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A proxy record of winter temperatures since 1836 from ice freeze-up/breakup in lake Näsijärvi, Finland

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Abstract One obstacle on the way to a comprehensive spatial reconstruction of regional temperature changes over the past centuries is the sparseness of long winter temperature records. This paper reconstructs a proxy record of April and November–December temperatures in south-central Finland for the interval from 1836 to 1872 from breakup and freeze-up dates and ice-cover duration of a lake. Emphasis is on detecting the suitable winter months and quantifying the calibrations with measured temperatures (1873–2002). The calibration slope for the breakup date (0.158°C/day) is larger than for freeze-up date (0.119°C/day) or duration (0.090°C/day). A comparison with results from other proxy records shows that the slope may depend also on the geographical site. Trend analyses of the full temperature records (1836–2002) indicate the existence of minor change-points at around 1867 (April temperature) and 1874 (November–December temperature), with warming rates thereafter of 1.67°C per century (April) and 1.16°C per century (November–December). Spectral analyses reveal peaks in the band between 2 and 5 year period, which may point to influences of the North Atlantic Oscillation, and less power in the decadal band (up to 42 year period).

Keywords Documentary data · Errors-in-variables regression · Block bootstrap resampling

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1 Introduction

Paleoclimatologists are creative in tapping climate archives and measuring proxy variables (Bradley 1999). This type of information about past climate dynamics helps also climatologists wishing to extend directly observed records further back in time. Accurate proxy calibrations require long paired time series (proxy, observations) and a suitable statistical methodology.

One obstacle on the way to a comprehensive spatial reconstruction of regional temperature changes over the past centuries, which hampers also efforts of a global temperature estimation (Mann et al. 1998), is the sparseness of winter temperature records. The tree-ring archive (Bradley 1999) is useless here since trees grow mostly in spring and summer, and tree-ring width is therefore not a proxy for winter temperatures. Observed dates of ice breakup and freeze-up in rivers and lakes—events that happen in winter—can help in this situation. This type of archive, documentary climate data, has the potential to extend our view back by several centuries for Europe (Weikinn 1958–2002; Pfister 1999; Tarand and Nordli 2001; Brázdil et al. 2005) and perhaps also Asia, and by a few centuries for North and South America.

So far, ice breakup or freeze-up dates in rivers or lakes have not been widely used in quantitative paleoclimatology. The Working Group I of the Intergovernmental Panel on Climate Change (IPCC–WG I) lists in its latest report (Lemke et al. 2007) just two papers on lakes (and two on rivers): the first (Magnuson et al. 2000) on the northern hemisphere with a focus on the interval from the mid-19th to the end of the twentieth century, the second (Duguay et al. 2006) on Canada and the interval from 1951 to 2000. The general impression (Lemke et al. 2007) is that a tendency exists towards earlier breakup and later freeze-up, compatible with regional warming trends (Trenberth et al. 2007). Other papers studied other regions, such as Switzerland (Franssen and Scherrer 2008), or looked at potential drivers (also other than temperature) of the dates of those ice events (Yoo and D'Odorico 2002; Prowse et al. 2007).

Proxy measurements on climate archives, in general, record not only climatic variations but inevitably also other influences. For example, ice breakup dates may depend not only on temperature but also on the amount of snowfall via the insulating effect of snow (Lemke et al. 2007). Inasmuch those influences cannot be taken into account (because they are unknown or poorly constrained by modelling, or measurements do not exist), they disturb the calibration between the proxy and the climate variable and introduce "proxy noise."

Since most studies on ice breakup and freeze-up dates focused so far on temporal trends, there is a lack of calculated calibrations (taking proxy noise into account) between the dates and regional temperatures. This deficit is considerable since it is not a priori clear to which winter month such an event is related via temperature. Furthermore, the calibration parameters may depend on the geographical position and other factors such as size of a lake, surrounding orography and the degree of proxy noise.

The present paper reconstructs a proxy record of winter temperatures from breakup and freeze-up dates of a lake in Finland. I put emphasis on detecting the suitable winter months and quantifying the calibrations by fully taking into account proxy noise and other potential statistical issues (non-normal distributional shapes, autocorrelation and errors-in-variables regression). Finally, I analyse deterministic trends and evidence for cyclical forcing factors in the long (time interval 1836–2002) winter temperature records. This is a pilot study on a single geographical site. The methodical tools described here, however, should facilitate studying other sites as well and thereby obtaining a clearer picture of winter temperature trends in space and time over the past centuries.

2 Data

Näsijärvi is a lake in southern Finland (Fig. 1), of size 257 km^2 and at altitude 95 m.a.s.l. Winter conditions at this site have, at least over the past two centuries, been such that freezing occurs and a proxy record can be obtained. Official observations of the ice conditions are made by the harbour master at Tampere, the city south of Näsijärvi (Fig. 1); supporting information comes from fishermen. The breakup date is defined as the day from which on the two major parts of the lake, Koljonselkä and Näsiselkä (Fig. 1), are navigable. Minor bays may still be ice-covered. The freeze-up date is defined as the day from which on the two parts are completely ice-covered. Accuracy of the estimated dates is 1 day (Matti Joki, personal communication, 2005); a reexamination showed that they fall also on Saturdays and Sundays. The reproduction of the original list by Matti Joki (Mustonen and Nieminen 2004, pp. 207–208 therein) gives the dates for the interval from 1836 to 2004, with the following winters missing: 1882/83 (breakup), 1883/84 (freeze-up) and 1885/86 (breakup). Two dates (24 November and 31 December 1992) are given for winter 1992/93 (freeze-up); the analyses use the later date and study the effects of using instead the earlier date. Time series of breakup and freeze-up dates are shown in Fig. 2d, e, respectively, as number of days relative to 1 January; also shown is the time series of the duration of the ice cover (Fig. 2f).

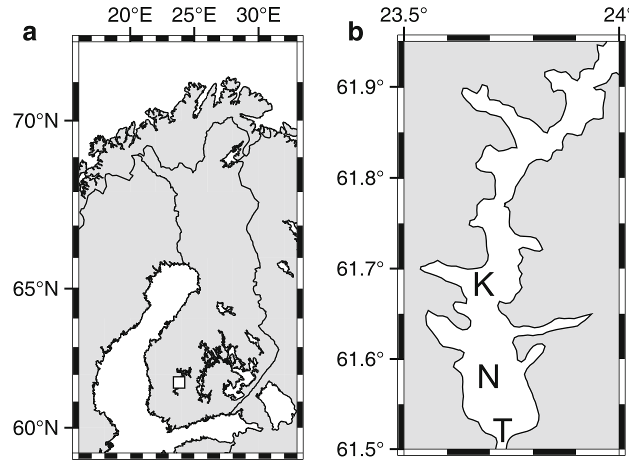

Fig. 1 **a** Finland, *white rectangle* denotes place of lake Näsijärvi (**b**), which contains parts Koljonselkä (K) and Näsiselkä (N) and is located north of Tampere (T)

Measured monthly mean temperatures are from Tampere (Fig. 1), station Härmälä (61.5°N, 23.75°E, 85 m.a.s.l.). The typical error of today's measurements is 0.03°C (Mudelsee 2010, Table 1.3 therein). However, errors from unknown changes in the observational system (radiation shelter, calibration of the instrument, etc.), that is, inhomogeneities, may add considerably to the pure measurement error, especially for earlier periods. Data had therefore been corrected for inhomogeneities, and interpolated for some missing months (Tuomenvirta 2004). The interval covered is from January 1873 to December 2002. Time series of April and November–December temperatures are shown in Fig. 2a, b/c, respectively.

3 Methods and results

3.1 Calibration

The first task is to detect the suitable winter month of the temperature data to be related to breakup date, freeze-up

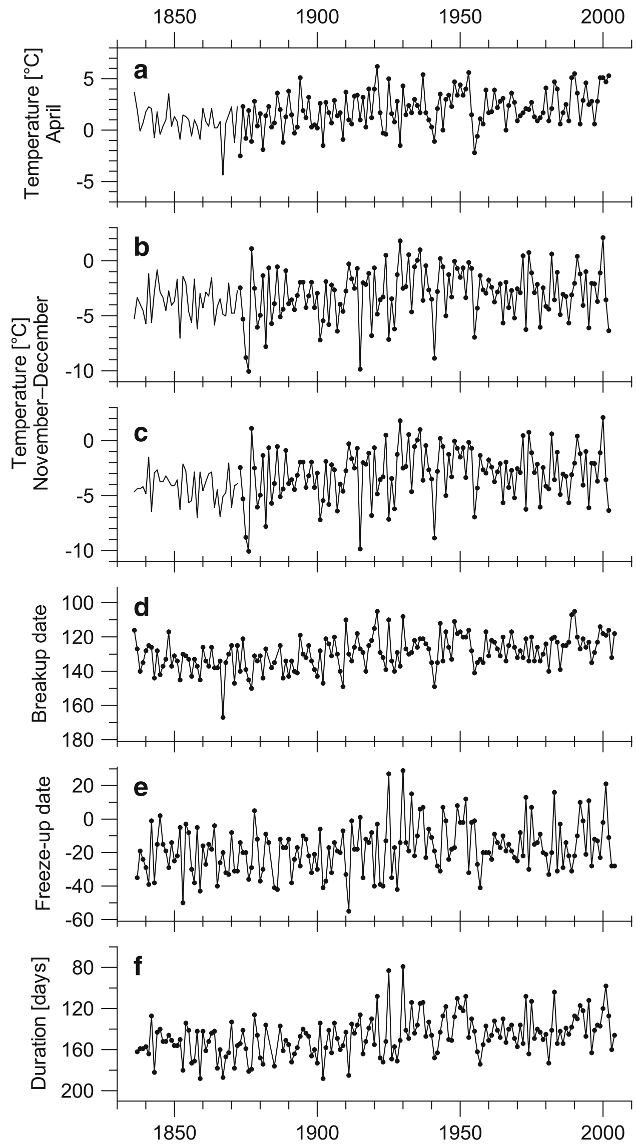

Fig. 2 Time series of air temperature at Tampere (**a–c**) and ice events (freeze-up, breakup and duration of ice cover) in lake Näsijärvi (**d–f**); measured values are shown as *lines* with *dots*; proxy-inferred temperature values for the interval 1836–1872 (**a** via breakup, **b** via freeze-up, **c** via duration) are shown as *lines*; day values (**d, e**) are relative to 1 January of a year; note inverted *y*-axes (**d, f**); data (**a, b**) given with confidence interval as Online Resource 1

Table 1 Correlation results with 95% confidence interval

Temperature in month	Ice event	*r*	
April	Breakup date	−0.82	[−0.86; −0.73]
November–December	Freeze-up date	0.77	[0.67; 0.83]
November–December	Duration	−0.77	[−0.82; −0.68]

date or duration. This requires construction of bivariate time series, $\{t(i), x(i), y(i)\}_{i=1}^{n}$, where *t*(*i*) is year, *x*(*i*) is ice-event date (day relative to 1 January) or duration, *y*(*i*) is temperature averaged over an interval of winter months and *n* is data size. Pearson's correlation coefficient, *r*, between *x*(*i*) and *y*(*i*) measures the degree of the linear relationship between both variables.

If *x*(*i*) is breakup date, then the strongest correlation (largest absolute value) is obtained for *y*(*i*) given by April temperature ($r = -0.82, n = 128$). If *x*(*i*) is freeze-up date and duration, then the strongest correlation is obtained for *y*(*i*) given by November–December temperature ($r = 0.77$, $n = 128$ and $r = -0.77, n = 126$), respectively.

A confidence interval (CI) for the correlation estimate helps to assess whether other selections of months for calculating *y*(*i*) yield significantly lower correlations. Pairwise-moving block bootstrap resampling (Mudelsee 2010, Chapter 7 therein) is a powerful tool for CI construction. By resampling (*x*(*i*), *y*(*i*)) pairs, we do not have to rely on the assumption of normally distributed data. By resampling the pairs over blocks in *t*(*i*), we do not have to assume that the time series are without autocorrelation. CI construction for *r* estimates is implemented in the PearsonT software (Mudelsee 2003).

The resulting *r* values with 95% CI are given in Table 1. Here for the Näsijärvi time series, the blocking procedure turns out as not necessary because the spacing of the *t*(*i*) of 1 year means that ice events and temperatures from the past winter do not influence ice events and temperatures in the current winter, respectively. In case of the breakup date, instead of the April temperature with $r = -0.82$ [−0.86; −0.73], one may also take April–May temperature with a correlation value ($r = -0.77$) that lies within the CI, but should not take May temperature ($r = -0.49$), March temperature ($r = -0.48$) or other temperatures. In case of the freeze-up date, the strongest correlation is for November–December temperature, $r = 0.77$ [0.67; 0.83]. Taking just November temperature would give $r = 0.65$, taking just December would give $r = 0.63$. In case of the duration, it is interesting that the strongest correlation is also for November–December temperature, $r = -0.77$ [−0.82; −0.68], and not for November–April temperature ($r = -0.69$).

Replacing the freeze-up date 31 December by 24 November 1992 yielded correlations (November–December temperature versus freeze-up date, $r = -0.75$; November–December temperature versus duration, $r = -0.76$) indistinguishable from the results in Table 1.

The second task is to calibrate the relation between ice-event date or duration, *X*(*i*), and average temperature in the detected monthly intervals (Table 1), *Y*(*i*), by means of a regression:

$$Y(i) = \beta_0 + \beta_1 [X(i) - X_{\mathrm{noise}}(i)] + Y_{\mathrm{noise}}(i), \quad (1)$$

$i = 1, \ldots, n$ (Following statistical convention, *Y*(*i*) is a random variable while *y*(*i*) is a numerical value, analogously for *X*(*i*)). The parameters β_0 and β_1 can be estimated using ordinary least-squares (OLS) regression. Note that not only temperature, $Y(i)$, has an error component, $Y_{\mathrm{noise}}(i)$, also the ice-event dates or duration, $X(i)$, has such, $X_{\mathrm{noise}}(i)$. This is an errors-in-variables regression model (Draper and Smith 1981), typical for proxy calibration problems. To obtain a bias-free slope estimate, a correction (Mudelsee 2010, Sect. 8.1.1.1 therein) has to be applied to the OLS slope estimate,

$$\widehat{\beta}_1 = \widehat{\beta}_{1,\mathrm{OLS}} / \left\{ 1 - S_X^2 / VAR[x(i)] \right\}. \quad (2)$$

S_X is the standard deviation of the noise component, $VAR[x(i)]$ is the variance of the $x(i)$. The reported S_X value of 1 day for the Näsijärvi dates (and, hence, 1.4 days for the Näsijärvi duration) means that the bias correction factor is close to unity (breakup, 1.011; freeze-up, 1.004; duration, 1.005). For other sites (i.e., other S_X or $VAR[x(i)]$), however, it could be unwise to ignore it.

The resulting calibration parameters with 95% CI (from pairwise-block bootstrap resampling) are given in Table 2, while the resulting fit curves are shown in Fig. 3.

Replacing the freeze-up date 31 December by 24 November 1992 yielded calibrations (November–December temperature versus freeze-up date, $\widehat{\beta}_0 = -1.09$°C, $\widehat{\beta}_1 = 0.116$°C/day; November–December temperature versus duration, $\widehat{\beta}_0 = 9.83$°C, $\widehat{\beta}_1 = -0.089$°C/day) indistinguishable from the results in Table 2.

Dividing the observation time intervals into two halves yielded calibrations (results not shown) indistinguishable from the results in Table 2. This supports the stationarity assumption for the calibrations.

There is little evidence in Fig. 3a for a nonlinear temperature response of ice breakup, which was suggested for Swedish lakes and annual temperature during 1961–1990 (Weyhenmeyer et al. 2004). An F test indicates that there may be larger variability for early freeze-up dates, such as in November, than for late freeze-up dates (Fig. 3b) and for longer durations than for shorter durations (Fig. 3c). Formulating a heteroscedastic regression model would likely not improve calibrations for reconstructed April temperatures, but it could—if a suitable model for the residual variance can be found—improve (i.e., reduce standard errors for) the calibration for November–December temperature. We do not pursue this point further here.

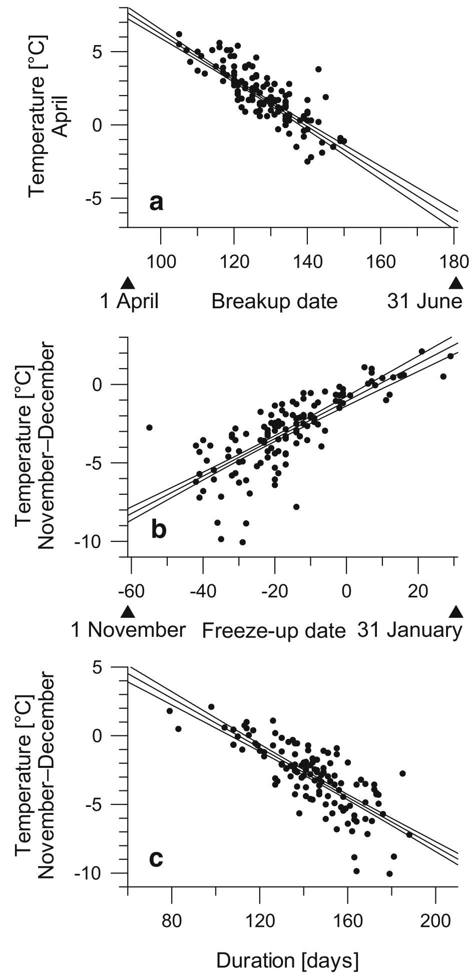

Fig. 3 Calibration of proxy variables by means of linear errors-in-variables regression; best fit with 95% confidence band (*solid lines*), data points (*dots*). **a** April temperature versus breakup date, **b** November–December temperature versus freeze-up date, **c** November–December temperature versus duration of ice cover

Note that the days were taken relative to 1 January and not the vernal equinox. Since the latter date shows a drift of about 0.78 days per 100 years (Sagarin 2001), the day values in Fig. 3a, b are influenced by this noise. A uniform distribution over [0; 0.78] has standard deviation $0.78/3^{1/2} \approx 0.45$ (Johnson et al. 1995, Chapter 26 therein). The drift effect would let S_X increase from 1.0 days to $(1.0 + 0.45^2)^{1/2} \approx 1.1$ days. We ignore this effect on bias correction (Eq. 2).

Table 2 Calibration results (intercept, slope) with 95% confidence interval

Temperature in month	Ice event	$\widehat{\beta}_0$ °C	$\widehat{\beta}_1$ °C/day
April	Breakup date	22.0 [20.6; 23.5]	−0.158 [−0.170; −0.146]
November–December	Freeze-up date	−1.07 [−1.47; −0.67]	0.119 [0.109; 0.130]
November–December	Duration	9.92 [8.95; 10.89]	−0.090 [−0.096; −0.084]

In light of the mentioned error sources, it seems that the slope for the breakup date calibration is significantly larger (in absolute value) than the slopes for freeze-up date or duration calibrations (Table 2).

The calibrations serve for predicting temperature for observed ice event dates or durations. The confidence bands around the linear fits are constructed from pointwise CIs obtained from pairwise-block bootstrap resampling (Mudelsee 2010, Chapter 8 therein). Extra care in interpreting results should be exercised when an observed date or duration falls outside of the calibration interval.

3.2 Trends

OLS fits of a linear regression model (Draper and Smith 1981) to the temperature records over the full interval (1836–1872 via proxy, 1873–2002 measured, $n = 167$) provide a first, simple trend estimation. The model seems suitable from a per-eye inspection (Fig. 4). Trend parameters are listed in Table 3. The linear model has $\nu = n - 2$ degrees of freedom. The reduced sum of squares (Table 3), given by the OLS sum of squares divided by ν, helps to compare the linear fit with other models.

The second trend model is a constrained ramp, fitted by OLS regression. While the unconstrained ramp (Mudelsee 2010, Chapter 4 therein) has four parameters, the constrained version has three ($\nu = n - 3$): a change-point in time (t_1), before which the level is constant (x_1) and after which the change has a constant slope (also called β_1). Bootstrap error bar construction for ramp parameters is implemented in the RAMPFIT software (Mudelsee 2000).

The fitted ramps (Fig. 4) exhibit a reduced sum of squares very close to the values for the linear fits (Table 3). However, it is interesting to note that the estimated change-points $\widehat{t}_1$ lie in the "proxy time interval" (April temperature) or close to it (November–December temperature). The extension by means of the proxy variable, thus, suggests that there may have been a period, namely the early to mid-19th century, with more or less constant April or November–December temperatures in the Tampere region.

April temperatures show smaller variations than November–December temperatures, which is also reflected by smaller estimation error bars. Results from using the duration proxy are similar to those for the freeze-up date proxy and not shown.

April 1867 was a cold event that falls outside the calibration range: proxy-inferred temperature was −4.4°C with 95% CI [−4.9°C; −3.8°C] and break-up date was as late as 17 June. Corroboration for this inferred cold event comes from Stockholm, where March-to-May temperature was measured as −4.8°C (Jones and Bradley 1992).

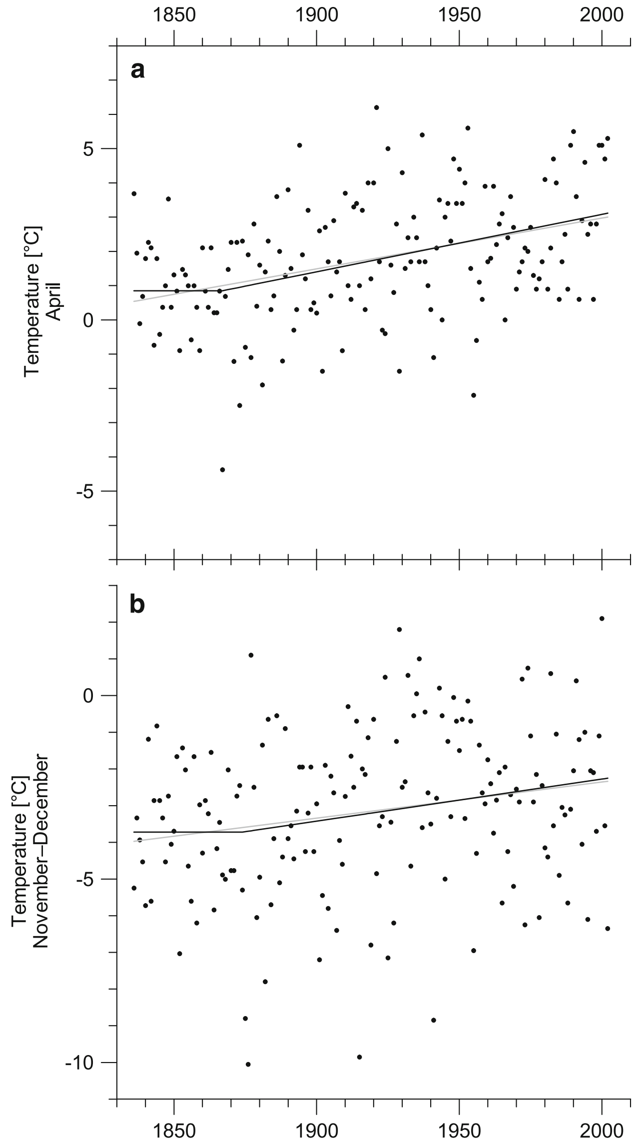

Fig. 4 Trends in temperature, **a** April, **b** November–December (for 1836–1872 via freeze-up date); best linear fits (*grey solid lines*) and best constrained ramp fits (*dark solid lines*)

Table 3 Trend estimation results (slope) with 1-σ bootstrap standard errors; ξ_ν^2, reduced sum of squares; for constrained ramp fits also change-point time, $\widehat{t}_1$, is given; November–December temperature for 1836–1872 inferred via freeze-up date

Temperature in month	Fit method	$\widehat{t}_1$	$\widehat{\beta}_1$°C/(100 a)	ξ_ν^2 (°C)^2
April	Linear		1.48 ± 0.21	2.78
April	Ramp	1867 ± 25	1.67 ± 0.44	2.76
November–December	Linear		0.98 ± 0.28	4.82
November–December	Ramp	1874 ± 45	1.16 ± 0.93	4.84

3.3 Cycles

A spectrum is helpful for distinguishing cyclical forcing mechanisms of the winter temperature in the Tampere region and broad-band resonances. The Lomb–Scargle periodogram combined with the Welch's Overlapped Segment Averaging procedure (Schulz and Stattegger 1997; Mudelsee 2010) is applied to the ramp-detrended series (Fig. 4). A hypothesis test of a red-noise alternative is also implemented in the REDFIT software (Schulz and Mudelsee 2002).

The resulting spectra (Fig. 5) reveal a number of peaks in the range of 2–5 year period. The noise alternative is "only slightly red," in correspondence to the weak autocorrelation mentioned in previous sections. The period values are confirmed by means of applying a second technique, multitaper estimation (Thomson 1982; Mudelsee 2010).

4 Discussion

Two possible sources of inhomogeneity influencing the occurrence of ice events in lake Näsijärvi come from Tampere: district heating and soot (via albedo). A measurable influence may have persisted since the 1990s (Matti Joki, personal communication, 2005). It appears, however,

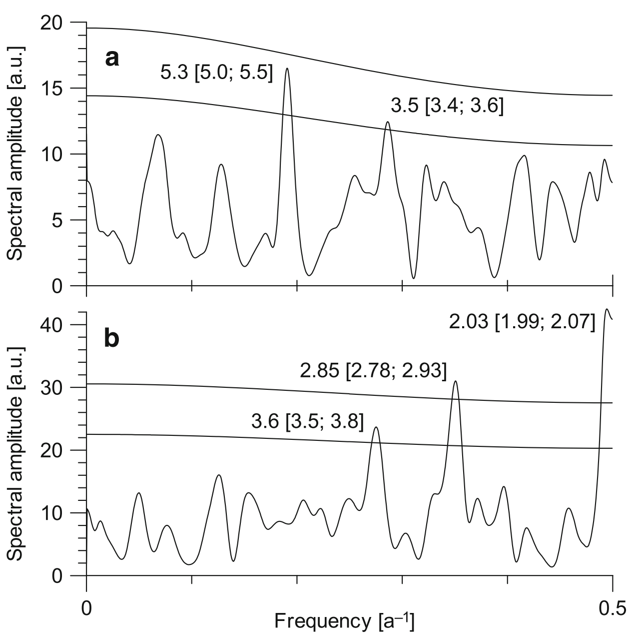

Fig. 5 Spectra of temperature changes, **a** April, **b** November–December (for 1836–1872 via freeze-up date); each panel shows spectral amplitude (*wiggly line*), upper 95% (*lower smooth line*) and upper 99% (*upper smooth line*) chi-squared bounds for AR(1) red-noise alternative; peaks are labelled with period value (in years) and 6-dB bandwidth interval; estimation parameters: 3 segments of 50% overlap, Welch I taper, oversampling factor 20; a.u., arbitrary units

that this had no effect on the temperature calibration parameters (Table 2), as a sensitivity study revealed, in which the upper time interval bound was successively set from 2002 back to 1950 (results not shown).

Ice breakup in Näsijärvi reflects mainly April temperature, while freeze-up and duration reflect mainly November–December temperature. Those months are not predefined by theory but empirically found. The lake system (water, ice and air temperature) is complex, and other lakes than Näsijärvi may show highest susceptibilities for other months.

Inasmuch freeze-up and breakup dates are influenced by other factors than temperature, those may in principle be included in a multivariate regression, yielding possibly a reduced calibration error. A specific factor may be the pattern of atmospheric circulation over the region (Prowse et al. 2007). We think that the spatial resolution and accuracy of state-of-the-art reconstructions of this factor for the past centuries (Luterbacher et al. 2010) do not yet let one expect a large reduction in calibration error.

A calibration requires that the regression is carried out of the response variable Y (temperature) on the predictor variable X (ice event proxy). A regression in the other direction gives, if noise is present, another result (Draper and Smith 1981). This effect, and possibly also ignored predictor noise, likely explains why Korhonen (2006) found a different slope for changes in Näsijärvi breakup date over changes in April temperature (−0.2°C/day) than what Table 2 shows (−0.158°C/day [−0.170°C/day; −0.146°C/day]).

Of particular practical relevance is the calibration slope, β_1, which relates changes in temperature with changes in ice event dates or duration of ice cover. My analysis (Table 2) suggests that the absolute value of β_1 for a single geographical site depends also on ice event type and winter month. Freeze-up date is less sensitive to November–December temperature than breakup date to April temperature. It may not be clear whether the assumption of a constant value of $|\beta_1|$ around 0.2°C/day "for many lakes and rivers around the Northern Hemisphere" (Magnuson et al. 2000, p. 1744 therein) is fulfilled or instead $|\beta_1|$ depends not only on the ice event type but also on the geographical position or the time interval. The following is a list of $|\beta_1|$ values found in the literature on northern Eurasia, which may help assessing the validity of the constant-slope assumption. It should be kept in mind that $|\beta_1|$ determination was not always the focus of the cited papers; the values given therein were usually without error bars or details about the statistical estimation procedure.

Breakup dates were related to April or April–May temperatures for lakes in Finland for the interval from the mid-19th to the end of the twentieth century (Korhonen 2006), with resulting slopes ($|\beta_1|$) between 0.28°C/day and

0.20°C/day; for freeze-up dates related to November or November–December temperature in the same space–time setting, Korhonen (2006) found slopes between 0.39°C/day and 0.19°C/day. As said in a preceding paragraph, the unavailability of error bars as well as the "other regression direction" (predictor temperature) may explain the deviations to the results in Table 2. Duration of ice cover was connected to annual mean temperature for a lake in Finland, mid-19th to end of 20th century, by means of a thermal degree-day modelling approach (Thompson et al. 2005), obtaining a slope of 0.06°C/day. Temperature in February to April versus the predictor breakup date for lake Randsfjord in Norway during 1875–2006 yielded a calibration (Nordli et al. 2007) with a slope of 0.17°C/day (no error bars). Temperature in the winter months December to March (from the region Stockholm–Tallinn–St. Petersburg) was calibrated against breakup date in the port of Tallinn (Gulf of Finland) for a long interval, 1757–1999 (Tarand and Nordli 2001), resulting in a slope of 0.098°C/day. Duration of ice cover in Polish lowland lakes was related to the predictor December–March temperature for the period 1961–2000 (Marszelewski and Skowron 2006), giving $|\beta_1|$ slopes between 0.06°C/day and 0.13°C/day (no error bars). Finally, air temperature at around March in Vladivostok versus breakup date in lake Khanka (about 160 km north) for the short time span 1984–2003 (Nonaka et al. 2007) exhibited a linear calibration with a slope of 0.26°C/day (no error bars).

As regards temperature trends since the 19th century, the inadequacy of the linear model is illustrated by the IPCC–WG I report (Trenberth et al. 2007, FAQ 3.1, Fig. 1 therein): the rate of the temperature increase depends strongly on the lower bound of the estimation interval. The more sophisticated ramp model may help achieving more quantitative information via the estimated change-point time in the 19th century. The observed warming rates of 1.16–1.67°C per century since about 1867–1874 (Fig. 4) seem to be consistent with the general observation (Trenberth et al. 2007) that rates in northern Europe and winter are stronger than in other seasons and at most other places. Note that the constrained ramp, which is the simplest model, makes the assumption of one unique change-point; other change-point models (e.g., Mudelsee 2009) may be used as well.

As regards forcing factors of winter temperature changes in northern Europe, the spectral analysis (Sect. 3.3) suggests that shorter-term influences (2–5 year period) may exist. These factors may include the North Atlantic Oscillation (NAO) (Yoo and D'Odorico 2002). A closer inspection of this association by means of regional climate model studies should therefore be interesting. On the other hand, longer-term cyclical influences (e.g., solar activity) may not be strong. However, due to the relative shortness of the Näsijärvi records of 167 years, this verdict applies only to periods up to about half of the length of the segments (Mudelsee 2010), that is, 42 years (Fig. 5). Testing for the presence of even longer cycles, such as those named after Gleissberg (1965) and de Vries/Suess (de Vries 1958; Münnich et al. 1958; Suess 1965), requires even longer time series.

5 Conclusions

1. Proxy documentary data about ice events (breakup, freeze-up and duration of ice cover) in lake Näsijärvi allowed to extend the measured record of winter temperatures back from 1873 to 1836.
2. Correlation techniques and errors-in-variables regression, both combined with block bootstrap resampling, are powerful tools for calibrating the relation proxy–measured variable for the overlapping period (which is from 1873 to 2002 for Näsijärvi).
3. In the case of Näsijärvi, changes in breakup date have the strongest correlation with changes in temperature during the month of April, while both freeze-up and duration are stronger related to November–December temperature.
4. In the case of Näsijärvi, the calibration slope for the breakup date proxy (0.158°C/day) is larger than for freeze-up date (0.119°C/day) or duration (0.090°C/day). A comparison with results from other proxy records showed that the slope may depend also on the geographical site. This has to be taken into account when reconstructing spatial temperature histories back in time.
5. Trend analyses of the full Näsijärvi temperature records (1836–2002) indicated the existence of minor change-points at around 1867–1874, with warming rates thereafter of 1.16–1.67°C per century.
6. Spectral analyses revealed peaks in the band between 2 and 5 year period, which may point to NAO influences, and less power in the decadal band (up to 42 year period).

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