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NONSTATIONARITIES IN TECHNICAL PRECIPITATION SERIES IN CZECHIA

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Under the proper consideration of climate change one cannot rely on the stationarity of the series fed to a hydrological model. Modelling should rather be based on the assumption of nonstationarity with appropriate discrimination between deterministic and stochastic trends in the series. Although some works regarding the identification of trends in precipitation across the entire territory of Czechia were undertaken, they primarily accounted for the deterministic and linear relationships only. What is more, the results in many former studies inevitably had to be biased due to missing values and inhomogeneities introduced during measurements. In the present paper, trends in 268 complete and homogeneous precipitation series from Czechia spanning from 1961 to 2012 are detected by means of several statistical techniques. Subsequently, places with different kinds of nonstationarities are determined.

KĽÚČOVÉ SLOVÁ: trends, Hurst exponent, unit root testing, seasonality, rainfall

NESTACIONARITY V TECHNICKÝCH ŘADÁCH ÚHRNŮ SRÁŽEK V ČESKU. Za předpokladu klimatické změny není možné nadále počítat se stacionaritou časových řad vstupujících do hydrologického modelu. Modelování by bělo být spíše založeno na předpokladu nestacionarity s rozlišením deterministických a stochastických trendů. Přestože již některé analýzy trendů ve srážkových úhrnech na území Česka v minulosti proběhly, byly zohledňovány především pouze deterministické trendy a lineární závislosti. Navíc výsledky mnohých dřívějších studií musely být zákonitě ovlivněny chybějícími hodnotami a nehomogenitami zavedenými do řad v průběhu měření. V tomto příspěvku bylo analýze trendu podrobeno 268 srážkových řad, které byly nedávno homogenizovány a doplněny a které reprezentují období 1961–2012. Na původní řady a další z nich odvozené byly aplikovány nejrůznější statistické metody s cílem zjistit, kde přesně v Česku dochází k významným změnám (nárůstům či poklesům) ve srážkových úhrnech.

KEY WORDS: trendy, Hurstův exponent, testování na jednotkový kořen, sezonalita, srážky

Introduction

In recent years, climate change (CC) impacts on different ecosystems as well as on human society have been more and more emphasized and many hydrological studies on them have emerged. However, prior to the investigation of the effects of CC themselves, there is a need to find out if at all climatic elements such as temperature or precipitation exhibit some important nonstationarities. Besides, adressing the question where exactly the changes occur is equally imperative since some regions may not show any, while others do.

Because floods and droughts belong to the most discussed CC consequences in Czechia, the studies aimed at the proper examination of several features of precipitation series are necessary here. Although the measurements of precipitation totals have a very long tradition, not many uninterrupted long precipitation series can be found across the country. This led to the fact that a lot of past studies focused only on few selected (mostly secular) precipitation series (e.g., Brázdil et al., 2012), whereas nationwide assessments were rather seldom. Currenty, on the other hand, a set of so-called technical series (including precipitation) is accessible through the Czech Hydrometeorological Institute (CHMI), which starts in 1961 and proceed to the present. This means that over 250 series of precipitation totals that are homogenized and complete (i.e., missing values were filled in) are available in daily time step for future works regarding the territory of Czechia (Štěpánek et al., 2011, 2013). This poses also a great potential for the next hydrological studies.

It is true that some studies conducted so far have already examined the changes in above mentioned technical precipitation series (see Střeštík et al., 2014). Nevertheless, they focused on quantifying the magnitudes of changes such as deterministic trends rather than on proper testing for their significance. Moreover, only techniques that require serial independence (e.g., linear regression based on ordinary least squares) were empoyed, which certainly ended with misleading results because hydrometeorological series commonly exhibit some sort of short-tem persistence (STP) or even long-term persistence (LTP) that cause the significant trends are detected too often in the case of positive autocorrelation. The opposite holds when a series is autocorrelated negatively, which is not so typical in hydrology or climatology.

The goal of the present study is therefore to take the same series as in Střeštík et al. (2014) that spatially cover the whole area of Czechia and try to find out if some nonstationarities are present in them. Note that only the nonstationarity in the mean is considered here which, to a large extent, imply the existence of eiher stochastic or deterministic trends. First, a distinction between these two kinds of trends is made using the unit root tests and the Hurst exponent estimation. Then, a statistical test for trends with null hypotheses set up adequately are performed to better understand what places in Czechia experience changes in precipitation. Abrupt changes are not subject to this analysis because it was assumed that many of them were eliminated during the construction of the technical series. The assessment of nonstationarities in the variance was postponed to the future work since this issue has not received much attention in hydrology yet and not much literature on it can be found (Modarres and Ouarda, 2013).

Material and methods

The same dataset of technical (i.e., homogenized and filled in) precipitation series that was used in Střeštík et al. (2014) was accessed from the CLIDATA database of the CHMI. The dataset comprised daily precipitation totals (in mm) from 268 stations from the whole territory of Czechia (see Figs. 1 and 2 for the spatial distribution of stations). The temporal span was from 1 January 1961 to 31 December 2012, so 52 years of uninterrupted daily data were available. In other words, there were exactly 18 993 entries for each station.

The daily series were then aggregated so that new series representing each month and the year were formed. This mainly enabled the possibility of comparing different changes wihin the year. Then, besides the basic series encompassing precipitation totals (r), other series were derived in the sense similar to the study by Brunetti et al. (2001) – annual number of days with a total equal or greater than 0.1 mm (wet days, WD) and annual precipitation intensity (I), meaning the ratio of annual r (i.e., R) to WD. Because Střeštík et al. (2014) mentioned also some changes in seasonality, the annual series of the three following indicators (with subscript k for each year spanning from 1 to 52) were additionally subjected to the analysis (see Feng et al., 2013):

- seasonality centroid (timing) C_k
- seasonality spread (duration) Z_k
- seasonality index SI_k

More specifically, the centroids were computed as:

$$C_k = R_k^{-1} \sum_{m=1}^{12} m r_{k,m} \tag{1}$$

where m was the second subscript denoting months from January (m = 1) to December (m = 12).

Having centroids, the spreads could be computed as:

$$Z_k = \sqrt{R_k^{-1} \sum_{m=1}^{12} (m - C_k)^2 r_{k,m}} \,. \tag{2}$$

Finally, the seasonality indices were acquired as follows:

$$SI_k = R_k \cdot \sum_{m=1}^{12} p_{k,m} \log_2 (12 p_{k,m}) \tag{3}$$

where $p_{m,k}$ was the probability connected with the precipitation total in the month m and the year k.

Note that unlike in the work of Feng et al. (2013), the normalization in Eq. (3) was not carried out because the goal here was not the quantification of changes in physical units (see Table 1 for them). Note also that sometimes there was no precipitation in a given month. For these cases, the following limit was used (similarly as in Sohoulande Djebou et al., 2014) in the entropy part of Eq. (3):

$$\lim_{p_{k,m} \to 0^+} p_{k,m} \log_2 (12 p_{k,m}) = 0 \tag{4}$$

Deseasonalization and determination of different kinds of nonstationarities in daily precipitation series

Since the results of trend analyses strongly depend on assumptions and namely on the null hypothesis of a trend test, where the independece among data belong to the most important requirements, special care was made in distinguishing between STP and LTP processes generating the basic (daily) time series. The tests accounting for serial independence, STP or LTP should be applied accordingly.

However, before this discrimination itself, there is a need to have the time series deseasonalized. Despite, in hydrology and climatology, often simple deseasonalization techniques are used (e.g., Salas, 1993; Grimaldi,

2004; Kantelhardt et al., 2006; Fatichi et al., 2009), here, a novel approach based on wavelets (see Szolgayová et al., 2014) was performed. In this approach, first, the multiresolution decomposition is made and, subsequently, only the component corresponding to seasonality is transformed back (level 8 in this case). This outcome is then subtracted from the original series to get the deseazonalized series. When performing the multiresolution decomposition, the maximum overlap discrete wavelet transform (MODWT) and the least asymmetric wavelet of order eight (s8) were utilized. More details on the multiresolution wavelet analysis for time series purposes can be found, for instance, in Percival and Walden (2000).

The discrimination of stochastic processes governing the precipitation series started with the so-called unit root tests, as recommended by Barbosa et al. (2008) or Fatichi et al. (2009). Namely, the Phillips–Perron (PP) test and the Kwiatkowski–Phillips–Schmidt–Shin (KPSS) test were employed. Their advantage lies in the fact that their null hypotheses are set up complementarily, which, among other aspects, allows for the recognition of the difference between deterministic and stochastic trends.

Let $D_t = \{d_1, d_2, \ldots, d_n\}$ be a deseasonalized precipitation series. The following model forms the basis of the PP test (Phillips and Perron, 1988):

$$D_t = \alpha + \beta \cdot t + \pi \cdot D_{t-1} + \varepsilon_t \tag{5}$$

where α and β are the parameters of linear regression, and ε_t is a stationary process, which may be searially correlated and heteroscedastic.

If the null hypothesis, H_0: $\pi = 1$, is rejected in favour of the alternative hypothesis, H_1: $\pi < 1$, the underlying process is not a random walk but the lag-one autoregressive (AR(1)) process about the deterministic linear trend. The key model of the KPSS test (Kwiatkowski et al., 1992) is:

$$D_t = \beta \cdot t + w_t + \nu_t \tag{6}$$

where β is the parameter of a linear trend, $w_t = w_{t-1} + \varepsilon_t, \varepsilon_t \sim N(0, \sigma_\varepsilon^2)$ is a random walk and ν_t is a stationary process.

There are two kinds of null hypotheses connected with the KPSS test. Here, the general case of trend stationarity (i.e., $H_0 : \sigma_\varepsilon^2 = 0, \beta \neq 0$ vs. $H_1 : \sigma_\varepsilon^2 > 0$) was tested. If the KPSS test is applied together with the PP test, up to three categories of stochastic processes may be identified:

- stationary short memory processes about a deterministic trend if the null hypothesis of the PP test is rejected and the null hypothesis of the KPSS test is accepted
- unit root processes along with random walks if the null hypothesis of the PP test is accepted and the null hypothesis of the KPSS test is rejected
- other nonstationary process (including difference stationatry long memory processes with their typical stochastic trends) if both null hypotheses are rejected

The fourth case, when both null hypotheses are accepted, should be considered as a failure of the tests implying that there is insufficient information in the data.

The unit root tests were also accompanied by the estimation of the Hurst exponent H whose values close to 0.5 indicate independent processes (see, e.g., Khaliq et al., 2008; Khaliq and Sushama, 2012). There are a lot of estimators in different fileds of geophysics (e.g., Fatichi et al., 2009), but, particularly in hydrology, several most precise techniques, which can work simultaneously with more than one parameter, were developed (Tyralis and Koutsoyiannis, 2011; Navarro et al., 2013). Here, the method of maximum likelihood was chosen (Tyralis and Koutsoyiannis, 2011).

Suppose that the deseasonalized daily observations of precipitation are organized in a vector $\mathbf{d}_n = D_t$. Then, the vector of estimated parameters of a long memory process may be written as $\boldsymbol{\theta} = (\mu, \sigma, H)$ with μ standing for the mean, σ standing for the standard deviation and H standing just for the Hurst exponent. The general form of the likelihood is (McLeod and Hipel, 1978):

$$p(\boldsymbol{\theta} \mid \mathbf{d}_n) = (2\pi)^{-n/2} [\det(\sigma^2 \mathbf{R})]^{-1/2} \exp[-(2\sigma^2)^{-1} (\mathbf{d}_n - \mu \mathbf{e})^{\mathrm{T}} \mathbf{R}^{-1} (\mathbf{d}_n - \mu \mathbf{e})] \tag{7}$$

where $\mathbf{e}$ is a column vector containing n unities, $\mathbf{R}$ is the n by n autocorrelation matrix and det(·) denotes the determinant of the matrix.

The estimate of H can be acquired by maximizing the single-variable function $g(H)$ in Eq. (8):

$$g(H) = -\frac{n}{2} \ln\left[\left(\mathbf{d}_n - \frac{\mathbf{d}_n^{\mathrm{T}} \mathbf{R}^{-1} \mathbf{e}}{\mathbf{e}^{\mathrm{T}} \mathbf{R}^{-1} \mathbf{e}} \mathbf{e}\right)^{\mathrm{T}} \mathbf{R}^{-1} \left(\mathbf{d}_n - \frac{\mathbf{d}_n^{\mathrm{T}} \mathbf{R}^{-1} \mathbf{e}}{\mathbf{e}^{\mathrm{T}} \mathbf{R}^{-1} \mathbf{e}} \mathbf{e}\right)\right] - \frac{1}{2} \ln[\det(\mathbf{R})] \tag{8}$$

.

Although some methods for it exist (see, e.g., Couillard and Davison, 2005, for rescaled range statistic), the significance of the distance of the estimates from the value of 0.5 was not specifically tested in this particular case. The same applied to the uncertainty in the estimation. Mainly, extremely time consuming cumputation process in conjuntion with such a huge dataset made it impossible.

Test for trend with suitable modification and looking for the relationships between trends and selected factors

As regards R series as well as all the series composed of

the derived indicators mentioned above, a lot of descriptive statistics were calculated. To name some, the values of minima, maxima, means, coefficients of variation, coefficients of skewness, coefficients of kurtosis, lag-one and lag-two autocorrelation coefficients filled a huge table that cannot be presented here. As these statistics are somewhat classical and well-known in hydrology, the reader is kindly referred, for instance, to Salas (1993).

Here, the author would like to describe rather the modification of the trend test in detail. Since the combination of unit root tests showed there are probably only the deterministic trends in daily series, only a modification of the Mann–Kendall (MK) test accounting for STP or serial independence was employed for all the derived annual series abbreviated as $X_k = \{x_1, x_2, ..., x_K\}$ with a subscript k again indicating years. Particularly, the Bayley–Hammersley–Matalas–Langbein–Lettenmaier equivalent sample size (BHMLLESS) approach with detrending prior to the estimation of lag-one autocorrelation coefficient was used, which was suggested by Yue and Wang (2004; hence the YW–MK test hereinafter). This procedure starts with the computation of so-called Sen's slope, which is an alternative estimator of a linear trend:

$$b = \mathrm{med}\left(\frac{X_j - X_i}{j - i}\right), \quad \forall i < j \quad . \tag{9}$$

This slope is used when detrending the original series, which results in the detrended series X_k^D:

$$X_k^D = X_k - bk \tag{10}$$

Having a detrended series, the next step is the estimation of the lag-one autocorrelation coefficient r_1 from it (for the well-known formula see Salas, 1993). Afterwards, two cases are distinguished.

If r_1 turns out to be close to zero (according to Anderson, 1942, for example), the original MK test is applied, whose statistic MK-S is calculated as (Mann, 1945; Kendall, 1970):

$$\text{MK-S} = \sum_{\forall i<j} \mathrm{sgn}(X_j - X_i) \tag{11}$$

where sgn(·) stands for the sign function.

The discrete distribution of this statistic can be approximated by the continuous standard Gaussian distribution if the number of years K is greater than 10 (Kendall, 1970). This condition was fulfilled here, so the standardized statistic MK-Z could be calculated:

$$\text{MK-Z} = \begin{cases} \dfrac{\text{MK-S} - 1}{\sqrt{\mathrm{var}(\text{MK-S})}} & \text{if} \quad \text{MK-S} > 0 \\ 0 & \text{if} \quad \text{MK-S} = 0 \\ \dfrac{\text{MK-S} + 1}{\sqrt{\mathrm{var}(\text{MK-S})}} & \text{if} \quad \text{MK-S} < 0 \end{cases} \tag{12}$$

where

$$\mathrm{var}(\text{MK-S}) = \frac{K(K-1)(2K+5) - \sum_{s=1}^{K} t_s s(s-1)(2s+5)}{18} \tag{13}$$

where t_s stands for the number of ties of extent s.

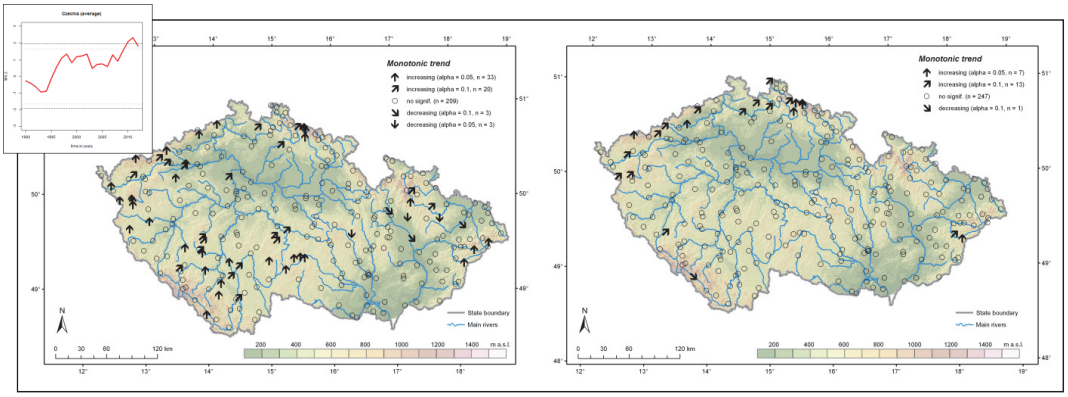


Fig. 1. Spatial distribution of trends in annual series of precipitation totals (left) and seasonality index (right) in Czechia during the period 1961–2012. The insertion on the top left depicts the evolution of the standardized Mann–Kendall statistics MK-Z for averaged annual precipitation series over the whole of Czechia using 30-year moving window (inner dashed lines – α = 0.1, outer dashed lines – α = 0.05).

Obr. 1. Prostorová distribuce trendů v ročních řadách srážkových úhrnů (vlevo) a indexu sezonality (vpravo) v Česku za období 1961–2012. Graf vlevo nahoře znázorňuje vývoj standardizované statistiky Mannova–Kendallova testu pro řadu ročních srážkových úhrnů vzniklé zprůměrováním všech 268 staničních řad. Při rekonstrukci vývoje byla využita 30letá období. Vnitřní, resp. vnější čárkované čáry vymezují 90%, resp. 95% interval spolehlivosti.

The second situation, when r_1 is significantly different from zero, requires the multiplication of Eq. (13) by a factor (Matalas and Langbein, 1962):

$$1 + 2 \cdot \frac{r_1^{K+1} - K \cdot r_1^2 + (K-1) \cdot r_1}{K \cdot (r_1 - 1)^2} \,. \qquad (14)$$

In both cases, finally, the value of the statistic MK-Z can be compared with the critical values of standard normal variates related to the prescribed significance level α.

Besides the trend detection, also the relationships between MK-Z statistics and variables such as elevation, longitude, latitude and long-term annual total at given precipitation station were experimentally determined. This was done by means of the Kendall rank correlation coefficient (see Kendall, 1970).

Results and discussion

There were a lot of outputs produced during the analysis including tables, maps and graphs. However, due to space limits, it is not possible to present them all. Among the main findings belongs the fact that the precipitation series in Czechia hardly exhibit stochastic trends. Definitely, it is worth noting there are probably no long memory processes generating the series. This is also well corroborated by the values of H that range from 0.556 to 0.705 with the median of 0.605, which is really close to 0.5 and, to a large extent, confirms the results of Kantelhardt et al. (2006). Such values may also be triggered by short memory processes (Maraun et al., 2004). As regards the estimates of the lag-two autocorrelation coefficients of the derived monthly or annual series, they were insignificant amost everywhere. A few exclusions were found concerning the WD and I series, and some seasonality indicators (namely C). However, despite this fact, only the YW–MK test that accounts maximally for STP (AR(1) in particular) was employed when detecting trends. It was thought that the absolute values of these coefficients decrease after the removal of trend (Eq. (10); see Yue et al., 2002). Nevertheless, no special check was performed here.

For the period 1961–2012, the YW–MK test suggested that not many changes reported by Střeštík et al. (2014) are significant. The stations with no trend clearly dominate throughout the whole territory of Czechia. On the other hand, some places really experience changes whose spatial distribution differ with time and studied indicator. As can be seen in Fig. 1, some increases in R series are evident in Bohemia and also in northern Moravia. In the centre of Moravia, along the middle reaches of the Morava River, one can observe even decreases. It seems that the rises in precipitation are closely related to elevation, which may be documented by Table 1 as well. Mainly mountain areas or mid-altitudes experience these rises. SI series exhibit increases connected rather with long-term totals, latitute and longitude than with elevation (Table 1). The domination of the most humid month becomes more important, which may be demonstrated by Fig. 2. While in June the precipitation totals drop, in July they increase. The July increases are evident in Bohemia, which confirms the statements of Střeštík et al. (2014). WD and I series show no clear spatial pattern. There is somewhat a mixture of their changes over Czechia. However, in the mountain areas the rises in I are caused by both the less WD and the higher R towards present days. C and Z series exhibit almost no changes.

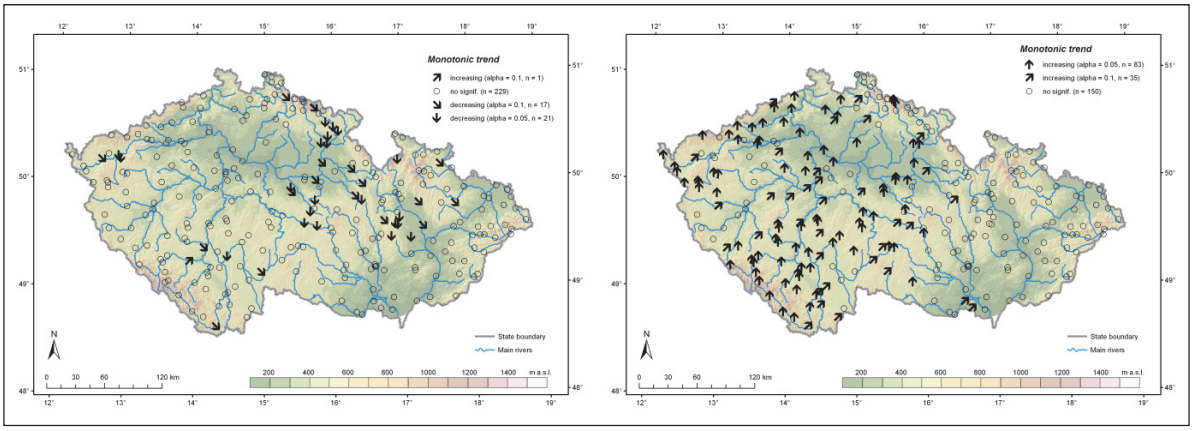


Fig. 2. Spatial distribution of trends in series of June (left) and July (right) precipitation totals in Czechia during the period 1961–2012.
Obr. 2. Prostorová distribuce trendů v řadách červnových (vlevo) a červencových (vpravo) srážkových úhrnů v Česku za období 1961–2012.

Table 1. Kendall's rank correlations between selected physical-geographical (PG) factors and standardized Mann–Kendall statistics MK-Z valid for the period 1961–2012
Tabulka 1. Závislosti standardizovaných statistik Mannova–Kendallova testu MK-Z platných pro období 1961–2012 na vybraných fyzickogeografických (PG) charakteristikách měřené Kendallovým koeficientem pořadové korelace

PG factor	*R* [mm]	WD [days]	*I* [mm/day]	*C* [month index]	*Z* [months]	SI [mm]
Elevation [m]	**0.254**	0.012	**0.123**	0.001	**0.183**	0.048
Longitude [m]	**-0.348**	-0.026	**-0.162**	**-0.461**	**-0.169**	**0.103**
Latitude [m]	0.043	-0.055	**0.104**	**0.161**	**0.136**	**0.216**
Long-term total [mm]	0.071	**-0.103**	**0.136**	**-0.256**	**0.117**	**0.263**

R – annual precipitation total, WD – number of wet days, *I* – precipitation intensity, *C* – seasonality centroid, *Z* – seasonality spread, SI – seasonality index
Note: Longitude and latitude were given in metres because projected coordinates were available.
R – roční úhrn srážek, WD – počet srážkových dnů, *I* – intenzita srážek, *C* – střed sezonality, *Z* – rozptyl sezonality, SI – index sezonality
Pozn.: Zeměpisná šířka a délka byly k dispozici v projektovaných souřadnicích, takže jsou v metrech.

Table 2. Continuation of Table 1, but for the precipitation totals in each month
Tabulka 2. Pokračování Tabulky 1, tentokrát pro srážkové úhrny v jednotlivých měsících

PG factor\Month	1	2	3	4	5	6	7	8	9	10	11	12
Elevation [m]	**0.111**	**0.213**	**0.230**	-0.009	**0.106**	**-0.135**	**0.135**	**0.141**	-0.035	**0.294**	**0.190**	-0.011
Longitude [m]	**-0.140**	0.003	**0.085**	**0.284**	**-0.272**	-0.072	**-0.539**	**-0.319**	**0.264**	**-0.412**	**-0.388**	**-0.414**
Latitude [m]	**0.083**	**0.140**	**-0.232**	**-0.222**	**-0.113**	-0.057	0.074	**0.125**	**-0.122**	**-0.087**	**0.278**	0.040
Long-term total [mm]	0.076	**0.401**	**0.210**	-0.051	**-0.118**	**-0.171**	**-0.137**	-0.022	0.013	-0.025	0.001	**-0.269**

Note: Boldface values are significant at the level $\alpha = 0.05$.
Pozn.: Hodnoty vysázené tučně jsou významné na hladině $\alpha = 0.05$.

Interesting insights were gained by looking at the series of every single month. Apart from the facts mentioned earlier, mainly winter and spring precipitation totals show important changes, if some. In winter (particularly in January) one can observe increases, whereas in spring (particularly in April) the decreases are apparent. Roughly speaking, among selected physical-geographical characteristics, the longitude seems to be the most essential, indicating Bohemia becomes wetter and Moravia becomes drier (Table 2).

After the sequential determination of the MK-Z statistics using 30-year window in spirit similar to studies such as Doleželová (2012, 2014) or Hannaford and Buys (2012), diverse results were acquired for different stations and indicators. Nevertheless, several generalized findings can be mentioned: (1) on average, *R* series had a tendency to decrease for the periods ending in the mid-1990s and then they started to increase, exceeding the limits of 95% confidence interval at the end of 2000s; (2) the series of seasonality indicators mostly remain within the confidence limits; and (3) SI series have their typical courses converse with respect to the courses of *Z* series, meaning that some changes in the shape of long-term annual pluviograph records may be expected, with dominant waves getting shorter and higher.

Conclusion

The present contribution dealt with several important features of technical precipitation series comming from 268 stations covering the entire territory of Czechia. Assuming climate change impacts, in particular, the detection of significant nonstationarities in the mean was carried out. In fact, the changes were expressed in terms of identified trends whose locations were determined. It was found that, with the highest probability, there are no stochastic trends in the precipitation series. However, there may be some deterministic ones resulting in increases in summer (July) and winter (January), while decreses in spring (April). Spatial patterns suggest that, if some, increases in precipitation totals are observable rather in Bohemia. In Moravia, the totals are generally stable or decrease. From the perspective of seasonality indicators, some changes in the long-term annual course may be expected in the future. First, some shifts of maxima from the month of June to the month of July have already been shown in the data. Second, in the next decades the months with maxima will probably become even more humid, while others will get drier at a few stations. For hydrologists, there is a very important finding revealed by the data, namely that the intensity of precipitation

significantly rises at places where many headwaters are situated. Thus, more frequent flash flooding may be expected in Czechia (especially in Bohemia).

Since there were a lot of outcomes produced during the analysis, naturally not all of them could be interpreted and presented appropriately. Instead, the study is believed to be as a good starting point for future works. For instance, the trends discovered here might be compared with the trends in discharge series. Thus, the technical precipitation series would undergo an alternative check of their validity. Further, the statement of Střeštík et al. (2014) on no significant period should be reconsidered in detail. The utilization of wavelet analysis and scalograms would be beneficial for this purpose (e.g., Torrence and Compo, 1998; Sleziak et al., 2015). Similar analysis may also be conducted concerning extremes.

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NESTACIONARITY V TECHNICKÝCH ŘADÁCH ÚHRNŮ SRÁŽEK V ČESKU

Příspěvek se zabývá hned několika statistickými vlastnostmi technických řad srážkových úhrnů z 268 stanic z celého území Česka, které z časového hlediska pokrývají období 1961–2012 v denním kroku. Řady jsou úplné (bez chybějících hodnot) a homogenní (tj. neobsahují žádně vychýlení, které by neodpovídalo samotné přírodě). V popředí zájmu příspěvku je analýza trendů coby reprezentantů jednoho z mnohých typů nestacionarit. Vlastně jde o nestacionaritu v průměru časové řady. Takové nestacionarity (trendy) mohou být dvojího druhu – deterministické, které by se měly v hydrologii modelovat pomocí některé z matematických křivek proložených časovou řadou, nebo stochastické, jež se modelují stochastickými modely s diferenčním parametrem. Při hydrologických a klimatologických aplikacích se lze setkat především s frakcionálně diferencovanými procesy, jejichž důležitým ukazatelem je tzv. Hurstův exponent. V této studii bylo pomocí testů na jednotkový kořen zjištěno, že velmi pravděpodobně řady srážkových úhrnů v Česku neobsahují stochastický trend a že je možné nestacionarity modelovat deterministicky. Také odhadnuté Hurstovy exponenty byly velmi blízké hodnotě 0.5 (medián 0.605), kterou se vyznačuje proces s nezávislými realizacemi. Proto byl v dalších analýzách využit v hydrologii velmi známý Mannův–Kendallův test, který je neparametrický, a tudíž nepřepokládá lineární trend ani přibližnou normalitu testovaných dat. V souvislosti s ním byla využita modifikace, která odstraňuje jeho nedostatek při přítomnosti (krátkodobé) autokorelovanosti v časové řadě. Tento test byl aplikován také sekvenčně pro 30letá období, čímž byla získána infomace o průběhu standardizované testové statistiky za období končící mezi lety 1990 a 2012. Pomine-li se fakt, že čeští hydrologové doposud základní charakteristiky vyjadřují za hydrologická období, jeví se i současné referenční období jako nestacionární z pohledu srážkových úhrnů. U řady za Česko zprůměrovaných ročních úhrnů za roky 1981–2010 se totiž jedná o statisticky významný trend na hladině 5 %.

Srážky v Česku významně rostou v létě (červenci) a v zimě (lednu). Naopak na jaře dochází k poklesům (zejména v dubnu). Při pohledu na rozmístění trendů je možné zjistit, že růsty jsou pozorovány v Čechách, zatímco na Moravě jsou zhruba srážky stabilní či někde klesají. Kromě klasických agregovaných hodnot pro měsíce a roky byly zkoumány též řady různých ukazatelů sezonality. Z nich je patrné, že měsíce s maximy budou ještě více humidnější a období s maximálními úhrny se bude zužovat, tj. soustředit se více do měsíců s maximy. Rovněž se potvrzuje, že tato maxima se stěhují z června do července. Nepříjemnou zprávou je, že srážková intenzita – roční úhrn srážek podělený počtem dnů s úhnrem nejméně 0,1 mm – zjevně roste právě ve vyšších polohách, tj. tam, kde se nacházejí pramenné toky. Toto do budoucna může znamenat zvýšení rizika výskytu přívalových povodní, a to více v Čechách než na Moravě.

Pravdou je, že příspěvek zdůrazňuje spíše metodologii a z tohoto důvodu není interpretace veškerých výsledků vyčerpávající. Předpokládá se ale, že výstupy mohou být nápomocny dalším studiím. Především se nabízí porovnání s trendy v odtokových charakteristikách integrujících vlastnosti povodí. Tím by vlastně technické řady srážek podstoupily jakýsi alternativní test své validity. Dále se nabízí ověření tvrzení, že zprůměrovaná řada ročních úhrnů za celé Česko nevykazuje žádnou periodocitu. K tomuto účelu autor doporučuje analýzu vlnek. V neposlední řadě by se detailnějším analýzám měly podrobit extrémy.

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