstly, we relax the definition of a change point so that it covers not only abrupt changes of the mean but also changes of piecewise linear trends and other changes of the statistics in time. Secondly, we use Pettitt's approach not as a *test* but as a mere *exploratory tool.* Thus large (small) $U_{t,T}$ are taken as indications for possible "downward" ("upward") change points, or points of inhomogeneity. Such change-points are accepted as physically meaningful when consistent change points are found in physically connected time series.

3.2. Canonical correlation analysis (CCA)

The CCA is a tool to find out linear relationships between two space-time dependent variables (Barnett and Preisendorfer, 1987; Von Storch, 1995). The CCA selects a pair of spatial patterns of two space-time dependent variables such that their time coefficients are optimally correlated. Since the coefficients are normalized to unity, so that the canonical correlation patterns represent the typical strength of the signal. Thus the coefficients may be seen as time series of weights which describe the strength and the sign of the patterns for each realization in time.

Prior to the CCA, the original data are projected onto their empirical orthogonal functions (EOFs) and only a limited number of them are retained, explaining most of the total variance. This also serves as a data-filtering procedure to eliminate noise (although it can exclude potentially useful information). Bretherton et al. (1992) suggest that a considered variance of about 70% to 80% represents a good compromise.

Note that the canonical correlations are overestimated when derived from a finite sample (Von Storch, 1995).

4. Results

The changes in the winter Romanian precipitation and SLP fields are discussed in several steps. First, the individual time series of winter mean precipitation at the 14 locations are examined (Subsection 4.1). Next, EOFs for precipitation and the European pressure distribution are derived and their coefficient time series are analysed with respect to change points (Subsection 4.2). Finally, in Subsection 4.3, the CCA of precipitation and pressure is done and the coefficient time series are screened for simultaneous change points.

4.1. *Trends and change-points in the Romanian precipitation records*

The time series of winter mean precipitation at the 14 Romanian stations (Fig. 1) exhibit small serial (winter-to-winter) correlations. Maximum values of this correlation $\hat{\alpha} \leqslant 0.22$ are found for Ocna Sugatag, Tg. Jiu, Bucharest, Calarasi, Sulina, Constanta and Iasi.

The trends of the winter precipitation in Romania during the 1901–1988 interval is presented in Table 1 (second column). On average precipitation is increasing by 17 mm/100 years. At nine stations the amount of precipitation has increased and at five the amounts have decreased. According to Mann's test (1945) only the trends for Brasov and Roman stations are statistically significant at the 5% level (after prewhitening).

Pettitt's statistics $U_{t,T}$, derived from all data, indicates upward change-points in the early 1930s and downward change points at about 1969/1970 for several stations (Table 1, 3rd column). When we calculate Pettitt's statistic only for the time after the first alleged change point then almost all stations exhibit a downward shift at about 1969/1970 (Table 1, 3rd column). In Table 2 the mean values calculated before and after change-points are listed as well as the shifts of the means (Δ) associated with the change-points in the mid 1930s and in 1969/1970. The shifts are sometimes large, for instance at Tg. Jiu in 1933 or in 1969/1970 at Brasov. The shift is in the order of 10 to 50 mm in the mid 1930s and -10 to -20 mm in 1969/1970. We will see later that these changes are consistent with changes in the large-scale circulation so that they are likely not due to changes in observational routines.

Fig. 2 shows the temporal evolution of the precipitation anomalies for the stations Tr. Severin and Brasov.

Our results are consistent with similar analyses of winter precipitation at Bulgarian and other Romanian stations (Boroneant et al., 1995).

4.2. *EOFs of Romanian precipitation and Central European SLP*

The first two EOFs for the winter Romanian precipitation and SLP have been computed from

Table 1. *Year-to-year correlations, trends (mm/season/100 years), change-points (year; a ↑ represents an upward shift and a ↓ a downward shift) over the complete interval 1901–1988 and the sub-interval 1935–1988 for the winter precipitation in Romania*

Stations	Year-to-year correlation	trend	1901–1988 change-point	1935–1988 change-point
Ocna Sugatag	−0.15	22	1939 ↑	1969 ↓
Baia Mare	0.02	−33	1923 ↓	1967 ↓
Bistrita	−0.08	25	1955 ↑	1954 ↑
Brasov	0.01	−31	1969 ↓	1969 ↓
Sibiu	0.02	11	1918 ↑	1970 ↓
Timisoara	0.01	16	1949 ↑	1948 ↑
Tr. Severin	0.02	58	1933 ↑	1970 ↓
Tg. Jiu	0.17	67	1933 ↑	1969 ↓
Bucharest	0.19	56	1919 ↑	1969 ↓
Calarasi	0.22	−5	1969 ↓	1969 ↓
Sulina	0.19	−4	1970 ↓	1970 ↓
Constanta	0.18	20	1951 ↑	1969 ↓
Roman	0.07	−30	1970 ↓	1969 ↓
Iasi	0.15	17	1930 ↑	1969 ↓

For the geographical locations refer to Fig. 1. The time series for Brasov and Tg. Severin are displayed in Fig. 2.

Table 2. *Winter mean precipitation amount over the intervals 1901–1934, 1935–1969 and 1970–1988 which are separated by change-points*

Stations	1901–1934	Δ	1935–1969	Δ	1970–1988
Ocna Sugatag	117.4	+34.7	152.1	−29.2	122.9
Baia Mare	217·6	0	217.6	−25.6	192.0
Bistrita	115.4	+10.4	125.8	+0.2	126.0
Brasov	98.4	+2.8	101.2	−29.3	71.9
Sibui	80.0	+15.6	95.6	−18.0	77.6
Timisoara	120.8	+7.3	128.1	−1.0	127.1
Tr. Severin	134.9	+52.8	187.7	−23.1	164.6
Tg. Jiu	135.9	+61.2	197.1	−28.0	169.1
Bucharest	98.6	+29.2	127.8	−5.6	122.2
Calarasi	101.2	+4.7	105.9	−20.0	85.9
Sulina	70.7	+14.3	85.0	−27.2	57.8
Constanta	80.0	+14.0	94.0	−8.1	85.9
Roman	75.2	−1.1	74.1	−18.6	55.5
Iasi	83.1	+14.4	97.5	−12.6	84.9

The change in the mean (Δ) between the intervals is also given. For the geographical locations refer to Fig. 1. The time series for Brasov and Tg. Severin are displayed in Fig. 2.

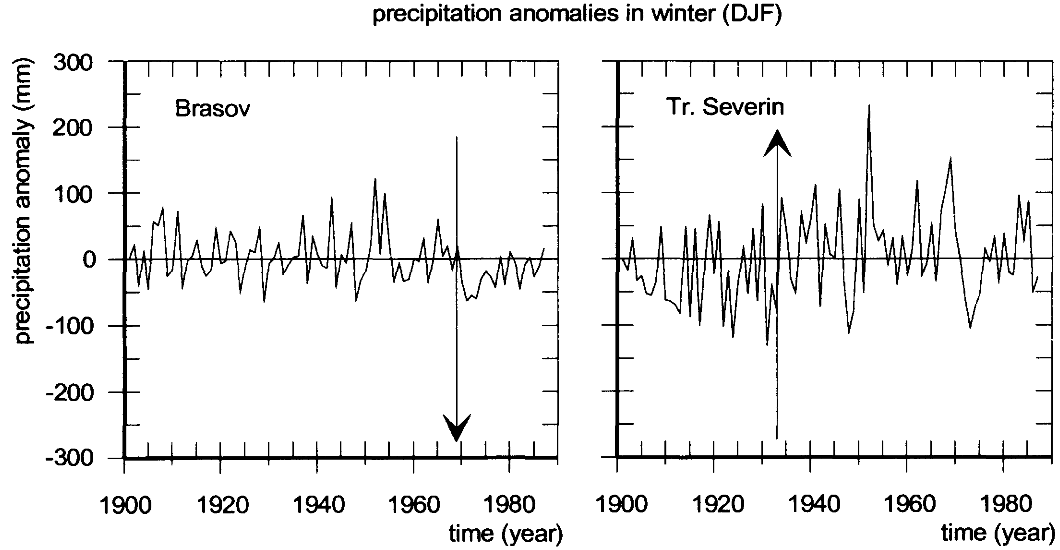

Fig. 2. Winter precipitation anomalies for the Tr. Severin and Brasov stations. The "change-points" are marked by vertical arrows. The location of the stations is given in Fig. 1.

the full data set 1901–1988. These patterns show the main spatial features of the two analysed variables and their coefficient time series describe the dominant variability in the data sets.

The first two EOFs for the Romanian precipitation are shown in Fig. 3. The first EOF explains 48% of the total variance and has the same sign over the entire area with the highest values in the southwestern part and decreasing to the northeast. This pattern suggests that in spite of the highly irregular topography of the region, there is a common physical process dominating the winter precipitation variability and this process could be linked to large-scale processes. We will be returning to this aspect in Subsection 4.3. The second EOF (20% explained variance) has a

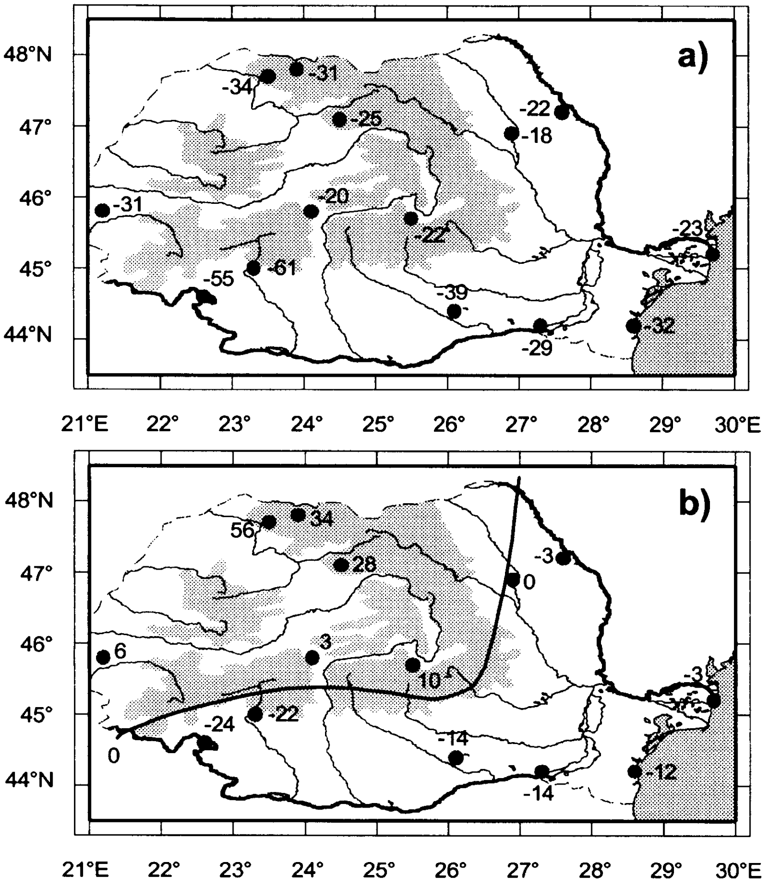

Fig. 3. The patterns of the first (a) and second (b) EOFs for the winter Romanian precipitation (1901–1988). The coefficients are normalized to one so that the patterns represent typical dimensional distributions (in mm).

dipole structure, the positive and negative values being separated by a line which follows almost exactly the Carpathian topography. This pattern suggests the influence of the Carpathian (especially in the higher southern part) on the precipitation variability. These results agree with those obtained by other studies, for instance Draghici (1988), Ion-Bordei (1988).

The first two SLP EOFs explaining 45% and 33% of the total variance are presented in Fig. 4. The first EOF shows a large anomaly area with the same sign and the second EOF shows a dipole structure with the gradient oriented from north-east to southwest.

The coefficient time series of the first two EOFs of SLP and Romanian precipitation are presented in Figs. 5, 6. The Pettitt statistic signals the presence of a downward shift in the precipitation-related first EOF coefficient and an upward shift in the first SLP EOF coefficient at about 1933. The Pettitt-statistic also offers some evidence that there might be a secondary change point at about 1969. The 2nd EOF coefficient time series of precipitation and SLP have also similar downward change-points (1923 for SLP and 1926 for precipitation).

Considering the EOFs patterns of the two parameters (Figs. 3, 4), these quasisynchronous change points seem to be dynamically consistent and indicate a possible link between them. The coefficient time series of the first SLP EOF was before 1933 on average negative so that the rain-bringing

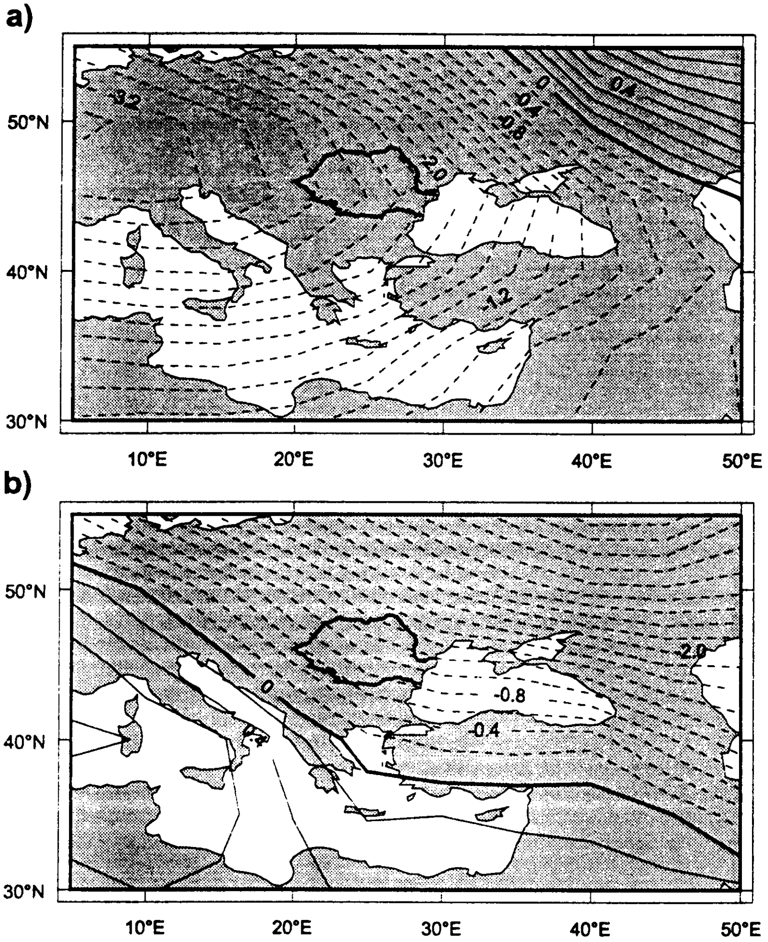

Fig. 4. The patterns of the first (a) and second (b) EOFs for the winter SLP (1901–1988). The coefficients are normalized to one so that the patterns represent typical dimensional distributions (contour interval of 0.2 hPa). Continuous lines mark positive values, and dashed lines negative values. The area of Romania is encircled by a heavy line.

southwesterly circulation over Romania was less frequent. Consistently was the first EOF-coefficient of precipitation on average positive so that less precipitation, with maximum deficits at the southwesterly flank of the Carpathians, was observed (Fig. 3a). Between 1934 and 1969 the situation was reversed, and after 1969 the southwesterly circulation became again less frequent. A similar interpretation is suggested by the patterns of the second pair of EOFs with less rainfall in the respective lee-side of the Carpathian (Fig. 3b). The intensity and/or frequency of the steering northwesterly circulation changed in the mid 1920s.

4.3. Connection between Romanian precipitation and the large-scale circulation

The CCA analysis determines the pairs of the spatial patterns of the SLP and Romanian precipitation such that their time components are optimally correlated. The first 6 EOFs for the SLP and 5 EOFs for the Romanian precipitation have been retained for the CCA.

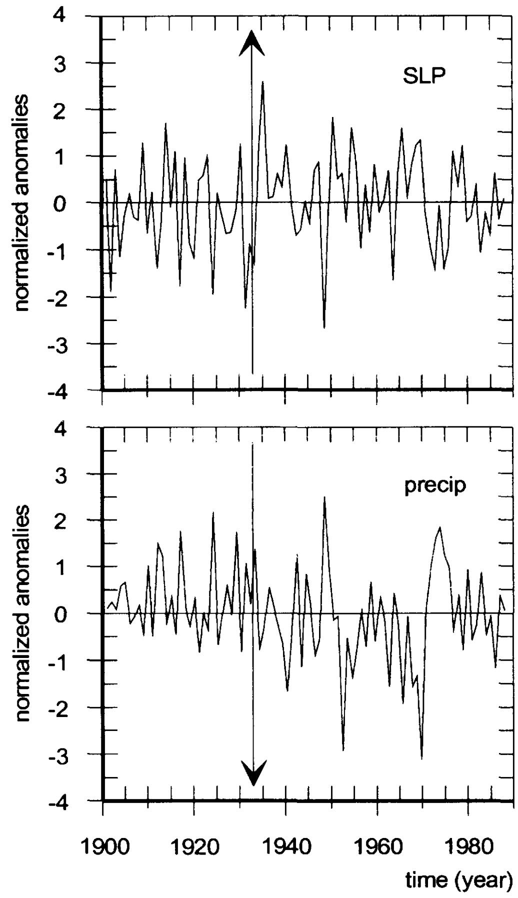

Fig. 5. Coefficient time series of the first EOF of the winter mean SLP and total winter Romanian precipitation. The "change-point" is marked.

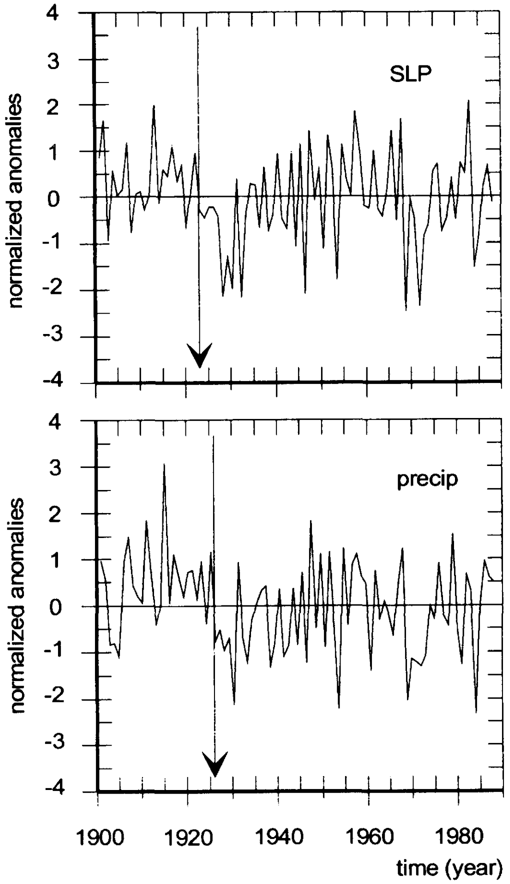

Fig. 6. Coefficient time series of the second EOF of the winter mean SLP and total winter Romanian precipitation. The "change-point" is marked.

The first CCA pair exhibits a correlation between the precipitation and SLP coefficient time series of 0.84. They explain 35% of the total seasonal mean SLP variance and 47% of the total precipitation variance. The patterns (Fig. 7) are similar to the first EOFs for both the variables (Figs. 3a, 4a) and represent a link that is very reasonable from the physical point of view: low pressure over Europe and the Mediterranean basin guides maritime air and precipitating weather systems into Romania, such that above normal precipitation is recorded. The maximum values of almost 59 mm are in the southwest and the minimum values of 17 mm in the northeast that shows the orographic perturbation effect of the Carpathian mountains.

The second CCA (0.65 correlation) explains 31% of the total SLP variance and 20% of the total precipitation variance. The patterns (Fig. 8) are similar to the second EOFs for both the variables (Figs. 3b, 4b). Again a physically plausible link is suggested by the patterns: the SLP pattern describes a northwesterly flow that affects mostly the intra-Carpathian region where the positive precipitation anomalies are emphasized, the highest (of almost 50 mm) being in the northwest.

In Fig. 9a, the time coefficients of the first CCA pair are shown. The year-to-year variations are highly coherent. Both curves have an upward

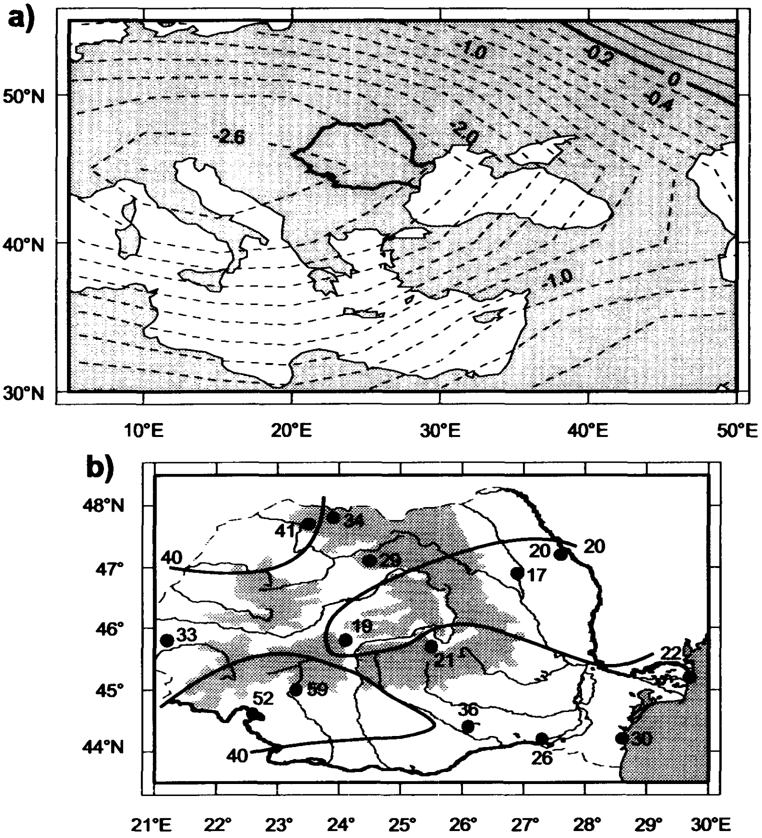

Fig. 7. The patterns of the first canonical pair of the winter mean SLP (contour 0.2 mb; the area of Romania is encircled by a heavy line) and total winter Romanian precipitation (contour 20 mm). Continuous lines mark positive values, and dashed lines negative values.

change-point at about 1933. The link is strong and now we can assert *that changes in the Romanian winter precipitation are due to changes in the large-scale circulation.* This phenomenon is very clear for the southwestern stations which are controlled by the first CCA. Less marked upward shifts which have taken place at other locations (compare with Table 1) such as Ocna Sugatag (1939), Bistrita (1954), Constanta (1951) may also be related to the same phenomenon. Only Bucharest does not fit into the picture, and we suggest that this station may be affected by the local process of urbanization.

Fig. 9b shows the time coefficients of the second CCA pair. The year-to-year variations are fairly coherent. Both curves share a simultaneous downward change-point in the mid 1920s. The CCA-patterns of the SLP and Romanian precipitation (Fig. 8) suggest the following physical link: an intensification of the northwesterly circulation over Romania leads to more precipitation in the intra-Carpathian region, with the stronger effect in the northwest. The downward shift for the Baia Mare station seems to be related to this mechanism.

5. Conclusions

Some conclusions may be drawn from this study, with respect to the methodology and to the physical mechanism of the Romanian precipitation changes during the winter.

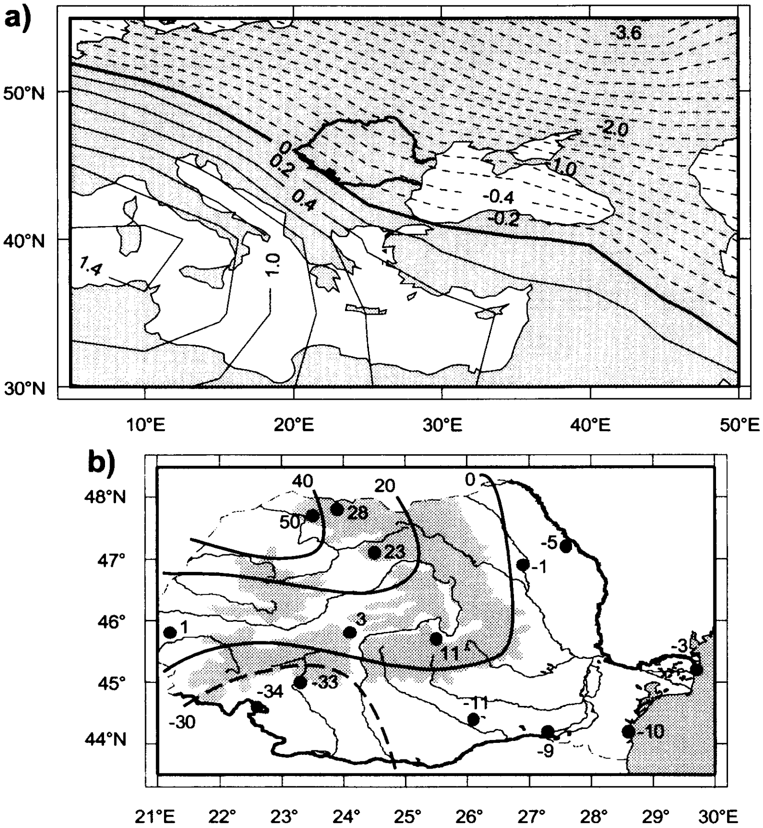

Fig. 8. The patterns of the second canonical pair of the winter mean SLP (contour 0.2 mb; the area of Romania is encircled by a heavy line) and total winter Romanian precipitation (contour 20 mm). Continuous lines mark positive values, and dashed lines negative values.

5.1. Methodical aspects

The Pettitt procedure is a good *exploratory* tool to detect change-points, if we relax the definition of change points from the strict case of abrupt changes in the mean to all kind of changes in statistics (i.e., a change in regime). When the Pettitt's procedure is applied to individual time series any determined change point may be due to a real change-point in the environment or to changes in observational, reporting or analysing techniques. To exclude such artificial signals we recommend the application of Pettitt's statistic to pairs of physically linked variables.

Generally, Pettitt's procedure is used as a *confirmatory* tool, namely as a method for testing the null hypothesis "no change point". However, this procedure qualifies as a test only if the data fulfill to non-trivial assumptions, namely lack of serial correlation and lack of a (linear) trend. In climatological applications these assumptions are often violated. These violations have severe implications (see Appendix).

The serial correlation effects the result of the Pettitt test such that the percentage of erroneous rejections of the null hypothesis is larger than permitted according to the significance level. We found that "prewhitening" of the time series prior to the test is sometimes a good cure.

The presence of linear trends causes the test also to reject the null hypothesis too often. However, when the estimated linear trend is subtracted from the time series, the power of the test to detect real change points is reduced in such a manner than the test becomes useless.

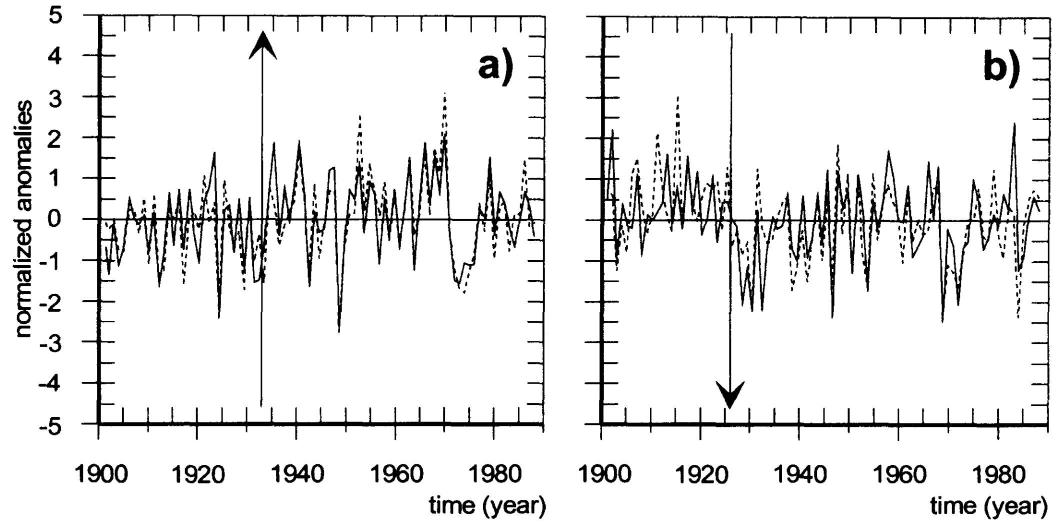

Fig. 9. Normalized time concepts of the first (a) and second (b) CCA patterns of SLP anomalies (continuous line) and Romanian precipitation anomalies (dashed line).

We recommend not to use Pettitt's procedure as a confirmatory tool when dealing with geo-environmental data. Such change points are accepted as physically meaningful when consistent change points are found in physically connected time series.

Other authors (Sen and Srivastava, 1975; Solow, 1987) have remarked earlier that the differentiation between abrupt changes and other temporal inhomogeneities is difficult. This can be accomplished only by assuming specific statistical models for the data under consideration; the justification of such models for climate data is difficult. Therefore we recommend to use the Pettitt-approach as it is, and to confirm and assess subjectively the found result.

5.2. *Physical aspects of the changes of Romanian precipitation*

The link between Romanian winter precipitation variability and the European SLP variability is strong. This link appears to be primarily related to the first SLP EOF or, in other words, to the strength of the southwesterly circulation from Mediterranean. A significant intensification of this circulation type since 1933 has been found. Because of the topographic structures, related to the Carpathians, this change is impacting mainly the southwestern stations which have an upward change-point about the same period. Since 1969, the southwesterly flow has weakened leading to the decreasing of the precipitation at some central and eastern stations.

Another mechanism, which seems to be responsible for a change point in the mid 1920s, is the strength of the northwesterly circulation which affects precipitation in the Intra-Carpathian.

We can not assert whether the dynamics of these changes in the large-scale circulation found in this study are natural fluctuations of the climate system or are determined by the external forcings.

6. Appendix

Most standard statistical techniques are derived with the explicit need for statistically independent and identically distributed (i.i.d.) data. However, almost all climatic data are somehow correlated in time so that the assumption of statistical independence is often violated. Also, low frequency variability on time scales of tens and hundredths of years introduces into time series extending over tens of years (linear) trends, i.e., instationarities in the mean.

In this Appendix we discusss the sensitivity of the Pettitt test against the effects of a serial correlation and of a linear trend.

6.1. *Serial correlation*

Following the approach pursued by Kulkarni and Von Storch (1995) we demonstrate the effect of serial correlation on the performance of Pettitt's test by means of a series of Monte-Carlo experiments with synthetic data X_t generated by an AR (1)-process

$$X_t = \alpha X_{t-1} + N_t, \tag{7}$$

where X_t is an AR (1)-process, α is the lag-1 autocorrelation of X_t, and N_t is a Gaussian "white noise" which is neither auto-correlated nor correlated with X_{t-k} for $k \geqslant 1$. 1000 independent identically distributed time series (i.i.d.) of different lengths n (100, 500) were generated and the Pettitt test was performed for $\alpha = 0.0 \ldots 0.95$. Since the time series have no change-points, we expect a reject rate of 5% if we adopt a risk (significance level) of 5%, i.e., 50 out of 1000 tests should return the result "reject null hypothesis". The actual rejection rate is much higher (Fig. 10). For autocorrelations $\alpha \leqslant 0.10$ the actual rejection is about the nominal rate of 5%, but for $\alpha > 0.20$ the rejection rate increases rapidly. After "prewhitening" the process (7) (Katz, 1988; Kulkarni and Von Storch, 1995) by

$$Y_t = X_t - \hat{\alpha} X_{t-1}, \tag{8}$$

where $\hat{\alpha}$ is the estimated autocorrelation at lag-1, the rejection rate becomes close to the nominal one (Fig. 10).

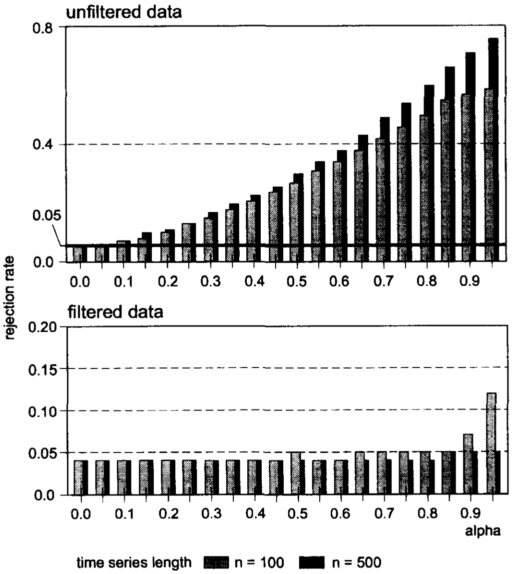

Fig. 10. Rejection rates of the Pettitt test of the null hypothesis "no change" when applied to 1000 time series of length $n = 100$ and $n = 500$ generated by an AR(1)-process (7) with prescribed α. The adopted nominal risk of the test is 5%. Top: results for unprocessed serially correlated data. Bottom: results after prewhitening the data (8).

6.2. *Linear trends*

To show the effect of the linear trend on the Pettitt's result we created synthetical time series X_t by means of the random process

$$X_t = N_t + \beta t, \tag{9}$$

with white noise N_t with variance 1, and a trend coefficient β.

For various trend coefficients β one thousand time series of length 100 were generated and both one-sided Pettitt's tests were done with a risk of 5%. In the following table the rates of (erroneous) identifications of an upward, or downward, change point are listed. A proper performance of the test would require a rate close to 0.05.

β	Rejection rate upward	Rejection rate downward
0.000	0.044	0.045
0.001	0.073	0.025
0.002	0.128	0.017
0.003	0.206	0.005
0.004	0.288	0.002
0.005	0.381	0.002

Obviously the linear trend does cause severe biases. Since the prescribed linear trend is an upward trend, upward change points are signalled much more often than permitted by the significance level. Also, much less downward change points are found. The bias of the test is already serious for $\beta = 0.002$. Note that a trend coefficient of $\beta = 0.002$ creates a moderate increase over 100 time steps of 0.2, which has to be compared with a variance of 1 of the white noise.

We have also counted how often a certain time in the time interval between 1 and 100 is picked by the Pettitt test as representing a change point.

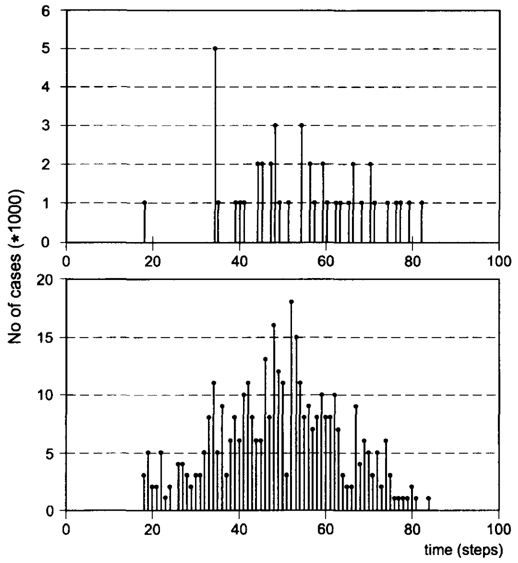

Fig. 11. The Pettitt procedure does not only deal with the null hypothesis of a change point but also identifies the time at which an alleged change takes place. The figure shows the frequency with which any of the 100 time steps in the Monte Carlo experiment is (erroneously) picked as a change point. Top: time series without a trend; bottom: time series with a trend ($\beta = 0.005$).

For $\beta = 0$ and $\beta = 0.005$ frequency distributions are plotted in Fig. 11. In both cases change points are incorrectly identified in the middle of the time interval. They are not, as one might wish, uniformly distributed over time.

We tried to solve the problem in a manner similar in concept to the prewhitening used above. After subtraction of the estimated trend from the data the rejection rate of the null hypothesis is closed to the nominal one. However, the subtraction of the fitted trend is punished by a reduction of the power of the test.

When more complicated instationarities prevail, such as a linear trend overlaid by one or two abrupt change points, then the situation becomes more difficult and can be dealt with only if additional information is available to build adequate a-priori statistical models.

In conclusion, we recommend the "prewhitening" of time series when the serial correlation is present and afterwards the use of the Pettitt test. When, however, a linear trend is present, Pettitt's test can not be used any longer as a confirmatory tool. We recommend using the Pettitt test as a mere exploratory tool and calculating Pettitt's statistic and dealing with possible change points as unproven hypotheses, which plausibility should be supported by physical arguments.

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