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Changing seasonality in North Atlantic coastal sea level from the analysis of long tide gauge records

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ABSTRACT

Sea level is a key variable in the context of global climate change. Climate-induced variability is expected to affect not only the mean sea level but also the amplitude and phase of its seasonal cycle. This study addresses the changes in the amplitude and phase of the annual cycle of coastal sea level in the extra-tropical North Atlantic. The physical causes of these variations are explored by analysing the association between fluctuations in the annual amplitude of sea level and in ancillary parameters [atmospheric pressure, sea-surface temperature and North Atlantic Oscillation (NAO) winter index]. The annual cycle is extracted through autoregressive decomposition, in order to be able to separate variations in seasonality from long-term interannual variations in the mean. The changes detected in the annual sea level cycle are regionally coherent, and related to changes in the analysed forcing parameters. At the northern sites, fluctuations in the annual amplitude of sea level are associated with concurrent changes in temperature, while atmospheric pressure is the dominant influence for most of the sites on the western boundary. The state of the NAO influences the annual variability in the Southern Bight, possibly through NAO-related changes in wind stress and ocean circulation.

1. Introduction

Sea level change has been attracting considerable attention in recent years, in part due to sea level rise scenarios emanating from global warming predictions (IPCC, 2001). Much research effort on present sea level change has been concentrated on the identification of trends from the available instrumental records. Climate variability, however, is expected not only to introduce variability in the mean value of atmospheric and oceanographic parameters, but also in the amplitude and phase of the seasonal cycle in these parameters (Thompson, 1995; Plag and Tsimplis, 1999). Changes in the amplitude and phase of the annual cycle of global surface temperature have been detected both in observations and model data, but the cause of such changes, in particular whether they result from variability in atmospheric CO_2 concentration, from an alternative forcing, or from internal variability, remains far from being settled (Thomson, 1995; Thomson, 1996; Karl et al., 1996; Mann and Park, 1996; White et al., 1996; Wallace and Osborne, 2002). Variability in global surface temperature influences atmospheric circulation and a number of surface climate variables, including sea level, a key indicator of climate change and an important observational constraint on climate models.

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Tide gauge records constitute a precious repository of information on 20th century climate variability. The seasonal cycle is an ubiquitous feature in tide gauge records. Sea level seasonality is influenced by several oceanographic and atmospheric processes (Gill and Niller, 1973). These include: atmospheric processes, associated to changes in air pressure and winds, which affect sea level through the inverted barometer response and wind setup; variability in solar radiation, which influences the gain/loss of heat by the upper layer of the ocean through steric density effects; hydrological processes (river runoff, glacier melt and mass exchange in the net evaporative cycle); oceanographic processes such as ocean circulation and wave setup.

Traditionally, a stationary seasonal pattern is assumed in the analysis of sea level variability from tide gauge records. There are two perspectives on the definition of a constant seasonal cycle: an empirical perspective, in which a mean seasonal cycle is computed from monthly or seasonal averages over a number of years, and a parametric approach in which the seasonal cycle is described by a small number of parameters (amplitude, phase) from the sum of harmonics at different frequencies. The mean seasonal cycle of sea level has been investigated by Pattullo et al. (1955) from global tide gauge records. Tsimplis and Woodworth (1994) mapped the global distribution of the seasonal cycle of coastal sea level by fitting to each tide gauge record a linear model of sinusoidal functions at the fundamental frequencies. Garcia-Lafuente et al. (2004) investigated sea level seasonality from Spanish tide gauge records focusing on the contribution of the different forcing factors but again through harmonic analysis and assuming a constant amplitude and phase for the seasonal cycle.

The description of seasonal variability through a constant pattern reflects the belief in a stationary behaviour of the climate system at the seasonal scale. However, seasonality reflects the complex non-linear response of the climate system to regular solar forcing, and there is no reason for the seasonal cycle to be treated as a constant pattern, invariant from year to year (Pezzulli et al., 2005). The extraction of a time-varying seasonal pattern from a time series of observations suffers from the obvious limitation resulting from the lack of a unique and precise definition of seasonality. The seasonal pattern is therefore dependent on the specific approach applied for its derivation and on the corresponding underlying assumptions. In order to separate seasonal variations from longer-term interannual variations in the mean, specific statistical approaches are required. Furthermore, a key aspect in a climate change context is the possibility to produce error bars for the detected seasonal variations and therefore assess whether they are statistically significant.

Moving harmonic analysis and other frequency-domain approaches aiming at describing the temporal evolution of periodicities with known frequency such as complex demodulation, lack adequate inferential procedures. The same applies to non-parametric approaches in the time domain, for which the description of the seasonal patterns depends on pre-set parameters reflecting the trade-off between the quantity of information extracted versus the degree of statistical confidence in that information. While empirical methods for deriving periodic components are constrained to a trade-off between goodness of fit and smoothness, model-based approaches use statistical model selection criteria to select an ‘optimal’ model from the data and uncertainty can be handled within a well-defined statistical framework. The use of model-based approaches in the estimation of time-varying seasonality became feasible in recent years due to the increase in computational power, propelling the development of explicit statistical models for the non-stationary analysis of seasonality, mainly in the time series and econometric literature. Model-based approaches such as constrained second order autoregressions (Newton and North, 1991), structural time series models (Harvey, 1989; Durbin and Koopman, 2001) and the closely related approach based on dynamic linear models (West and Harrison, 1997) have proved effective in the analysis of seasonal patterns within a geophysical context (West, 1995; West, 1996; Schwing and Mendelssohn, 1997; Bograd et al., 2002)

In this study the changing seasonality of coastal sea level in the extra-tropical North Atlanic is examined from the analysis of long and continuous tide gauge records. Changes in the seasonal cycle of sea level have been previously investigated by Ekman and Stigebrandt (1990) and Ekman (1999) through a moving Fourier analysis of the long tide gauge record from Stockholm. These authors found an increase in the amplitude of the dominant annual component associated to changes in winter wind conditions. Plag and Tsimplis (1999), also using moving harmonic analysis, investigated the spatial and temporal variability in the seasonal cycle of sea level from tide gauge records in the North and Baltic Seas, and related the detected variability with changes in the position and extent of the separation zone between maritime and continental regions.

Here, a model-based approach is used to derive time-varying seasonal patterns of sea level variability along with associated uncertainties. The main objective is to extract statistically significant fluctuations in the amplitude and phase of the annual sea level cycle at the coastal sites. Furthermore, beyond documenting variations in the seasonal cycle of coastal sea level in the North Atlantic, the possible physical causes which are responsible for these variations are explored by analysing the association between fluctuations in the annual pattern of sea level and of ancillary parameters (atmospheric pressure, sea-surface temperature and the North Atlantic Oscillation index).

The analysed tide gauge and climate data are described in Section 2. The autoregressive decomposition approach applied in the estimation of time-varying seasonal patterns is addressed in Section 3. The temporal variability of the annual cycle of sea level and forcing parameters is described in Section 4. The relation between changes in the seasonal pattern of sea level and changes in forcing parameters is discussed in Section 5, and concluding remarks are given in Section 6.

2. Data

Monthly tide gauge records adjusted to a Revised Local Reference (RLR) datum defined relative to the tide gauge benchmark are obtained from the Permanent Service for Mean Sea Level (PSMSL) database (Woodworth and Player, 2003). Long and continuous sea level records are required for the reliable estimation of changes in seasonality. Tide gauge stations in the North Atlantic with long ($>$50 yr) and continuous records (gaps $<$1 yr, missing values $<$2.5%) are considered (Fig. 1, Table 1). Although longer time series are available for some of the records (e.g. Brest, Halifax), shorter periods have been selected for analysis in order to avoid gaps in the time series and maintain coherency with criteria for missing observations; in the case of New York, values prior to 1927 are not considered since they are interpolated from half-tide levels.

Ancillary data are further considered. Monthly time series of SST from 1854 to 2000 are extracted from the $2^\circ \times 2^\circ$ grid extended reconstructed SST dataset of Smith and Reynolds (2004) at the gridpoints nearest to each tide gauge location. The same sea-surface temperature (SST) gridpoint is considered for the stations at Chesapeake Bay (Baltimore, Kiptopeke and Hampton). Monthly time series of sea level pressure (SLP) are obtained from the corrected dataset of gridded SLP data from 1900 until present (Trenberth and Paolino, 1980) and collocated to the tide gauge

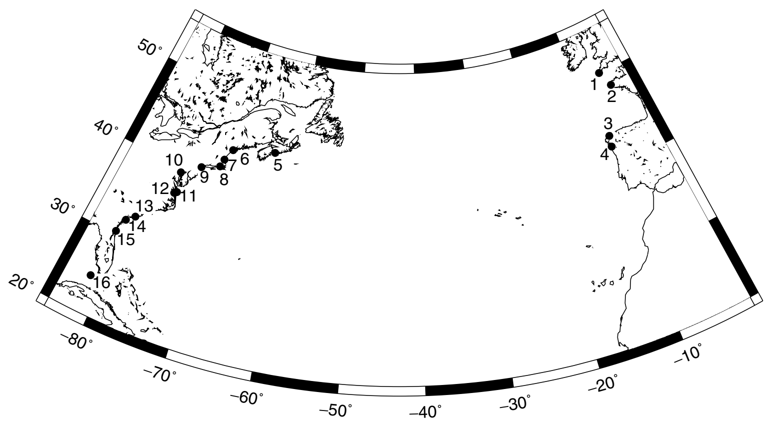

Fig. 1. Study area with tide gauge locations, 1: Newlyn; 2: Brest; 3: Coruña; 4: Vigo; 5: Halifax; 6: Portland; 7: Boston; 8: Newport; 9: New York; 10: Baltimore; 11: Kiptopeke; 12: Hampton; 13: Charleston; 14: Fort Pulaski; 15: Mayport and 16: Key West.

Table 1. Analysed tide gauge records: λ denotes longitude, φ latitude and *N* the total length (number of observations) of the record

Station	λ (°E)	φ (°N)	Period	*N*	%missing
North Atlantic Eastern Boundary					
Newlyn	−05.55	50.10	1916–2003	1056	0.19
Brest	−04.50	48.38	1953–2000	576	0.30
Coruña	−08.40	43.37	1944–2001	696	1.60
Vigo	−08.73	42.23	1944–2001	696	0.86
Scotian Shelf and Gulf of Maine					
Halifax	−63.58	44.67	1920–2002	996	1.30
Portland	−70.25	43.67	1912–2003	1104	0.27
Boston	−71.05	42.35	1921–2003	996	0.80
Mid-Atlantic Bight					
Newport	−71.33	41.50	1931–2003	876	1.26
New York	−74.02	40.70	1927–2003	924	1.08
Baltimore	−76.58	39.37	1903–2003	1212	0.16
Kiptopeke	−75.98	37.17	1952–2003	624	0.80
Hampton	−76.33	36.95	1928–2003	912	0.00
Southern Bight					
Charleston	−79.93	32.78	1922–2003	984	0.00
Fort Pulaski	−80.90	32.03	1935–2003	828	1.21
Mayport	−81.43	30.40	1929–2000	864	0.35
Florida Strait					
Key West	−81.80	24.55	1913–2003	1092	0.73

locations. The yearly winter index of Jones et al. (1997) is used as a representation of the state of the North Atlantic Oscillation (NAO).

3. Methodology

In this work the seasonal cycle of sea level is estimated within a model-based perspective by autoregressive decomposition. This method was developed by West (1997) and its application to the analysis of palaeoclimatic time series is considered in West (1995) and West (1996). Here, the approach is briefly outlined.

An autoregressive model expresses the present value of a time series $\{X_t : t = 1, \ldots, N\}$ as a linear combination of past values plus a random stochastic term:

$$X_t = \sum_{j=1}^{p} \phi_j X_{t-j} + \varepsilon_t, \tag{1}$$

where $p \in \mathbb{N}$ is the order of the process, $\varepsilon_t \sim N(0, \sigma^2)$ is a Gaussian white noise and ϕ_j are the autoregressive parameters.

The autoregressive decomposition method is based on the fact that an autoregressive process of order p, $AR(p)$, can be written as

$$X_t = \sum_{j=1}^{p} \gamma_t^j, \tag{2}$$

where each component $\gamma_t^j, j = 1, \ldots, p$ is either a first order, $AR(1)$, or a second order $AR(2)$ autoregressive process (West, 1997). Physically, this means that if a system is described by an autoregressive process it can be interpreted as a combination of pure (non-periodic) relaxators and damped (periodic) oscillators. The autoregressive approach has thus a strong physical motivation, being very close to the first order linear treatment for the dynamics of a system in classical physics.

Furthermore, the components γ_t can be computed directly from the eigenvectors of the $p \times p$ matrix of autoregressive parameters G,

$$G = \begin{bmatrix} \phi_1 & \phi_2 & \cdots & \phi_p \\ 1 & 0 & & 0 \\ 0 & 1 & & \vdots \\ \vdots & & \ddots & \\ 0 & 0 & \cdots & 0 \end{bmatrix} \tag{3}$$

and the type of the jth component, γ_t^j, is determined only by the jth eigenvalue of G, $r_j e^{\pm i w_j}$: real eigenvalues ($w = 0$) yield first order, $AR(1)$ components while complex eigenvalues ($w \neq 0$) yield second order, $AR(2)$ components (see West, 1997, for mathematical derivations and proofs).

From the theory of linear difference equations, first order autoregressions $X_t = \phi_1 X_{t-1} + \varepsilon_t$, associated with real eigenvalues

r, correspond to solutions of the form r^t and thus a stationary $AR(1)$ process ($r < 1$) exhibits the behaviour of a pure relaxator. Second order autoregressions $X_t = \phi_1 X_{t-1} + \phi_2 X_{t-2} + \varepsilon_t$, associated with complex eigenvalues $re^{\pm iw}$, correspond to solutions of the form $r^t \cos(wt)$ and thus the process exhibits the behaviour of a damped sinusoid with period $2\pi/w$ and damping factor r.

The estimation of a time-varying seasonal pattern with a given period T by autoregressive decomposition therefore involves the following steps: (i) fitting of a $AR(p)$ model to the time series of observations, that is, estimation of the autoregressive parameters $\phi_1, \phi_2, \ldots, \phi_p$; (ii) construction of the matrix G of autoregressive parameters and computation of the eigenvalues and eigenvectors of G; iii) selection of the complex conjugate eigenvalues with $w \simeq 2\pi/T$; and iv) computation of γ_t from the time series of observations and from the eigenvectors associated with the eigenvalues $re^{\pm iw}$.

The main issue in this scheme is the specification of the order p of the autoregressive process, that is, the number of autoregressive parameters. The Akaike Information Criterion (AIC) can be used to select the number p of autoregressive parameters by comparing the AIC values from models with different orders and selecting the model for which the AIC is minimum.

Since the autoregressive decomposition approach is determined by the autoregressive parameters $\phi_j (j = 1, \ldots, p)$ in (3), inference on the resulting components is derived from inference on the autoregressive parameters, that is, the uncertainty on the autoregressive components is directly related to the uncertainty on the autoregressive parameters.

After specifying the model order p (through AIC) the parameters of the autoregressive process (1) can be estimated by linear regression (ordinary least squares). Therefore, uncertainty can be handled within the usual linear regression framework. The linear model can be considered from a classical perspective, yielding a single estimate for each autoregressive parameter and the associated error, or from a Bayesian perspective, in which instead of estimating a single parameter a whole statistical distribution is derived for the parameter of interest.

Assuming a Bayesian linear regression model with a non-informative prior, the vector of the autoregressive parameters $[\phi_1 \phi_2 \ldots \phi_p]$ has a $p-$multivariate normal distribution with mean given by the least squares estimate of the autoregressive parameters and variance derived from the residuals of the least squares fit. Knowing its statistical distribution, the uncertainty on the autoregressive parameters can be handled by taking a sample of n independent draws from that distribution. Then, instead of a single vector of autoregressive parameters a set of n vectors is obtained. The final autoregressive component γ_t is derived by taking the mean or, for a more robust estimate, the median, of the n autoregressive estimates $\gamma_t^{(1)} \ldots \gamma_t^{(n)}$ obtained from each set of autoregressive parameters.

This method has been implemented in R (R Development Core Team, 2006) as R-package ArDec, released under GPL (General Public License) Version 2 and available from the CRAN repository (http://cran.r-project.org/src/contrib/Descriptions/ArDec.html)

4. Results

4.1. Analysis of tide gauge records

The annual cycle of coastal sea level is analysed from the tide gauge records described in Section 2 using the autoregressive decomposition method summarized in Section 3 and implemented in the R-package ArDec.

For each tide gauge record autoregressive models of different orders are fitted and the optimal model order p is selected through AIC. The autoregressive model of the selected order is cast in a linear regression form and the corresponding autoregressive parameters are estimated by least squares. In order to account for uncertainties in the estimated autoregressive parameters, a sample of $n = 500$ independent draws from the corresponding multivariate normal distribution is considered, yielding a sample of n autoregressive annual components. Histograms for the period ($T = 2\pi/w$) of these n simulated annual components are displayed in Fig. 2. As expected for the annual and persistent component, the periods are close to 12 months and the damping factors (not shown) are very close to 1 for all records. However, the spread in the distributions of periods is distinct for the different records. Uncertainty increases for the records which exhibit a weak annual signal (Boston, Portland) and decreases for the records with a strong annual signal. Baltimore has a strong and well defined annual pattern and thus the uncertainty on the derived component is very low. For Portland and Boston, exhibiting a weak annual cycle, uncertainty is larger. In the case of Vigo and Coruña the estimate obtained for the annual cycle must be viewed with caution since it comprises a very high degree of uncertainty. For these records the damping factors are low and the spread in the distributions of both the period and the damping factor is large.

The final annual component for each record is taken as the median of the $n = 500$ autoregressive components. The variability of the annual cycle is summarized by representing the annual time series, standardized through division by the corresponding standard deviation (σ), within 3 classes of variability (Fig. 3), corresponding to variations relative to the mean below 1.5σ, between $\pm 1.5\sigma$ and above 1.5σ, respectively. Figure 3 displays a complex pattern of seasonal changes taking place over the 20th century in the North Atlantic. Large variations in the annual sea level cycle are mostly concentrated on specific temporal clusters, alternating with ‘quiet’ years for which the seasonal values are closer to the mean value. Such an alternating pattern is clearly visible, for example, at Newlyn and Brest, in the quiet period of the early 1970s, followed by a period of higher annual variability in the late 1970s and again a quiet period during the 1980s. Figure 3 shows coherent patterns of annual changes for stations in the same region, for example a good agreement between the

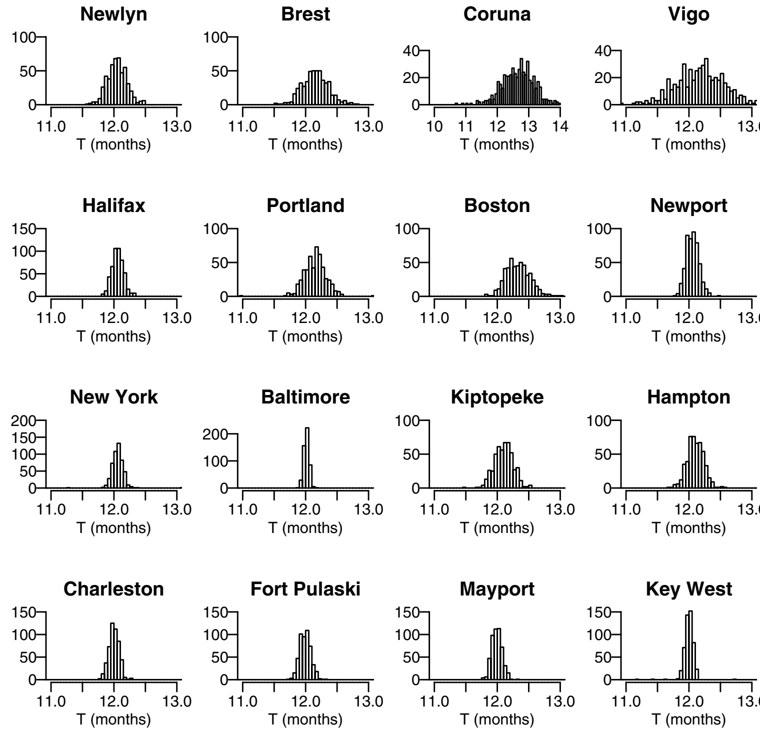

Fig. 2. Histograms for the periods ($T = 2\pi/w$) from the sample of $n = 500$ autoregressive annual components.

patterns on the western boundary and between the Southern Bight stations. The variations are spatially coherent, indicating consistency of the results at the regional level. However, there are considerable differences over the North Atlantic, for example the different variability patterns in the early 1980s, reflecting the distinct oceanographic areas of the coastal sites.

The amplitude of the annual cycle is computed for each tide gauge record as the difference between the corresponding maximum and minimum values for each year. The temporal variability of the amplitude of the annual cycle over the analysed period is shown in Fig. 4, along with the corresponding linear trend if statistically significant at 95% confidence level. The amplitude of the annual cycle exhibits an increasing trend at Newlyn (0.38 ± 0.11 mm yr^{-1}) and Vigo (1.15 ± 0.32 mm yr^{-1}) and a decreasing trend at Baltimore (-0.29 ± 0.12 mm yr^{-1}) and Mayport (-0.64 ± 0.26 mm yr^{-1}).

The description of changes in the annual pattern in terms of changes in amplitude only reflects changes in the maxima/minima of the annual sea level. A more detailed perspective on the variability of the annual cycle of sea level is obtained by looking at the overall annual pattern. The variability of the annual component on each individual season is represented by a monthplot of the annual cycle for each tide gauge station (Fig. 5). The horizontal lines represent the mean value of the annual component for each month over the period spanned by the record. The interannual variability of the annual component is displayed by superimposing to each monthly average the values of the annual component for each individual year (grey lines). Temporal variability in the seasonal pattern is assessed by fitting a linear regression model to each annual subseries (the time series of all January values, of all February values, . . .). The corresponding linear trends are displayed in the monthplot (solid line, black) if the estimate is statistically significant at 95% significance level (Table 2).

The monthplots of the annual component show changes in the annual pattern of the tide gauges on the eastern boundary of the Atlantic, at Chesapeake Bay and at Mayport. For Newlyn the annual subseries for the autumn months (September, October and November) exhibit small but statistically significant increasing trends, indicating that the characteristics of the annual cycle of sea level have changed over the analysed period, from lower sea level values in autumn in the early part of the record to higher autumn values in the latter part. Thus the maximum values of the annual cycle are increasing, contributing, along with the slight decrease in April, to an increase of the overall amplitude of the annual cycle. Brest and Coruña exhibit significant trends for individual months (Table 2), although the linear variability of the amplitude of the annual sea level cycle is not statistically significant for the two records. For Brest the values of the annual cycle are increasing in the late summer (July and August) and decreasing in March, while for Coruña the annual pattern is increasing from July to October. These changes affect the shape of the annual pattern but have a compensating effect in terms of amplitude. For Vigo the maximum values of the annual cycle, in December and January, exhibit a considerable increase over the analysed period, leading to an increase in the overall annual

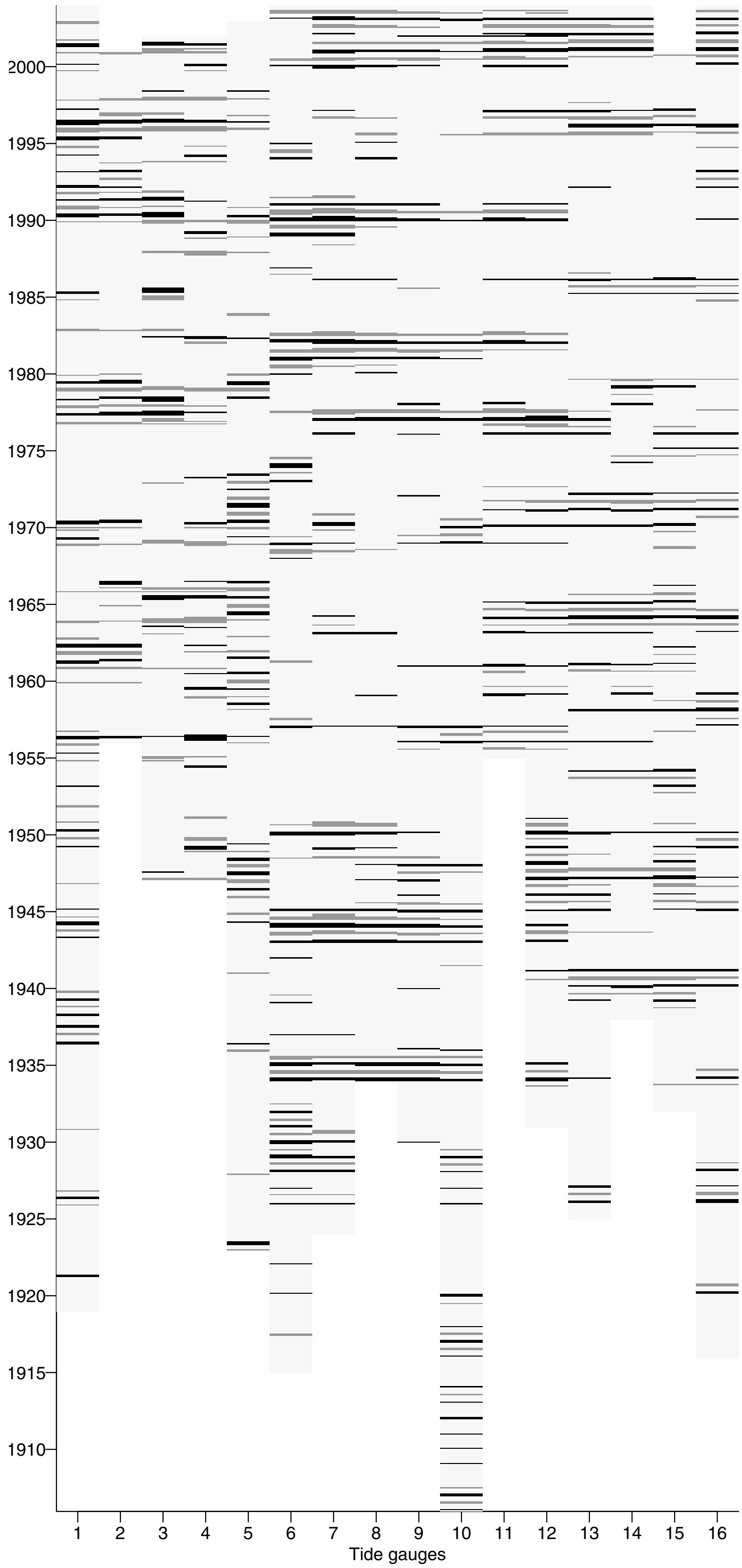

Fig. 3. Classes of annual sea level variability: $\leq -1.5\sigma$ (black), $]-1.5, 1.5[\sigma$ (light grey) and $\geq 1.5\sigma$ (grey). The symbol σ denotes standard deviation.

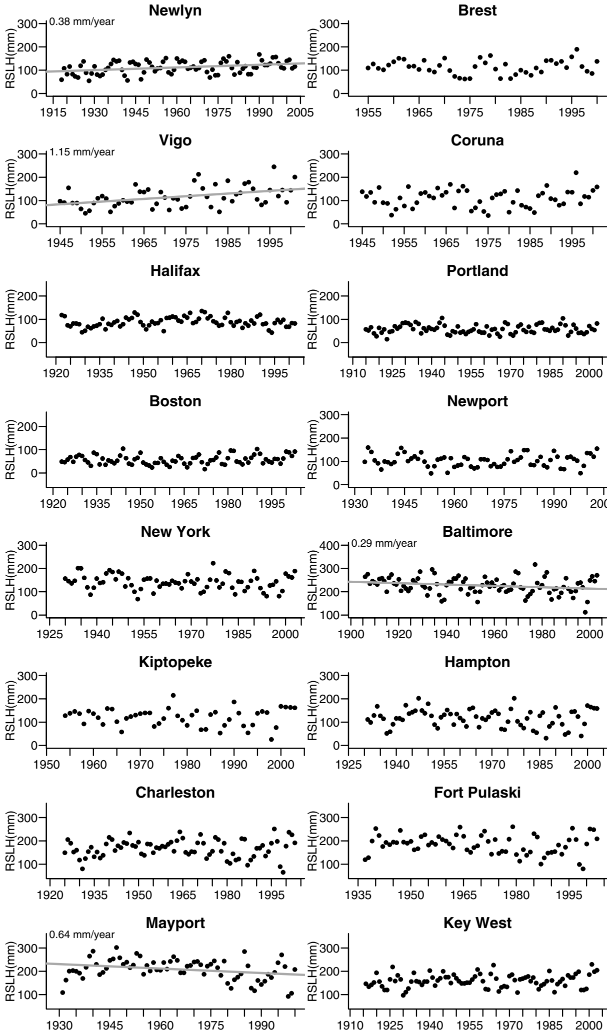

Fig. 4. Amplitude of the annual cycle for each tide gauge record.

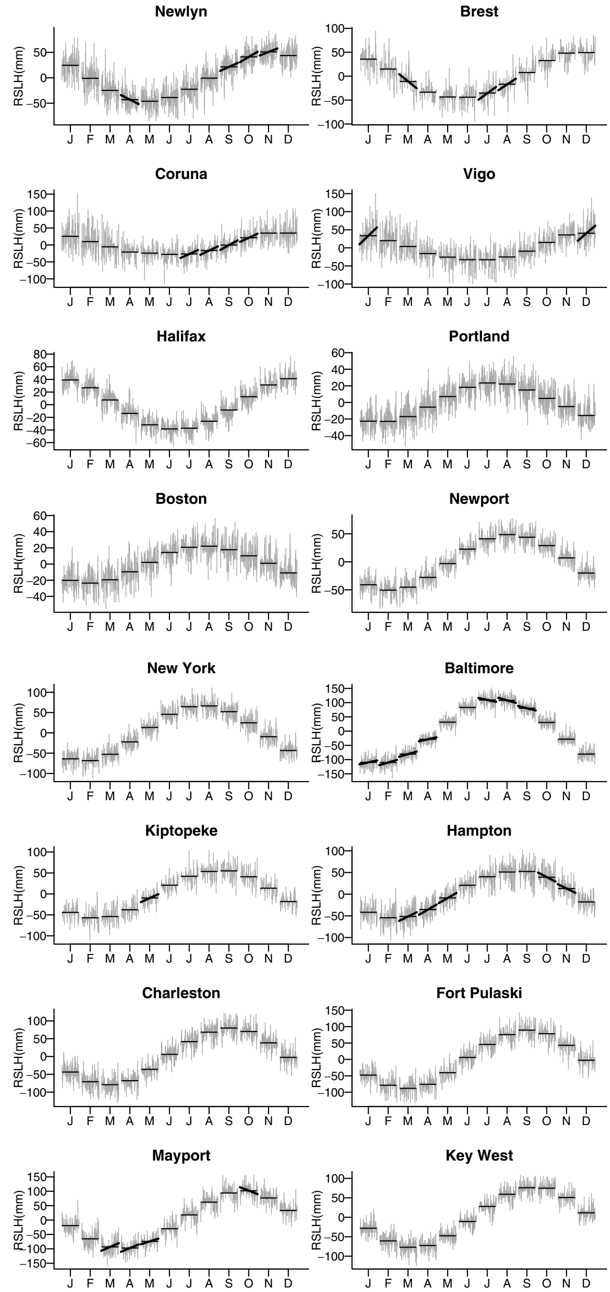

Fig. 5. Plot of the annual component for each tide gauge record. Linear trends for each month are represented if statistically significant at 95% confidence level.

amplitude. For the stations on the Chesapeake Bay (Baltimore, Kiptopeke and Hampton), only Baltimore exhibits a significant (decreasing) linear trend in the amplitude of the annual component, resulting from small but consistent negative slopes in the maximum values, from July to September, and positive slopes in the minimum, from January to April. Mayport also exhibits a decreasing trend in the amplitude of the annual pattern, which is associated with an increase in the minimum values in spring and a decrease in the maximum in October.

The temporal evolution of the phase of the annual sea level cycle is represented in Fig. 6 through the time (months) of maximum and minimum annual values. At Newlyn and Brest, although the minima of the annual cycle occur always in the late spring and the maxima in the early winter there is considerable variability from year to year. Despite the differing lengths, the two records show similar features, for example, in 1980 the later maximum (in January) followed by an unusual early minimum (in February) in 1981. During the 1990s, maxima and minima tend to occur earlier at Newlyn (in October/November rather than November/December, and in March/April rather than April/May), in agreement with the changes in amplitude documented in Fig. 5 and Table 2. At Brest the minima occur later in the first half of the record (May/June) and earlier (March/April) during the 1980s and 1990s. At Halifax, the annual cycle occurs

Table 2. Linear trends from the monthly subseries of the annual cycle of sea level

		Linear trend ($mm\,yr^{-1}$)	Standard error ($mm\,yr^{-1}$)
	April	−0.20	0.08
Newlyn	September	0.18	0.09
	October	0.24	0.07
	November	0.16	0.07
	March	−0.63	0.24
Brest	July	0.57	0.24
	August	0.52	0.23
	July	0.41	0.20
Coruña	August	0.44	0.18
	September	0.50	0.18
	October	0.42	0.20
Vigo	December	0.66	0.27
	January	0.82	0.31
	January	0.14	0.06
	February	0.18	0.07
	March	0.18	0.06
Baltimore	April	0.15	0.05
	July	−0.12	0.06
	August	−0.16	0.06
	September	−0.14	0.06
Kiptopeke	May	0.37	0.16
	March	0.26	0.12
	April	0.31	0.11
Hampton	May	0.32	0.10
	October	−0.26	0.12
	November	−0.25	0.11
	March	0.37	0.16
Mayport	April	0.36	0.13
	May	0.27	0.13
	October	−0.34	0.14

1 month earlier after the 1990s, with most minima occurring in June rather than in July and maxima occurring in December rather than in January. Portland and Boston exhibit similar phase patterns, in particular the unusually late maxima of 1970 and 1971 (in December and November) and the unusually early maximum in 1998, occurring in the winter rather than in the summer. The winter maximum of February 1998 is also visible in the data from Kiptopeke and Hampton.

4.2. *Analysis of climate variables*

In order to gain additional insight on the mechanisms involved in the detected variations in sea level seasonality, a similar analysis is carried out for SST and SLP data. Atmospheric pressure exerts a downward force on the sea surface and influences directly sea level through the inverted barometer response of the ocean: a decrease (increase) in atmospheric pressure of 1 mb raises (depresses) sea level by 1 cm. Furthermore, pressure influences indirectly sea level through wind effects, since winds are correlated with pressure (e.g. Pugh, 2004). Steric variations in sea level result from the thermal expansion/contraction of the water column due to variations in temperature. At the seasonal scale, temperature variations reflect the periodic changes in solar radiance which influence the heat content in the upper layer of the ocean. Therefore, although sea level changes due to thermal expansion represent variations over the entire water column, while changes in SST only reflect variations near the ocean surface, SST is used as a proxy for steric sea level variations.

The annual cycle of SST and SLP is computed by autoregressive decomposition, in the same way as for the annual cycle of sea level. The temporal evolution of the amplitude of the annual cycle of SST is displayed in Fig. 7. Both SST and SLP exhibit appreciable variability in the amplitude of the annual pattern over the analysed period. However, while in the case of SST the variations in the annual amplitude at all sites except Coruña can be described by a linear trend, in the case of SLP only at Kiptopeke and Key West the changes in amplitude correspond to a statistically significant linear trend (0.2 mb/decade and −0.06 mb/decade, respectively). At all locations except for Vigo and Coruña, in the northern Spanish coast, the amplitude of the annual cycle of SST exhibits an increasing trend. On the western boundary the amplitude increases as 0.05–0.06 °C per decade, and almost as twice for New York and Chesapeake Bay.

The correlation between the amplitudes of the annual cycle of sea level, SST and SLP and also between the annual sea level amplitude and the winter NAO index for the corresponding year is described through the (Pearson) correlation coefficient (Table 3). The NAO is one of the most important modes of climate variability in the North Atlantic, being associated with changes in the strength and path of storm systems crossing the Atlantic, changes in sea-ice cover, sea-surface temperature, wave heights and to temperature and precipitation anomalies over Europe, mainly during the winter season (e.g. Marshall et al., 2001; Hurrell et al., 2003). The state of the NAO also influences winter sea level variability (Wakelin et al., 2003; Woolf et al., 2003; Yan et al., 2004; Jevrejeva et al., 2005). Therefore, it is reasonable to investigate the association between variations in the amplitude of the annual cycle of sea level and the state of the NAO, as described by the winter NAO index, despite the fact that an index based on two points at fixed locations can only yield a limited representation of atmospheric circulation variability over the North Atlantic.

The correlation between sea level and SST is positive at Newlyn, Brest and Halifax, and non-significant for the remaining sites. Variations in the annual amplitude of sea level are negatively correlated with changes in the annual amplitude of SLP on the western boundary, from Boston to Fort Pulaski. In the Southern Bight, the changes in the annual amplitude are negatively correlated with the NAO index. A statistically significant negative correlation with the NAO index is also found at Vigo

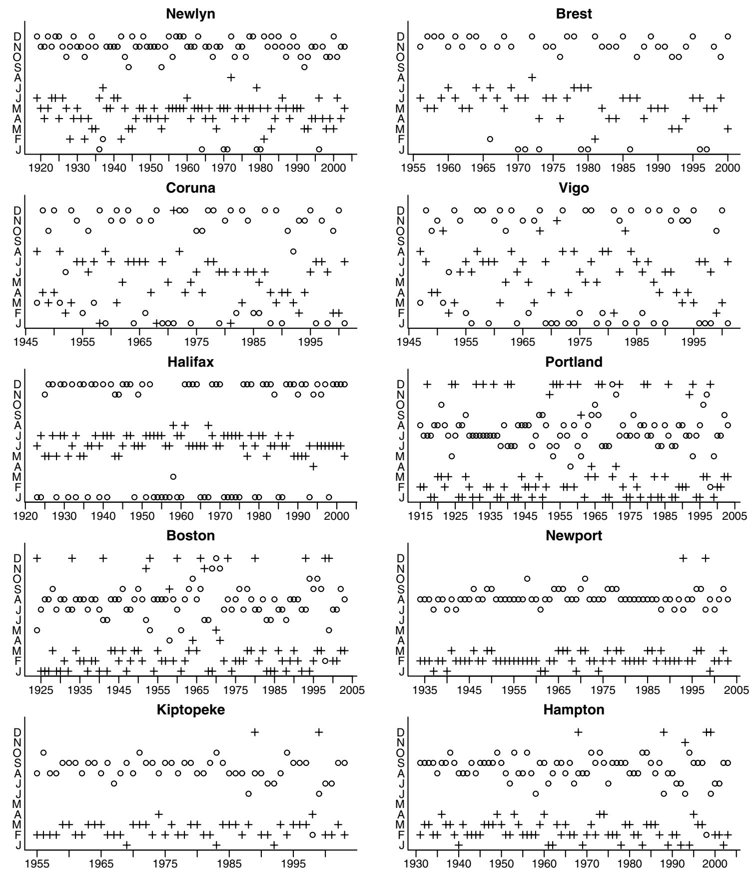

Fig. 6. Times (month) of maximum (circle) and minimum (cross) annual amplitudes.

while at Portland variations in the annual amplitude exhibit a low but significant positive correlation with the NAO index.

The correlation values obtained, although significant from a statistical point of view, are not high. This is not surprising, since it is examined the association between sea level values from tide gauge stations, influenced by very local atmospheric and oceanographic conditions, and gridded climate variables, representing more regional conditions. The use of station measurements rather than gridded values would be preferable, but is hindered by the lack of available data. Despite this obvious limitation, the fact that most variations in the annual pattern are spatially coherent (as can be seen, for example, in Fig. 3) suggests that regional variability can play at least a partial role in the variability of the annual cycle of sea level. Furthermore, the correlation coefficient has been computed from the amplitudes for each year, thereby reducing the effect of autocorrelation in the estimate of the correlation value, giving more confidence on the obtained low but statistically significant correlations.

5. Discussion

The annual cycle of coastal sea level exhibits considerable spatial variability. Baltimore, well inside the Chesapeake Bay, exhibits the largest annual cycle, in agreement with Tsimplis and Woodworth (1994) that found larger annual amplitudes for stations inside Chesapeake Bay than for neighbouring stations on the open coast. Chesapeake Bay is particularly influenced by hydrologic effects and the seasonal pattern at Baltimore is enhanced by the seasonal cycle of run-off. The amplitude of the annual cycle is higher on the western than on the eastern boundary, decreasing northwards.

For all records the amplitude of the annual cycle is not constant over the analysed period, exhibiting substantial interannual variability. The amplitude pattern is adequately described by a negative linear trend for Baltimore and Mayport and by a positive linear trend at Newlyn and Vigo. For Baltimore and Mayport the decreasing trend in the annual amplitude is due to a decrease in the maximum levels, occurring in summer/early autumn, and an increase in the minimum levels, occurring in winter/early spring. For Newlyn, the increase in amplitude is associated to an increase in sea level values in autumn and a decrease in spring, while for Vigo it is only due to an increase in winter sea level values. For the remaining stations, the variability from year-to-year cannot be summarized by a linear trend, either due to the short-period spanned by the record (as seems to be the case for Brest, which has a very similar pattern but a much shorter length than the nearby Newlyn record) or due to the non-monotone nature of the annual amplitude pattern. Since linear trends are very sensitive

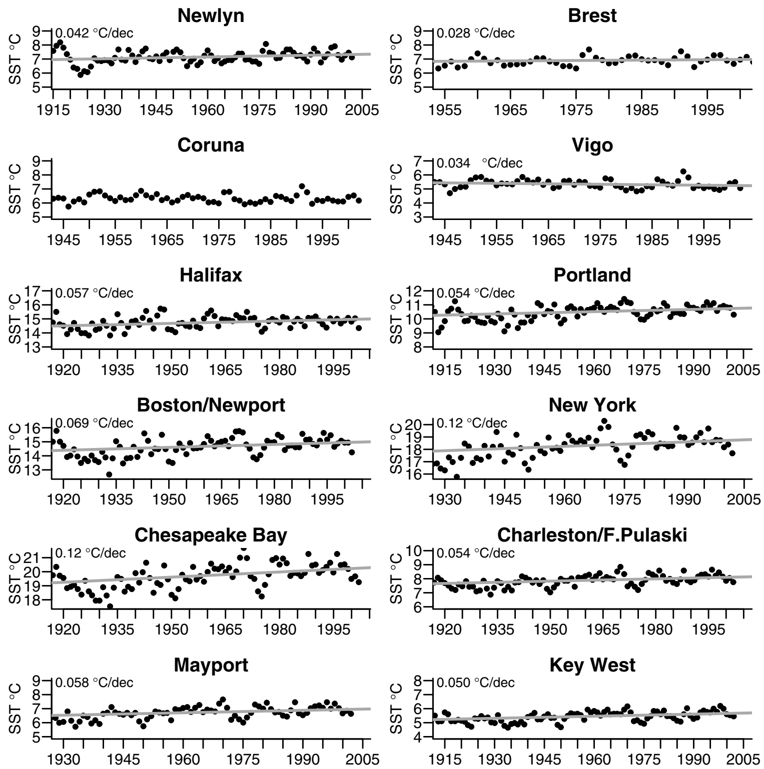

Fig. 7. Amplitude of the annual cycle of sea-surface temperature (SST). The same time series is considered for the stations in Chesapeake Bay (Baltimore, Kiptopeke and Hampton).

Table 3. Pearson correlation coefficient between annual amplitudes of sea level at tide gauge sites and SST, SLP and NAO winter index. Statistically significant correlations at 95% confidence level are displayed in bold. For the sites on the eastern boundary the annual cycle of SLP is below noise level

	$r_{RSLH-SST}$	$r_{RSLH-SLP}$	$r_{RSLH-NAO}$
Newlyn	**0.27**		−0.041
Brest	**0.44**		−0.26
Coruña	0.042		−0.24
Vigo	−0.058		**−0.32**
Halifax	**0.47**	0.20	−0.20
Portland	−0.056	−0.13	**0.21**
Boston	0.14	**−0.27**	0.17
Newport	0.081	**−0.32**	0.02
New York	0.030	**−0.34**	−0.14
Baltimore	0.091	**−0.29**	−0.14
Kiptopeke	−0.079	**−0.45**	−0.23
Hampton	−0.059	**−0.38**	−0.20
Charleston	0.048	**−0.32**	**−0.33**
Fort Pulaski	0.060	**−0.32**	**−0.47**
Mayport	−0.19	−0.10	**−0.32**
Key West	0.057	0.082	−0.087

to small changes at the beginning and/or end of the time series, it must be noted that the representation of variability by a linear trend (for example, in summarizing the changes over the years for each individual month as was carried out in the monthplots), is always highly constrained by the length of the time series and the analysed period.

Concerning the phase of the annual cycle, the temporal pattern of maxima and minima is considerably stabler for the records with a well defined annual cycle and more variable at the sites for which the annual cycle is weak (Coruña, Vigo, Portland and Boston). Although this can be due, in part, to difficulties in the estimation of the annual component, particularly at Coruña and Vigo (as suggested by Fig. 2), the similarity in the pattern of maxima and minima for Coruña and Vigo, and also for Portland and Boston, suggests that the dispersion results from real variability in these records with a week annual cycle rather than from an artefact due to the lower signal-to-noise ratio. At Newlyn, Brest and Halifax, the annual cycle exhibits a 1 month shift, with maxima and minima occurring about a month earlier since the mid-1980s and particularly during the 1990s.

The correlation between the amplitudes of the annual cycle of sea level and SST is positive at Newlyn, Brest and Halifax, suggesting a steric influence, through thermal expansion/contraction, on the variability of the annual amplitude of sea level. However, although at the three sites the SST amplitudes exhibit an increasing linear trend, sea level amplitudes only show a statistically significant trend at Newlyn. Therefore, an apparent contradiction results, where fluctuations in sea level amplitudes are related to concurrent variations in SST amplitudes, but trends in the SST amplitudes do not translate into trends in sea level amplitudes. This fact can be explained by taking into account that linear trends are a limited description of temporal variability and, more importantly, that there is rarely a single factor, such as temperature or any other, influencing the annual cycle of coastal sea level. This can be seen by noting that although the annual cycle of SST is very similar at Newlyn, Brest and Halifax, reflecting the maximum of solar radiance in the summer, the annual cycle of sea level is distinct at Halifax, with minima in summer and maxima in the winter, rather than minima in spring and maxima in the autumn, as for Brest and Newlyn. While sea level at Brest and Newlyn is more strongly affected by the annual cycle of temperature, reflected in the autumn maximum, Halifax is also influenced by atmospheric pressure (Hilmi et al., 2002), and as a consequence the expected maximum in sea level in the early autumn due to thermal expansion is attenuated by the opposing hydrostatic effect due to the maximum in the annual cycle of atmospheric pressure occurring in the early autumn. Although the annual cycle at Halifax is only partially the result of SST variations, with atmospheric pressure also playing an important role, variations in the amplitude of the annual cycle are associated with variations in the annual cycle of SST. This is indicated not only by the statistically significant correlation value, but can also be observed in the similarities in the amplitude patterns of sea level and SST at Halifax (see Figs. 4 and 7, e.g. the common period of increasing amplitudes from the early 1930s until 1950, the common period of decreasing amplitudes in the early 1970s, or the common period of decreasing amplitudes in the early 1990s).

The increasing trend in the amplitude of the annual cycle at Newlyn agrees with previous studies of sea level records from the North Sea, which have found an increase in the amplitude of the annual cycle, possibly associated with atmospheric forcing through winds and the NAO (Ekman and Stigebrandt, 1990; Ekman, 1999; Plag and Tsimplis, 1999). Our results indicate that at Newlyn the changes in the annual cycle of sea level are mainly associated to a steric rather than an atmospheric mechanism, since they are driven by changes in spring and autumn, rather than in winter, when the influence of the NAO state is more pronounced. Furthermore, variations in the annual amplitude of sea level are correlated with variations in the annual cycle of SST, but not with the NAO winter index.

Newlyn, Brest and Halifax display a similar behaviour not only in terms of the association of changes in sea level annual amplitude with variations in the amplitude of the annual cycle of SST, but also in terms of phase variability, with maximums occurring about 1 month early after the 1990s for the three stations. The cause of this change in phase is not clear, warranting further investigation.

At Vigo, the amplitude of the annual sea level cycle exhibits a positive trend, resulting from an increase in winter sea levels. Furthermore, the variations in the amplitude of the annual sea level cycle are negatively correlated with the winter NAO index. On the other hand, the amplitudes of the annual cycle of SST display a negative trend. Previous studies addressing sea level variability at Vigo found a strong response of sea level to meteorological forcing, particularly to winds (Garcia-Lafuente et al., 2004) and a negative correlation between sea level and the winter NAO index (Woolf et al., 2003; Fenoglio-Marc et al., 2005). Considering hydrostatic effects alone (inverse barometer response), a negative correlation is expected between the NAO winter index and sea level, since positive values of the NAO index reflect a stronger than usual subtropical high pressure centre, and therefore lower heights of the sea surface. However, the state of the NAO influences sea level not only through a direct hydrostatic response but also through non-hydrostatic effects in wind stress, surface fluxes and oceanic circulation. The variability in the annual cycle of sea level at Vigo, driven by changes in the winter period, seems to be associated mainly with these non-hydrostatic influences of the NAO state, since a positive correlation with the winter NAO index has been found, while for SST and atmospheric pressure the correlations are not significant. Changes in wind patterns associated with NAO variability, in particular the predominantly positive NAO index since the early 1980s, are a possible explanation for the detected variations in winter sea level values at Vigo and for the lower amplitudes of the SST annual cycle in the later part of the record.

The annual sea level amplitudes at Portland exhibit a weak but statistically significant positive correlation with the winter NAO index, suggesting an influence of the NAO state on sea level, very possibly through both hydrostatic and non-hydrostatic effects. From a purely inverse barometer response, a positive correlation with the NAO would be expected, due to the influence of the cyclonic low-pressure centre in the Northern Atlantic. Positive values of the winter NAO index are associated with a deeper than normal pressure low and thus higher sea levels. A positive correlation with the NAO index is also found on the eastern boundary for tide gauges at northern latitudes (Wakelin et al., 2003; Woolf et al., 2003; Yan et al., 2004).

On the western boundary, tide gauges from Boston to Fort Pulaski exhibit a negative correlation for the amplitudes of the annual cycles of sea level and SLP. This hydrostatic influence reflects the tendency of higher (lower) than average sea level maxima or minima to be associated with lower (higher) than average atmospheric pressure. Although at these locations the annual amplitudes of SST exhibit a positive trend, particularly high at New York and in Chesapeake Bay, the variations in the amplitude of the annual cycle are uncorrelated with SST, suggesting that atmospheric effects are the dominant factor in the variability of the annual sea level cycle. Once again, coastal sea level is not influenced by a single factor, but by a complex interaction of several climate and oceanographic effects. Chesapeake Bay is particularly influenced by hydrologic effects, which can induce sea level changes not only through variations in water content but also through density variations due to changes in salinity. Since SST exhibits a high positive trend, and variations in the annual amplitude of sea level and SST are not correlated, the negative trend in the annual amplitude of sea level at Baltimore is unlikely to be related to steric effects through thermal expansion/contraction, but can have a steric contribution associated with variations in salinity.

The amplitude of the annual sea level cycle for the stations in the Southern Bight (Charleston, Fort Pulaski and Mayport) is negatively correlated with the NAO index. Charleston and Fort Pulaski display an additional negative correlation with atmospheric pressure. It is not possible to distinguish whether the association between sea level and pressure variations at these two stations results from an independent, pressure-driven, hydrostatic contribution or reflects an hydrostatic response to the state of the NAO, or both. At Mayport, the annual sea level amplitudes are correlated only with the winter NAO index, suggesting a non-hydrostatic influence of the NAO on sea level variations at this site. A non-hydrostatic contribution associated to NAO variability cannot also be excluded for Charleston and Fort Pulaski. Results from an ocean general circulation model indicate that interannual sea level changes in the North Atlantic (including the area of the Southern Bight) are mainly related to the response of large-scale ocean circulation to changes in wind stress associated with NAO variability, while local heat fluxes are of minor importance (Esselborn and Eden, 2001). Therefore, a possible explanation for the correlation between variations in the annual amplitude of sea level and the winter NAO index is that changes in sea level are reflecting variations in oceanic circulation due to NAO-driven changes in wind stress, particularly noting the coastal path of the Gulf Stream and the association between coastal sea level variations in the Southern Bight and Gulf Stream transport (Ezer, 2001).

6. Conclusions

Sea level seasonality has been analysed from 16 tide gauge records at coastal sites in the eastern and western boundary of the North Atlantic Ocean. The use of a larger number of records, although desirable, is hampered by the lack of suitable long and continuous records for the study area.

For each sea level record a time-varying description of the annual cycle, which is the dominant periodicity in the analysed time series, has been obtained by autoregressive decomposition. A key issue in the analysis of time-varying seasonality is how to separate seasonal from interannual and long-term variability in the mean. This is a challenging task, requiring the application of specific methods (see also the discussion in Pezzulli et al., 2005). The advantages of the method used in this work are two-fold. First, it is a model-based approach and thus the resulting seasonal pattern is not dependent on an arbitrary concept of smoothness; the use of an explicit statistical model provides an adequate inferential framework. Second, within the modelling perspective, autoregressions are one of the simplest models. The approach is fairly straightforward, avoiding convergence issues which often plague more complex models (such as structural models). Furthermore, the Bayesian analysis can be accomplished in a simple and direct numerical framework with no need of Monte Carlo simulation methods.

The application of this autoregressive decomposition approach in the examination of the seasonality of coastal sea level has provided new estimates of the annual pattern for the analysed tide gauges and a new detailed description of fluctuations in the annual amplitude. It is shown that sea level in the North Atlantic exhibits considerable variability in the amplitude of the annual cycle. Furthermore, the changes in the annual cycle of sea level are regionally coherent and associated with changes in the annual amplitude of forcing variables. However, there is not a single factor responsible for the detected annual variations. Distinct sites show distinct influences reflecting the local nature of tide gauge measurements and the diversity of local effects influencing coastal sea level.

Thermal effects are the dominant influence at the northern stations, specifically Newlyn, Brest and Halifax, with fluctuations in the annual amplitude being associated with concurrent changes in SST. Atmospheric effects are the dominant influence at most of the stations on the western boundary of the North Atlantic. The state of the NAO influences sea level both through hydrostatic and non-hydrostatic effects and is the dominant factor determining the fluctuations in the annual amplitude of sea level at Vigo and Portland. The state of the NAO also influences the annual variability in the Southern Bight, possibly through the effect on sea level of NAO-related changes in wind stress and ocean circulation.

Sea level seasonality is affected by many different climate and oceanographic forcing processes acting over different spatial and temporal scales. Besides temperature and atmospheric pressure, sea level at coastal sites is influenced by winds and hydrologic variations, which have not been directly considered in this study.

An useful extension of this work would be the analysis of the seasonal variability of sea level and forcing variables from runs of climate models.

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