Temporal variations of streamflow in a mid-latitude Eurasian steppe watershed in the past half century

Xixi Wang

ABSTRACT

Previous studies either did not identify abrupt change or identified such change but did not exclude it from the detection of trend in streamflow. As a result, an overall downward trend might be erroneously detected as an upward trend because of abrupt increase, while an overall upward trend could be faked as a downward trend due to abrupt decrease. The objectives of this study were to: (1) present a methodology to analyze trend in streamflow in the presence of abrupt change; and (2) use this methodology to detect trend and extreme occurrence of streamflow in the Upper Balagaer River watershed, a mid-latitude nearly pristine precipitation-fed Eurasian steppe watershed in north China. The results indicate that streamflow abruptly decreased around 1994 and exhibited no significant trend from 1960 to 1993 but a significant decrease trend since 1994 (in particular after 1999). In addition, the occurrence of days with a low streamflow was greater after 1994, whereas the occurrence of days with a high streamflow was smaller. Further, the inclusion of the abrupt change in the analysis could compound the detection of the pre-1994 trends but had minimal influences on the detection of the post-1994 trends. These results can be representative across the Eurasian steppe region beyond the study watershed.

Key words | change detection, climate change, Mann–Kendall test, nonstationarity, runoff

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INTRODUCTION

Streamflow (i.e. runoff) constitutes a significant phase of the hydrologic cycle at watershed scale (Mather 1981; USEPA 2013), and its temporal variations have been extensively assessed in order to develop adaptation measures to climate change (Bates *et al.* 2008; Danneberg 2012; Fan *et al.* 2013). Usually, at least 15–25 years of systematic observations are needed to statistically characterize temporal patterns in streamflow (Mather 1981). For example, using the conventional Mann–Kendall test technique (Mann 1945; Kendall 1970), Lins & Slack (1999) examined streamflow trends in the USA. In that study, 395 hydro-climatic data network stations were selected and subdivided into six groups, respectively, with 30, 40, 50, 60, 70, and 80 years records by 1993. Those authors found that the percentage of stations with significant increasing/decreasing trends were larger for the groups with record lengths of 40–70 years than for the groups with record lengths of 30–80 years. However, regardless of the groups, the test revealed that the 395 stations exhibited more increasing trends in the low to moderate flows but roughly equal upward and downward trends in high flows. Those authors concluded that overall, the USA appeared to be getting wetter but less extreme. Nevertheless, the predominant stations in the mid-latitude regions (e.g. northern Iowa) showed significant decreasing trends in flows at any magnitudes.

Kundzewicz & Robson (2004) and Khaliq *et al.* (2009) overviewed methods for identifying trends in hydrologic series and concluded that for change detection at watershed scale, the non-parametric Mann–Kendall test with trend-free-pre-whitening (designated the modified Mann–Kendall test technique for description purposes) (Hamed & Rao 1998; Yue *et al.* 2002; Yue & Wang 2004; Petrow & Merz 2009) is more appropriate than the conventional Mann–Kendall test technique because the latter technique assumes

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independence, which may be usually invalid for streamflow series (Yue *et al.* 2002; Chebana *et al.* 2013). The purpose of pre-whiting is to eliminate the effect of serial correlation on the Mann–Kendall test (Yue & Wang 2004). With this regard, a number of researchers used the modified Mann–Kendall test technique to examine trends in streamflow across the world.

Among them, Lindström & Bergström (2004), Kundzewicz *et al.* (2005), Svensson *et al.* (2005), Mudelsee *et al.* (2006), Petrow & Merz (2009), Stahl *et al.* (2010), and Danneberg (2012) examined flow data of high-latitude (>50° N) European glacier-fed watersheds and found positive trends in streamflow (especially in winter). A similar positive trend was found by Wang *et al.* (2013) in the glacier-fed arid region of northwest China. In contrast, Cunderlik & Ouarda (2009) found that in south Canada, snowmelt floods exhibited significant downward trends at a significance level of $\alpha = 0.05$ while rainfall-dominated floods exhibited either insignificant or marginally significant upward trends. One reason for these opposite trends is that the warmer temperature increased melting of the high-latitude glaciers and made snowmelt earlier in timing. Another possible reason is that human activities (e.g. withdrawing and damming) might have altered the natural hydrology. However, the above studies did not consider possible abrupt changes and thus could invalidate the stationarity assumption of the test technique (Rougé *et al.* 2013; Wang *et al.* 2014) because the mean of a streamflow series with abrupt change(s) is likely inconsistent over a long time period.

On the other hand, some other studies examined abrupt change but did not exclude it from trend analysis. For example, Yang & Tian (2009) and He *et al.* (2013) used the sequential Mann–Kendall test technique (Sneyers 1990) to identify abrupt changes, while this technique is conventionally used to test the null hypothesis that a stationary time series shows no beginning to develop a trend. Subsequently, they used the modified Mann–Kendall test technique to detect trends without excluding the identified abrupt changes. Yang & Tian (2009) found that runoff in the Haihe River catchment of China abruptly declined between 1978 and 1985, and exhibited a significant downward trend due to agricultural water use. Similarly, He *et al.* (2013) found that streamflow along the Yellow River of China exhibited a significant downward trend that can be attributed to warmer climate. Moreover, Zhang *et al.* (2014) used 'ensemble empirical mode decomposition' to identify abrupt changes and 'scanning t and F statistics' to detect trends, and found that the fluctuations and abrupt changes of streamflow in the East River basin of China were largely influenced by both water reservoir constructions and precipitation changes. He *et al.* (2013) and Zhang *et al.* (2014) also used the wavelet transform technique to determine periodicities in streamflow. As with other primary trend analyses presented in existing literature, these three studies did not take into account the nonstationarity resulting from abrupt change, and thus likely confounded or faked the detected trends.

The objectives of this study were to: (1) present a methodology to analyze trend in streamflow in the presence of abrupt change; and (2) use this methodology to detect trend and extreme occurrence of streamflow in the Upper Balagaer River watershed (UBRW) located in northeast China (Figure 1). The methodology seamlessly integrates the most appropriate methods (Chandler & Scott 2011) for identifying abrupt change and detecting trend in streamflow. Although this methodology was first formulated by Wang *et al.* (2014) to detect trends of precipitation in the same watershed, this study is the new application of the methodology in streamflow analysis. The UBRW was selected because it: (1) belongs to the mid-latitude (40–45°N) inland, where effects of climate change on hydrology have been, and will continue to be, evident (Meehl *et al.* 2007; Swann *et al.* 2012); (2) has typical hydrologic characteristics (e.g. small mean streamflow) of arid/semi-arid Eurasian steppe grasslands; and (3) has long-term (i.e. >25 years) observed data on streamflow, which is rare in the Eurasian region. The hydrologic characteristics of this watershed are distinctly different from those located in the other (e.g. south-central) parts of China, such as the watersheds studied by Yang & Tian (2009), Zhang *et al.* (2014), and He *et al.* (2013).

MATERIALS AND METHODS

Study area

The 2,896 km^2 upper portion of the Balagaer River watershed (Figure 1), located in the northeastern Inner Mongolia Autonomous Region of China, was selected for

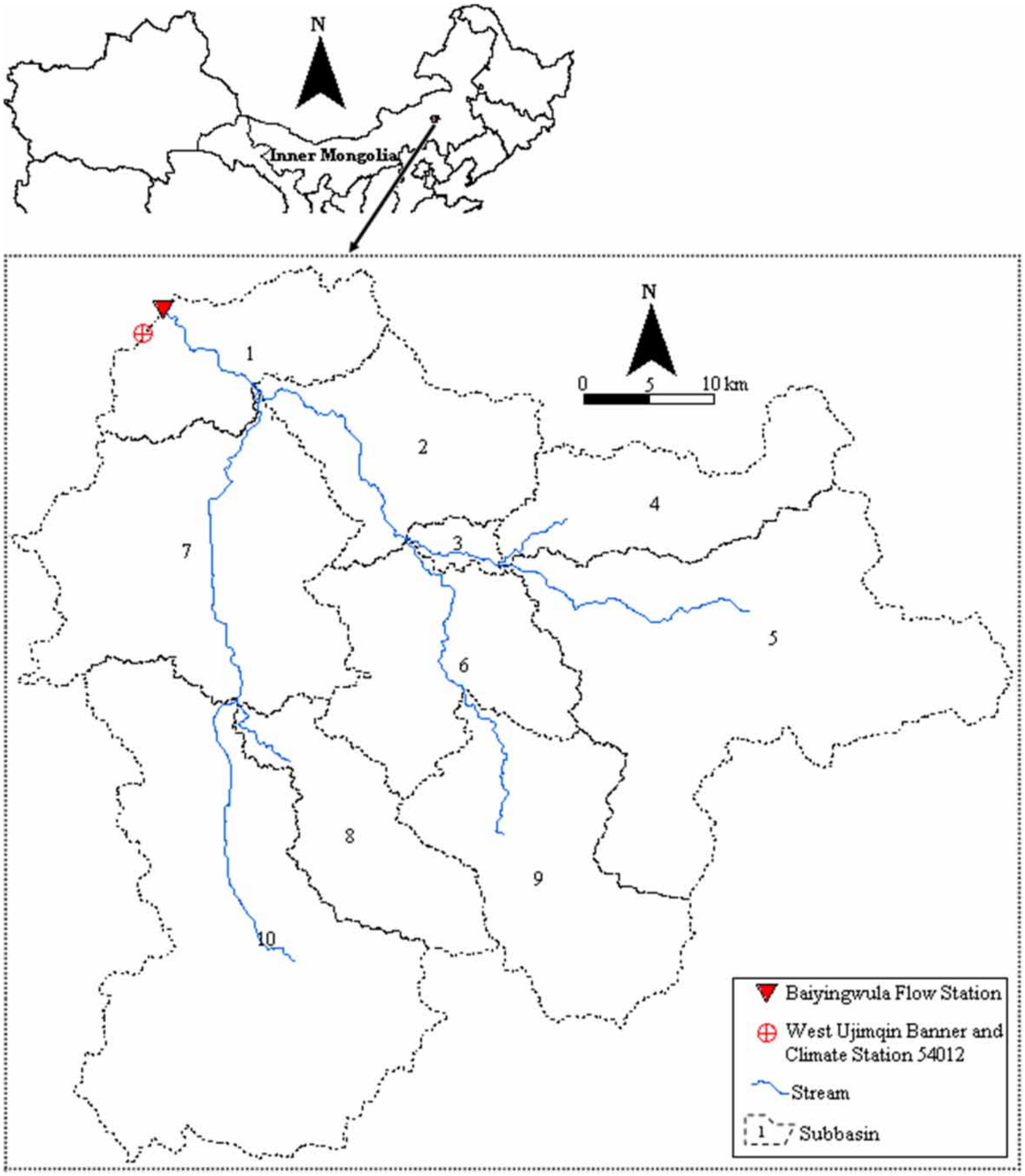

Figure 1 | Map showing the UBRW and its subbasins, along with the flow and climate stations where data were used in this study.

study. The study area (designated as the UBRW for description purposes) is controlled by the Baiyingwula flow gauging station (117°37′ E, 44°36′ N) and fed by precipitation. The UBRW is nearly pristine (i.e. negligibly impacted by human activities) and has a land cover of almost uniform typical steppe grasses of *leymus chinensis* and *stipa grandis* (Wang *et al.* 2012). The elevation of the UBRW varies from 870 to 1,055 m above-mean sea level, leading to a mean topographic gradient of 0.07. The watershed is dominated by loamy soils, with less than 30% clay (particle diameter smaller than 0.002 mm) and 70% sand (particle diameter 0.05–2.0 mm) and silt (particle diameter 0.002–0.05 mm) (McCarthy 2002).

Based on the long-term observations at West Ujimqin Climate Station 54012 (Figure 1), the UBRW receives 170–615 mm precipitation annually, with a mean of 335 mm. Most of the precipitation falls between July and September as rain, and between October and January as snow. However, the watershed has an annual potential evapotranspiration (PET) of 1,165 mm or higher, which is much greater than the mean annual precipitation. Thus, the watershed has a semi-arid climatic condition (Wang *et al.* 2012). The annual mean daily average air temperature is 1.2 °C, with a maximum daily air temperature of up to 37.5 °C in summer and a minimum daily temperature of as low as –38.5 °C in winter. The watershed has an annual mean daily average wind speed of about 15 km h^{-1} and can have 28–148 windy days for a given year. For a given windy day, the maximum wind speed can reach as high as 125 km h^{-1}.

The annual mean discharge at the watershed outlet is 0.86 $m^3 s^{-1}$. The streamflows in the months of March, April, May, October, and November are primarily generated by melting snow, while the streamflows in the months of June, July, and August are mainly generated by rainfall. The three winter months of December, January, and February are completely frozen and have a zero streamflow. The major water users are animal husbandry, domestic, and power industry. Because of a very low population density (<5 people km^{-2}) and small amount of water consumption, the UBRW is nearly pristine with negligible effects of human activities on hydrology. This is similar to the US watersheds studied by Lins & Slack (1999).

Data and preprocessing

Data on daily streamflow at Baiyingwula flow gauging station (Figure 1) were downloaded from the National Meteorological Information Center website (www.nmic.gov.cn) for a record period of 1 January 1959 to 31 December 2011, during which global climate exhibited detectable changes (NIPCC 2011). The station, located near the center of the watershed, is maintained by the China Meteorological Administration (CMA; www.cma.gov.cn) and has never been relocated during the record period. A broad-crested concrete weir was constructed to stabilize the river cross section at the station, while a traditional current meter had been used to measure flow velocities at various depths and locations across the section. Also, in the past 60+ years, there has been no significant construction of hydraulic structures (e.g. diversions and reservoirs) along the Balagaer River and its tributaries as well as throughout the watershed. The data were collected and compiled by well-trained field personnel, and subsequently submitted to the CMA for quality and consistency check by subject experts, making sure that the data have a best quality. The record is complete with no missing value and was considered to be long enough for statistical analysis of trend (Mather 1981; Hamed & Rao 1998; Burn & Hag Elnur 2002; Yue & Wang 2004). Because the streamflows in all winter days were zero, the three winter months (December–February) were excluded out of further analysis.

For a given month in a hydrologic year (September–August) from 1960 to 2011, the daily values were arithmetically averaged to get the mean streamflow for this month. Also, the daily values for the 9 months from March to November of a hydrologic year were arithmetically averaged, and this average value was designated as 'mean annual streamflow' of this hydrologic year for description purposes. In addition, the daily values in the months of a season of this hydrologic year were arithmetically averaged to get the mean streamflow of this season. Herein, the three analyzed seasons are named as spring (March–May), summer (June–August), and fall (September–November). Further, for a given month (excluding December, January, and February) spanning the period from March 1960 to November 2011, the daily values were arithmetically averaged to get the mean streamflow of this month. All these computations were executed in Microsoft® Excel 2007. As a result, nine datasets of annual mean monthly streamflow, three datasets of annual mean seasonal streamflow, and one dataset of mean annual streamflow were generated. This study used these 13 datasets to examine abrupt change, trend, and occurrence of streamflow in the UBRW at annual, seasonal, and monthly time scales.

Trend analysis approach

Figure 2 shows the analysis approach proposed by, and used in, this study. For a given time series (i.e. dataset), the distribution-free cumulative sum (CUSUM) technique proposed by McGilchrist & Woodyer (1975) was used to detect whether, and at which year, a significant abrupt (i.e. upward-to-downward or downward-to-upward) change occurred at a significance level of $\alpha = 0.05$. Herein, the null hypothesis is that there was no abrupt change (i.e. the mean value of the time series is statistically independent of number of computational years). If no significant abrupt change was detected, the modified Mann–Kendall (Mann 1945; Kendall 1970; Hirsch *et al.* 1982; Hamed & Rao 1998) was applied to the entire dataset to determine whether a statistically significant ($\alpha = 0.05$) temporal (i.e. downward or upward) trend existed. Otherwise, the modified Mann–Kendall test was applied independently to the subdatasets before and after the abrupt-change year to determine any trends. In addition, the sequential Mann–Kendall method (Taubenheim 1989; Sneyers 1990) was used to determine the year when a trend started, while the Sen's slope (Sen 1968) was computed and used to measure change rate of the trend.

If $\{x_1, x_2, \ldots, x_N\}$ is a time series with a sample size of N, the CUSUM test statistics, V_k, is computed as (McGilchrist & Woodyer 1975):

$$V_k = \sum_{i=1}^{k} \text{sign}(x_i - x_{\text{median}}) \tag{1}$$

where $k = 1, 2, \ldots, N$; and x_{median} is the median of the time series.

The sign function is defined as:

$$\text{sign}(x_i - x_{\text{median}}) = \begin{cases} 1 & \text{if} \quad x_i > x_{\text{median}} \\ 0 & \text{if} \quad x_i = x_{\text{median}} \\ -1 & \text{if} \quad x_i < x_{\text{median}} \end{cases} \tag{2}$$

At a significance level of $\alpha = 0.05$, the null hypothesis of no abrupt change will be rejected if a maximum or minimum value for V_k exists and $\max\{|V_k|\} > 1.36\sqrt{N}$ is satisfied (Conover 1971; McGilchrist & Woodyer 1975). That is, an abrupt-change point is the year when the maximal or minimal V_k falls out of the 95% confidence limits of $\pm 1.36\sqrt{N}$.

The Sen's slope, Q, for the time series is computed as (Sen 1968):

$$Q = \begin{cases} T_{\frac{m+1}{2}} & \text{if} \quad m = \dfrac{N(N-1)}{2} \text{ is odd} \\ \dfrac{T_{\frac{m}{2}} + T_{\frac{m}{2}+1}}{2} & \text{if} \quad m = \dfrac{N(N-1)}{2} \text{ is even} \end{cases} \tag{3}$$

T_i is computed as:

$$T_i = \frac{x_j - x_k}{j - k} \tag{4}$$

where $i = 1, 2, \ldots, \dfrac{N(N-1)}{2}$; $k = 1, 2, \ldots, N\text{–}1$; $j = k+1, k+2, \ldots, N$.

The modified Mann–Kendall test takes autocorrelation into account (Hamed & Rao 1998) and is executed by seven sequential steps: (1) compute the Sen's slope of the time series using Equations (3) and (4); (2) subtract the Sen's slope from the original values to obtain a trend-free pre-whitening time series of $\{x'_1, x'_2, \cdots, x'_N\}$ and to assign ranks to these new values; (3) compute autocorrelation coefficients, $\rho_s(i)$, $i = 1, 2, \ldots, N\text{–}3$, using the ranks (Sen 1968; Zetterqvist 1991) and keep the statistically significant coefficients at a significance level of $\alpha = 0.05$ based on a t-test statistics (Equation (5)); (4) compute the ratio of $\dfrac{N}{N_S^*}$

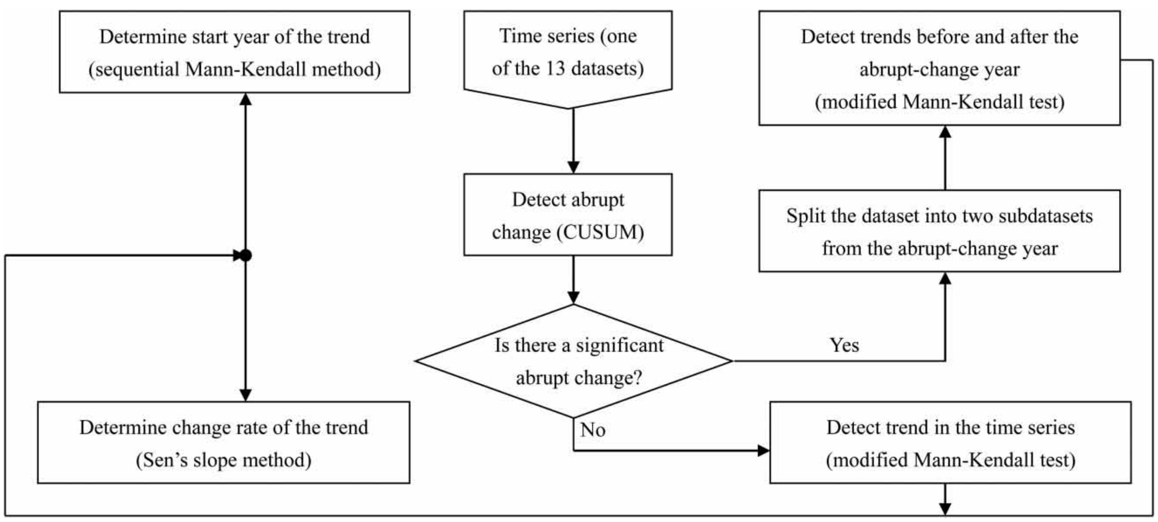

Figure 2 | Schematic diagram of the trend analysis approach proposed by, and used in, this study.

(Equation (6)); (5) compute variance $\mathrm{var}^*(S)$ (Equation (7)); (6) compute the conventional Mann–Kendall statistics of S (Equations (8) and (9)) (Mann 1945; Kendall 1970); and (7) compute the modified Mann–Kendall test statistics of Z^* (Equation (10)).

If Z^* is positive and greater than the standard normal percentile of $1\text{–}\alpha = 0.95$, a statistically significant upward trend was detected. In contrast, if Z^* is negative and smaller than the percentile of $\alpha = 0.05$, a statistically significant downward trend was detected. Otherwise, no significant trend was detected.

$$t_i = \frac{\rho_s(i)}{\sqrt{\dfrac{1-[\rho_s(i)]^2}{N-i-2}}}\,(i = 1, 2, \cdots, N-3) \tag{5}$$

$$\frac{N}{N_S^*} = 1 + \frac{2}{N(N-1)(N-2)} \sum_{i=1}^{N-3} [(N-i)(N-i-1)(N-i-2)(\rho_s(i))] \tag{6}$$

$$\mathrm{var}^*(S) = \frac{N(N-1)(2N+5) - \sum_{p=1}^{N_p} \left[m_p(m_p-1)(2m_p+5)\right]}{18} \cdot \left(\frac{N}{N_s^*}\right) \tag{7}$$

where N_p is the number of distinct ties; and m_p is the number of tied data points in pth tie.

$$S = \sum_{k=1}^{N-1} \sum_{j=k+1}^{N} \mathrm{sign}(x_j - x_k) \tag{8}$$

$$\mathrm{sign}(x_j - x_k) = \begin{cases} 1 \text{ if } & x_j > x_k \\ 0 \text{ if } & x_j = x_k \; (k = 1, 2, \cdots, N-1; j = k+1, k+2, \cdots, N) \\ -1 \text{ if } & x_j < x_k \end{cases} \tag{9}$$

$$Z^* = \begin{cases} \dfrac{S-1}{\sqrt{\mathrm{var}^*(S)}} \text{ if } S > 0 \\ 0 \text{ if } S = 0 \\ \dfrac{S+1}{\sqrt{\mathrm{var}^*(S)}} \text{ if } S < 0 \end{cases} \tag{10}$$

The sequential Mann–Kendall method (Sneyers 1990) tests the null hypothesis that a stationary time series shows no beginning to develop a trend. It computes two statistics: one is for the time series in progressive order (i.e. from the start to the end year), while another is for the time series in retrograde order (i.e. from the end to the start year). When these two statistics are plotted versus year, the intersection between these two curves, from which onward one of these two statistics is either monotonically smaller or greater than another, is judged to be the beginning of the trend (Figure 3). Otherwise, the trend is considered to continue throughout the entire period of the time series.

The progressive statistics, $u(pt_k)$, $k = 1, 2, \ldots, N\text{–}1$, is computed as (Sneyers 1990):

$$u(pt_k) = \frac{pt_k - E(pt_k)}{\sqrt{\mathrm{var}(pt_k)}} \tag{11}$$

where pt_k is an asymptotically distributed random variable; $E(pt_k)$ is the mean of pt_k; and $\mathrm{var}(pt_k)$ is the variance of pt_k.

pt_k is computed as:

$$pt_k = \sum_{j=k+1}^{N} n_j \quad (k = 1, 2, \ldots, N-1) \tag{12}$$

where $n_j = 1$ if $x_j > x_k$ while $n_j = 0$ otherwise.

$E(pt_k)$ is computed as:

$$E(pt_k) = \frac{k(k-1)}{4} \tag{13}$$

$\mathrm{var}(pt_k)$ is computed as:

$$\mathrm{var}(pt_k) = \frac{k(k-1)(2k+5)}{72} \tag{14}$$

The retrograde statistics, $u(rt_k)$, $k = N, N\text{–}1, \ldots, 2$, is computed as (Sneyers 1990):

$$u(rt_k) = \frac{rt_k - E(rt_k)}{\sqrt{\mathrm{var}(rt_k)}} \tag{15}$$

where rt_k is an asymptotically distributed random; $E(rt_k)$ is the mean of rt_k; and $\mathrm{var}(rt_k)$ is the variance of rt_k.

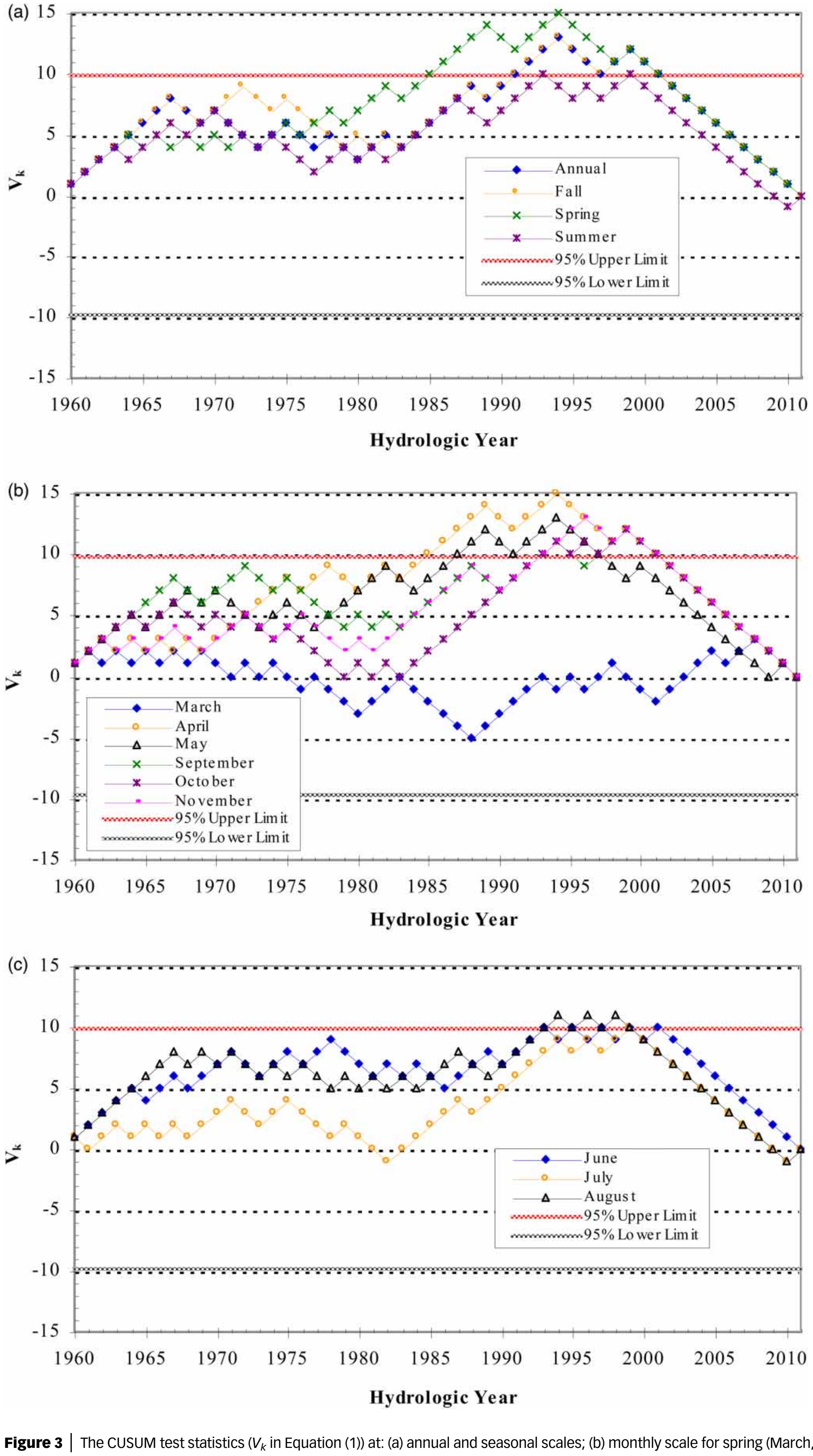

Figure 3 | The CUSUM test statistics (V_k in Equation (1)) at: (a) annual and seasonal scales; (b) monthly scale for spring (March, April, and May) and fall (September, October, and November) months; and (c) monthly scale for summer (June, July, and August) months.

rt_k is computed as:

$$pt_k = \sum_{j=k-1}^{1} n_j \quad (k = N, N-1, \ldots, 2) \tag{16}$$

where $n_j = 1$ if $x_j > x_k$ while $n_j = 0$ otherwise.

$E(rt_k)$ is computed as:

$$E(rt_k) = \frac{(N-k)(N-k+1)}{4} \tag{17}$$

$\mathrm{var}(rt_k)$ is computed as:

$$\mathrm{var}(rt_k) = \frac{(N-k)(N-k+1)[2(N-k)+5]}{72} \tag{18}$$

Extreme occurrence analysis method

The dataset of mean annual streamflow was used to compute five quartiles of minimum (Q_{min}), 25% (Q_{25}), 50% (Q_{50}), 75% (Q_{75}), and maximum (Q_{max}) (Neter *et al.* 1996). Because of the possible nonstationarity in streamflow, existing frequency analysis methods (e.g. the Bulletin 17B approach; Thomas 1985; Viessman & Lewis 2003) became invalid and could not be used to determine frequency flows (Hirsch & Ryberg 2011; Vogel *et al.* 2011). For this reason, this study used these five quartiles as surrogates of frequency flows (Lins & Slack 2005) to represent long-term average extremes across the entire record years from 1960 to 2011. Herein, Q_{min} and Q_{25} were used to represent below-average flows, whereas Q_{75} and Q_{max} were used to represent above-average flows. Subsequently, at the annual and seasonal scales, the percentages of days with streamflows either less than or greater than Q_{50} were computed and assessed for any differences before and after an abrupt change year if any.

RESULTS

Abrupt change

The annual streamflow exhibited an abrupt change around 1994 (Figure 3(a)), when V_k has a maximum value of greater than the 95% confidence upper limit (13 versus 9.8). This is also evidenced by examining the plot showing observed annual streamflow versus hydrologic year (Figure 4). The streamflow rapidly dropped from 2.85 m^3 s^{-1} in 1993 to 1.33 m^3 s^{-1} in 1994, and then to 0.66 m^3 s^{-1} in 1995. The coefficient of variation (C_v, the ratio of standard deviation to mean) of the streamflows from 1960 to 1993 is much smaller than that of the streamflows from 1994 to 2011 (0.51 versus 0.73), indicating that the prior-1993 streamflows exhibited a random pattern but the post-1993 streamflows showed a consistent pattern. The annual average streamflow from 1960 to 1993 was more than twice higher than that from 1994 to 2011 (1.34 versus 0.65 m^3 s^{-1}).

At seasonal scale, similar abrupt changes occurred around 1994 for streamflow in spring and fall, but no abrupt change was detected for streamflow in summer (Figures 3(a) and 5). The abrupt change in spring was mainly caused by the significant changes of streamflow in April and May, while the abrupt change in fall was mainly caused by the changes of streamflow in October and November (Figure 3(b)). In contrast, because only marginally significant abrupt change occurred in streamflow in August and there were no abrupt changes in June and July (Figure 3(c)), the streamflow in summer did not exhibit a significant abrupt change at a significance level of $\alpha = 0.05$. The abrupt change in spring was stronger (i.e. more significant), but the abrupt change in fall was almost as significant as that at annual scale, as indicated by the value of max$\{V_k\}$ for spring streamflow being larger than those for fall and annual streamflows while the value of max$\{V_k\}$ for fall streamflow is almost the same as that for annual streamflow.

Given these seasonal abrupt changes, the annual abrupt change was most likely caused by sudden decreases of snowmelt runoff in late spring and rainfall/snowmelt runoff in late fall. However, because significant abrupt changes occurred in April and May (late spring) and October and November (late fall) (Figure 3(b) and 3(c)), it was judged that the change of snowmelt runoff contributed more to the annual abrupt change than the change of rain-generated runoff. As stated above, runoffs of the UBRW in these four months is mainly generated by melting snow.

Temporal trend

The significant abrupt changes were found to confound or fake the detected prior-1994 trends regardless of time

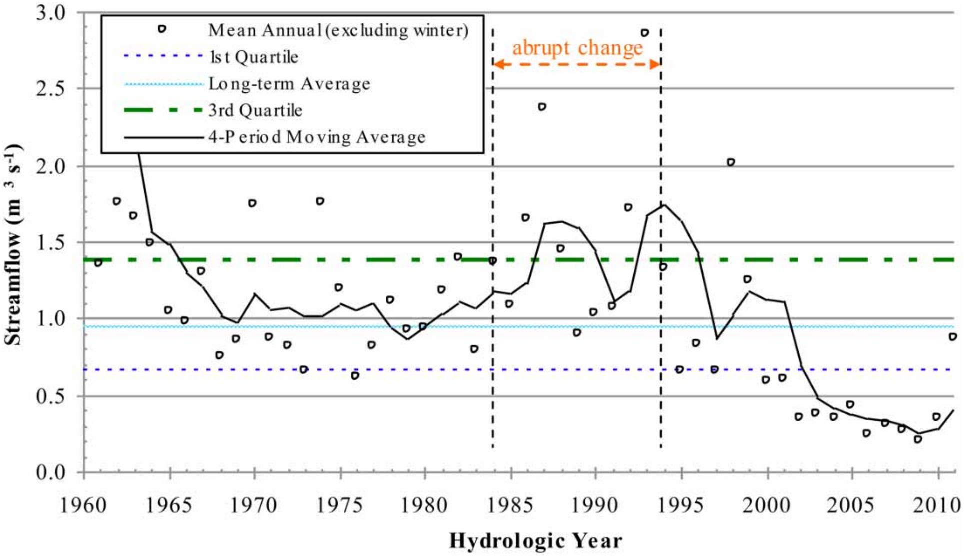

Figure 4 | Plot showing the observed mean annual (including spring, summer and fall only) streamflow, superimposed by a four-period moving average curve and the selected frequency streamflows. Herein, a hydrologic year is defined as from 1 September to 31 August.

scale, but did not influence the identification of post-1994 trends (Table 1). At annual scale, for the entire record, the modified Mann–Kendall test indicated a significant downward trend (p-value = 0.000+). However, for the pre-1994 record, the test indicated no significant trend (p-value = 0.929), while for the post-1994 record, the test indicated a significant downward trend (p-value = 0.001). Similarly, at seasonal scale, significant downward trends for the entire record (p-value ≤ 0.005) were detected, whereas there were no significant trends for the pre-1994 record (p-value ≥ 0.423), but significant downward trends for the post-1994 record (p-value ≤ 0.012), were detected. At monthly scale, when a month (e.g. April) was detected to have a significant downward trend for the entire record, no significant trend for the pre-1994 record, but a significant downward trend for the post-1994 record, of this month was determined. On the other hand, when a month was detected to have no significant trend for the entire record, no significant trend was found either for the pre- or post-1994 record.

Unlike the abrupt changes, the significant post-1994 downward trend at annual scale was probably a combined result of both decreasing rain-generated and snowmelt runoff. The streamflows in June, August and September, when runoff is primarily generated by rainfall, exhibited significant downward trends. Also, the streamflows in April, May, October and November, when runoff is predominantly generated by melting snow, showed significant downward trends too. Overall, since 1994, the streamflow had been decreasing by 0.041 m^3 s^{-1} per year. This is equivalent to a loss of 1.3×10^6 m^3 water resources every year, which will likely increase the challenge in sustaining the animal husbandry and economy of this watershed if such decreasing trend in streamflow continues in the future.

When the entire record for a given time scale was used for trend analysis, the trend (if significant) was determined to start either around 1994 or in another year (Table 2 and Figures 6(a) and 7(a)). This implies that the commonly used trend analysis methods, including the modified and sequential Mann–Kendall test techniques, cannot always differentiate an abrupt change point from trend start year. Alternatively, the approach proposed by this study and illustrated in Figure 2 can overcome this issue. The post-1994 downward trends started from a year between 1997 and 2001 (the last column of Table 2), depending on time scale. The post-1994 annual downward trend was determined to start from 1999 (Figure 6(b)), which was probably triggered by the downward trends of August–November that started in or before 1999 (Table 2). The downward trends of these four months also triggered the downward trends in summer (Figure 7(b)) and fall (Figure 8(a)). The downward trends in April and June started later, while the downward

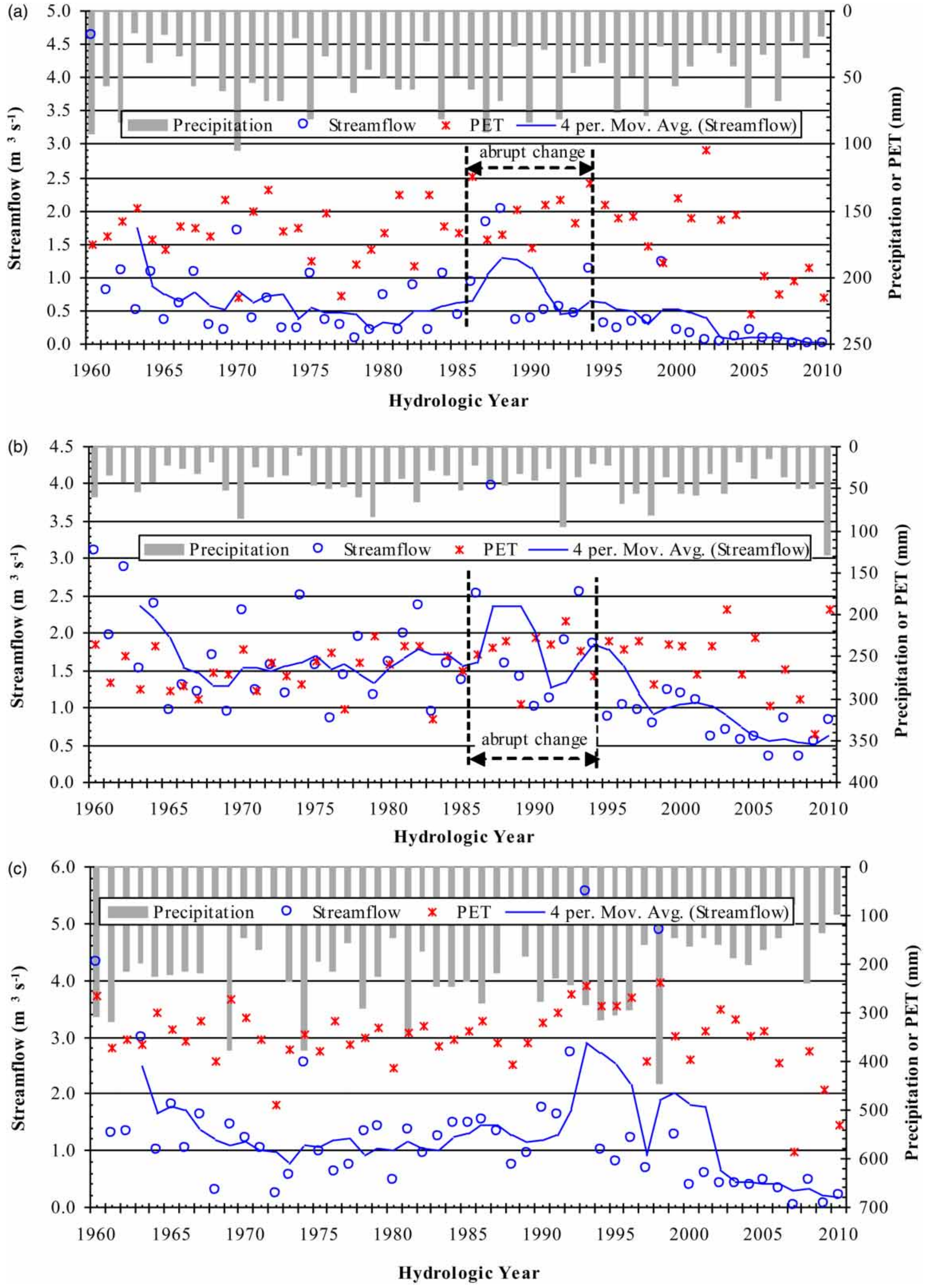

Figure 5 | Plots showing the seasonal streamflow, precipitation, and PET for: (a) fall; (b) spring; and (c) summer. Herein, a hydrologic year is defined as from 1 September to 31 August.

trend in May continued throughout the record year. The downward trend in spring (Figure 8(b)) started from 2000, which was likely triggered by the combined downward trends of April and May.

Extreme occurrence

Based on the mean annual streamflows from 1960 to 2011, the five quartiles were determined to be $Q_{\min} = 0.20\ \mathrm{m^3\ s^{-1}}$,

Table 1 | Results of the modified Mann–Kendall tests[a]

		Entire record[c]			Pre-1994 record			Post-1994 record		
Time scale	**Specific scale**	**Q[d]**	**Z*[e]**	***p*-value**	**Q[d]**	**Z*[e]**	***p*-value**	**Q[d]**	**Z*[e]**	***p*-value**
Annual	Hydrologic year[b]	–0.019	–4.080	**0.000**	0.002	0.089	0.929	–0.041	–3.257	**0.001**
Seasonal	Spring	–0.027	–4.743	**0.000**	–0.003	–0.208	0.836	–0.044	–3.106	**0.002**
	Summer	–0.019	–2.833	**0.005**	0.009	0.801	0.423	–0.049	–2.500	**0.012**
	Fall	–0.014	–4.412	**0.000**	–0.004	–0.741	0.459	–0.023	–3.964	**0.000**
Monthly	March	0.000	–0.387	0.698	0.000	–0.475	0.635	–0.004	–0.722	0.471
	April	–0.053	–4.190	**0.000**	–0.007	–0.237	0.813	–0.099	–2.651	**0.008**
	May	–0.023	–4.569	**0.000**	–0.004	–0.504	0.614	–0.036	–2.273	**0.023**
	June	–0.017	–3.685	**0.000**	0.002	0.000	1.000	–0.045	–2.651	**0.008**
	July	–0.011	–1.365	0.172	0.029	1.542	0.123	–0.044	–1.591	0.112
	August	–0.022	–3.551	**0.000**	–0.006	–0.237	0.813	–0.033	–2.540	**0.011**
	September	–0.019	–4.362	**0.000**	–0.008	–0.801	0.423	–0.027	–3.562	**0.000**
	October	–0.017	–4.523	**0.000**	–0.007	–0.860	0.390	–0.037	–4.193	**0.000**
	November	–0.005	–3.970	**0.000**	–0.002	–0.593	0.553	–0.011	–3.812	**0.000**

[a]The tests in bold fonts are statistically significant at a significance level of $\alpha = 0.05$.
[b]From 1 September to 31 August.
[c]Calendar years from 1959 to 2011.
[d]Sen's slope (in $m^3\ s^{-1}\ yr^{-1}$) defined in Equation (3).
[e]Modified Mann–Kendall statistics defined in Equation (10).

Table 2 | Start years of the significant trends in Table 1 at a significance level of $\alpha = 0.05$[a]

		Entire record[c]		Pre-1994	Post-1994
Time scale	**Specific scale**	**Start year**	**Due to the abrupt change?**	**Start year**	**Start year**
Annual	Hydrologic year[b]	1999	No	–	1999
Seasonal	Spring	1993	Yes	–	2000
	Summer	2000	No	–	1999
	Fall	1999	No	–	1999
Monthly	March	–	–	–	–
	April	1994	Yes	–	2007
	May	1982	No	–	none
	June	1999	No	–	2001
	July	–	–	–	–
	August	1999	No	–	1998
	September	1994	Yes	–	1999
	October	1994	Yes	–	1999
	November	1994	Yes	–	1997

[a]'–' means insignificant trend.
[b]From 1 September to 31 August.
[c]Calendar years from 1959 to 2011.

$Q_{25} = 0.65\ m^3\ s^{-1}$, $Q_{50} = 0.96\ m^3\ s^{-1}$, $Q_{75} = 1.38\ m^3\ s^{-1}$, and $Q_{max} = 4.02\ m^3\ s^{-1}$. The chance for a year with a mean annual streamflow of smaller than $Q_{50} = 0.96\ m^3\ s^{-1}$ was greater after 1994 than before this year (Figure 5). This is also true for daily streamflows (Table 3). On average, 51% of the post-1994 days, but only 35% of the pre-1994 days, had a streamflow less than $Q_{min} = 0.20\ m^3\ s^{-1}$. In contrast, the percentage of the days with a streamflow of larger than Q_{50} was determined to be much greater before 1994 than after this year regardless of time scale. However, the occurrences of the daily spring streamflows in any magnitudes were roughly the same before and after the abrupt change year.

DISCUSSION

Recently, Wang *et al.* (2014) found that both annual and summer precipitations in the UBRW (study watershed) had an abrupt increase between 1993 and 1994. As expected, the abrupt change of annual precipitation likely caused the abrupt change of annual streamflow around 1994. However, the abrupt change of summer precipitation did not result in an abrupt increase of summer streamflow, which seems inconsistent with the general hydrologic principles (Viessman & Lewis 2003). This may be one of the specific hydrologic characteristics of the Eurasian steppe grassland. In summer, the transpiration demand tends to exponentially increase with precipitation (Figure 5) because of the better growth of

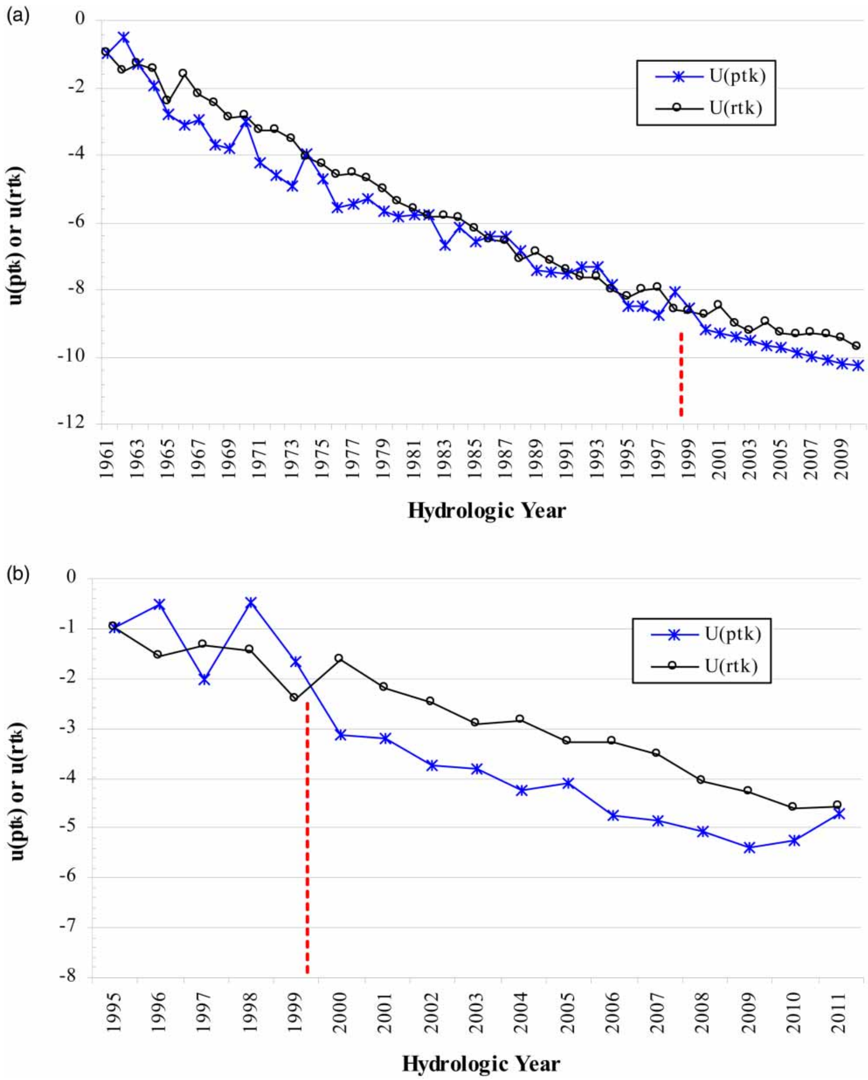

Figure 6 | Plots showing the sequential Mann–Kendall test statistics $u(pt_k)$ (Equation (11)) and $u(rt_k)$ (Equation (14)) of annual streamflow for the: (a) record year, and (b) hydrologic years from 1994 to 2011. Hydrologic year is defined as from 1 September to 31 August. The intersects between $u(pt_k)$ and $u(rt_k)$ curves indicate the start years of significant trends at a significance level of $\alpha = 0.05$.

steppe grasses (Jiang *et al.* 1999), which in turn can reduce the soil moisture to a level at which soil storage becomes phenomenal. As a result, the abruptly increased precipitation in summer might primarily be used for increased transpiration and soil storage, rather than being converted into runoff. Albeit, the remediation of soil storage in summer probably elevated the antecedent soil moistures of fall and spring and thus resulted in the abrupt increases of streamflow around 1994 in these two latter seasons.

In addition, as with precipitation (Wang *et al.* 2014), streamflow had no trends before 1994 regardless of time scale (Tables 1 and 2), whereas both annual and summer streamflow exhibited significant ($\alpha = 0.05$) downward trends after 1994. This indicates that the study watershed likely has a sophisticated nonlinear precipitation-runoff relationship. Moreover, because the watershed is nearly pristine and has experienced negligible human activities/developments in the past centuries (Wang *et al.* 2012), it was judged that the detected abrupt changes and significant downward trends can mainly be attributed to natural rather than anthropogenic factors. That is, the changes and trends were probably either a result or an indicator of climate change.

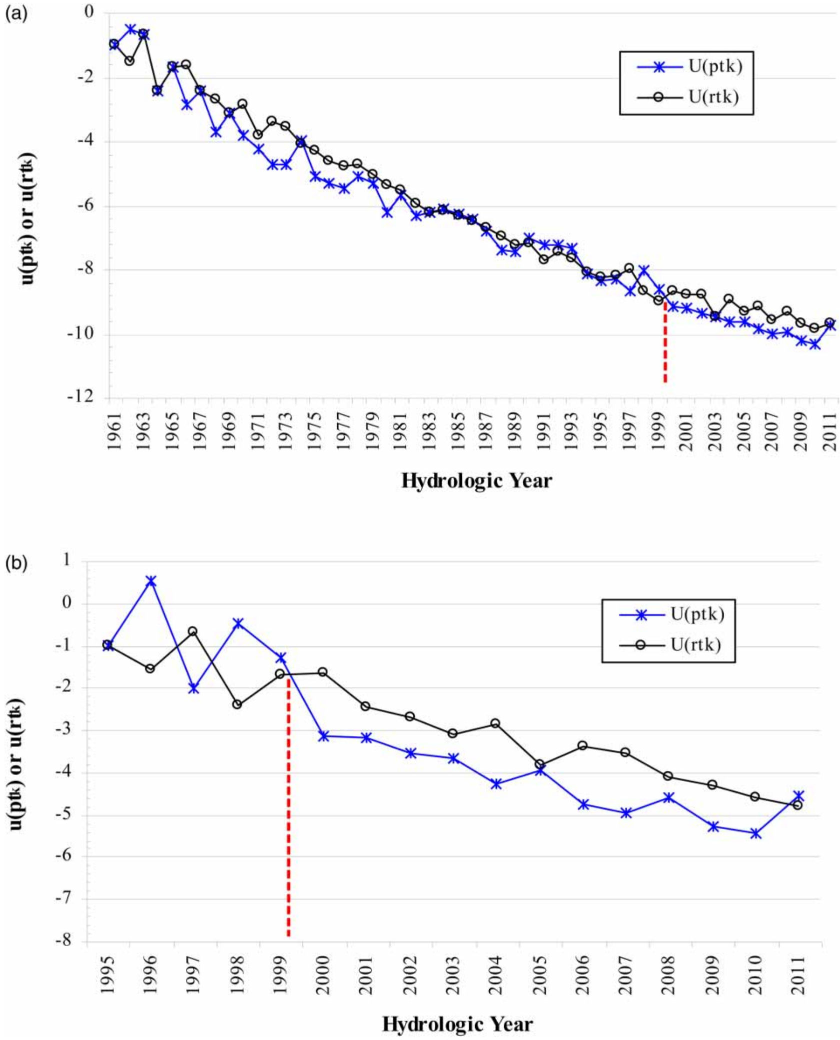

Figure 7 | Plots showing the sequential Mann–Kendall test statistics $u(\mathrm{pt}_k)$ (Equation (11)) and $u(rt_k)$ (Equation (14)) of seasonal streamflow in summer for (a) record year and (b) hydrologic years from 1994 to 2011. Hydrologic year is defined as from 1 September to 31 August. The intersects between $u(pt_k)$ and $u(rt_k)$ curves indicate the start years of significant trends at a significance level of $\alpha = 0.05$.

As mentioned above, few stations in the Eurasian steppe region have long-term (>25 years) high-quality streamflow data. The stations with good-quality data primarily monitor glacier-fed watersheds (e.g. Yang *et al.* 2004; Petrow & Merz 2009; Danneberg 2012; Wang *et al.* 2013), which have distinctly different hydrologic characteristics from the precipitation-fed watersheds with negligible effects of human activities, including the UBRW. For this reason, this study only analyzed the 51-years flow data at the single station. Nevertheless, given the similar topography, land cover and land use, and climate of steppe grasslands within a large spatial scope (Wang *et al.* 2012), the results of this study can still provide some insightful inferences of streamflow trends in the precipitation-fed areas of the Eurasian steppe region. Lins & Slack (2005) found that for the nearly pristine watersheds in two US prairie water-resources regions of Missouri and Upper Mississippi, the below-average streamflows tended to occur more frequently with the lapse of year, whereas the above-average streamflows tended to occur less frequently. This is consistent with the finding of extreme occurrences in this study: the chance for a year with a below-mean annual streamflow was greater, but the chance for a day with an above-mean streamflow was smaller, after 1994. As with the UBRW, those two water-resources regions

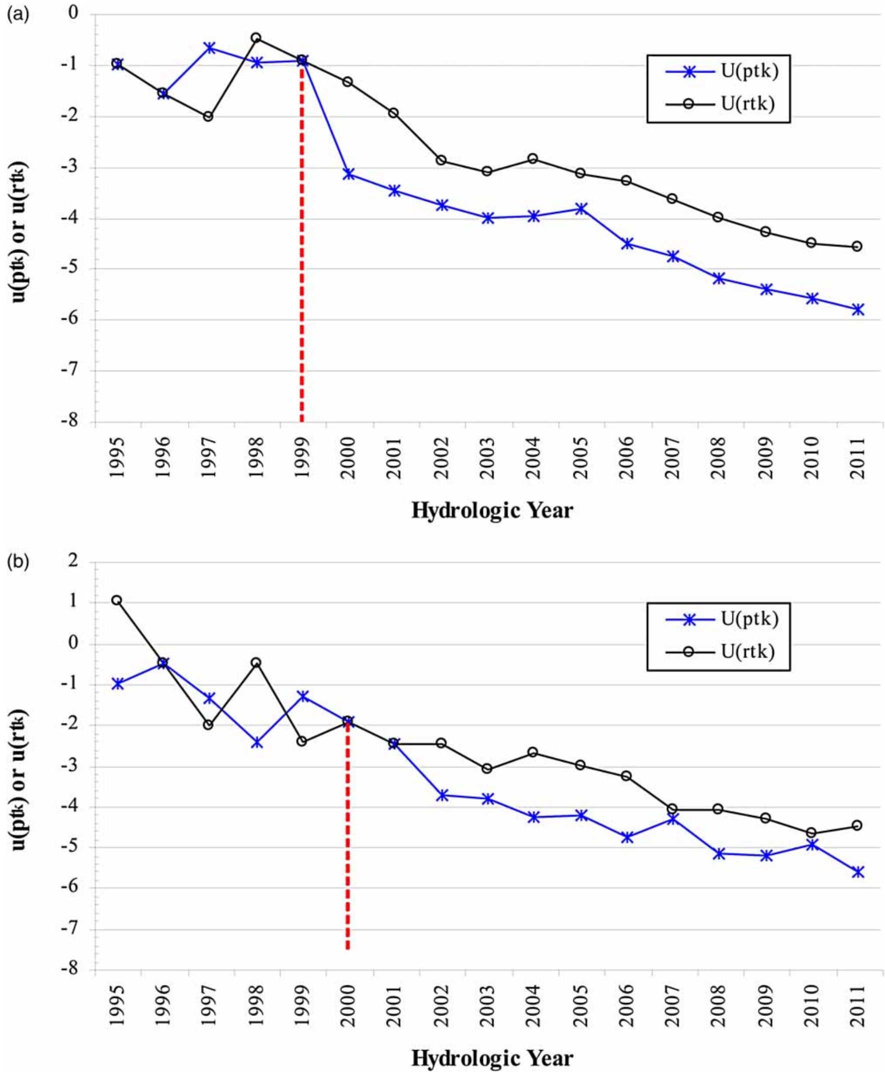

Figure 8 | Plots showing the sequential Mann–Kendall test statistics $u(pt_k)$ (Equation (11)) and $u(rt_k)$ (Equation (14)) of seasonal streamflow in (a) fall and (b) spring, during the hydrologic years from 1994 to 2011. Hydrologic year is defined as from 1 September to 31 August. The intersects between $u(pt_k)$ and $u(rt_k)$ curves indicate the start years of significant trends at a significance level of $\alpha = 0.05$.

have mid latitude, and are primarily fed by precipitation, covered by grasses and nearly pristine. Thus, the consistency of the findings indicates that the trends detected by this study are likely representative beyond the study watershed though a single flow station was used.

CONCLUSIONS

This study presented a methodology that can be used to differentiate abrupt change from trend in streamflow time series. The new application of this methodology in the UBRW revealed that streamflow abruptly decreased around 1994. If this abrupt change was included in trend analysis, it could fake the detected pre-1994 trends and start years but had minimal influences on the detection of the post-1994 trends. The streamflow from the watershed exhibited no significant trend from 1960 to 1993, while it showed a significant decrease trend since 1994 (in particular after 1999). In addition, the occurrence of days with a below-average (i.e. low) streamflow was greater after 1994 than before this year, whereas the occurrence of days with

Table 3 | Occurrence of daily streamflow[a]

Period and time scale	Percent of days with a streamflow (m^3 s^{-1}) <= Q_{min}	<= Q_{25}	>= Q_{50}	>= Q_{75}	>= Q_{max}
1960–1993	35	49	37	26	6
Spring	88	91	3	2	0
Summer	2	16	65	46	11
Fall	14	42	43	30	7
1994–2011	51	75	17	11	2
Spring	91	95	2	1	0
Summer	10	53	31	19	2
Fall	51	77	17	12	4

[a]Q_{min}, Q_{25}, Q_{50}, Q_{75}, and Q_{max} are five quartiles (respectively, minimum, first quartile, second quartile/mean, third quartile, and maximum) of the mean annual streamflows from 1960 to 2011. $Q_{min} = 0.20$ m^3 s^{-1}; $Q_{25} = 0.65$ m^3 s^{-1}; $Q_{50} = 0.96$ m^3 s^{-1}; $Q_{75} = 1.38$ m^3 s^{-1}; and $Q_{max} = 4.02$ m^3 s^{-1}.

an above-average (i.e. high) streamflow was smaller. This implies that the available water resources to sustain the steppe grasses of the watershed had been decreasing, which would likely threaten the animal husbandry and vulnerable economy if no effective adaptation measures were taken. The consistency with the findings of Lins & Slack (2005) indicates that the trends detected by this study are likely representative across the Eurasian steppe region beyond the study watershed though a single flow station was used.

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