Int J Biometeorol (1990) 34:136–145

International Journal of Biometeorology IJB

Original articles

The effects of monthly temperature fluctuations on mortality in the United States from 1921 to 1985

Ulla Larsen

Department of Sociology, State University of New York at Stony Brook, Stony Brook, NY 11794-4356, USA

Received August 28, 1989; revised May 29, 1990
Accepted May 30, 1990

Abstract. The impact of short-term temperature fluctuations on mortality has been studied mainly on historical populations, thus providing a limited ability to generalize to contemporary conditions, which would be more useful in determining public health policies aimed at reducing mortality. Therefore, this study examined the effects of monthly temperature fluctuations on mortality in the United States from 1921 to 1985. Monthly data about mortality from the Vital Statistics and temperature from the National Oceanic and Atmospheric Administration and the US Department of Agriculture Weather Bureau were used. Six states were selected to be studied (Massachusetts, Michigan, Washington, Utah, North Carolina, and Mississippi). The analysis was carried out using distributed lag models. The analysis showed that warmer than usual temperatures in July and August, and unusually cold temperatures from January to June are linked to higher mortality. From September to December unusually low temperatures are associated with higher mortality in most states, while temperature has no significant effect on mortality in June and September. In January and February mortality is especially affected by unusually cold weather in the southern states of Mississippi and North Carolina. For example, a one degreee drop in the mean temperature in 1921 is associated with a more than 3.5% increase in the February crude death rate in Mississippi and North Carolina and a less than 1% increase in the four other states examined. Finally, in the months from January to March the relationship between monthly fluctuations in the crude death rate and temperature declined over time and became relatively weak by 1985.

Key words: Distributed lag model – Short-term fluctuation – Seasonal pattern – Temperature – Monthly crude death rate

Introduction

Previous studies of historical populations in Europe have demonstrated that unusually cold and warm weather is followed by higher mortality (Galloway 1987; Lee 1981). In contemporary societies the seasonal pattern in mortality has been examined extensively (Lerner and Anderson 1963), but relatively little attention has been paid to the relationship between mortality and extreme weather conditions. It is conceivable that this link between mortality and short-term weather fluctuations has vanished during the 19th and 20th centuries as a result of, for instance, modern housing and clothing which may have ameliorated the effects of unusual weather conditions. However, there is no evidence to reject the hypothesis that short-term weather fluctuations affect mortality in contemporary United States, and if this relationship were significant it could be taken into consideration in public health policies aimed at reducing mortality. Furthermore, public health policies might be more effective if we knew whether specific population groups (e.g. children and elderly people, blacks and hispanics, or economically disadvantaged people) were especially vulnerable to unusually cold or hot weather. More knowledge about the interaction between temperature and mortality might also guide us in understanding which diseases are more likely to cause death or to have a higher incidence at times with unusually cold or hot weather. For instance, differentials in mortality (e.g. the sudden infant death syndrome) may be partly explained by short-term temperature fluctuations.

The weather varies substantially across the United States. To obtain a more homogeneous unit of analysis six states were selected to be studied (Washington, Utah, North Carolina, Mississippi, Massachusetts, and Michigan). These states are geographically scattered across the United States and form a representative sample of the various geographic and climatic regions (Visher 1954). Registration of deaths started at different points in time in each state. Hence, to compare the results obtained across states the analysis is confined to the period from 1921 to 1985 for which all the selected states have registered number of deaths by month. Furthermore, by 1921 the repercussions of the big influenza epidemic in 1918 have disappeared. The year 1985 is the most recent year for which data are available.

This study analyzes the relationship between weather and mortality. More specifically, we examine the seasonal pattern and the lag structure between monthly fluctuations in temperature and mortality, as well as the variations by state of residence, and the changes over time from 1921 to 1985.

Model

Weather can be measured by precipitation, temperature, and humidity. Previous studies showed that short-term fluctuations in mortality are associated mainly with short-term fluctuations in temperature (Eckstein et al. 1985; Lee 1981). Hence, the present study is confined to analyzing short-term fluctuations in temperature.

Either very cold winters or very hot summers can lead to higher mortality. For instance, the incidence of hypothermia increases with colder temperatures (Knochel 1985). In most cases, unusually cold weather does not lead to hypothermia and subsequent death; instead it enhances the incidence of certain diseases and the risk of mortality. The major diseases associated with mortality at times with unusually cold winter temperatures are respiratory diseases, strokes, and heart diseases (Bull and Morton 1975, 1978; Keatinge et al. 1984; Clark and Edholm 1985). Very warm summers are associated with an increase in infectious and intestinal diseases, as well as a higher incidence of heat strokes and heart failures (Ellis 1972; Knochel 1985). In historical England, Lee (1981) found that colder temperatues from December to May, and warmer temperatures in the months from June to November were associated with higher mortality.

During this century man has obtained more control over the environment and hence has become better able to protect himself against climatic features, such as extreme temperatures. For instance, better heating, air conditioning and clothing, as well as improved hygiene, nutrition, and medical innovations are linked to less pronounced seasonal patterns in mortality throughout this century (Lerner and Anderson 1963). As an example, refrigerated food storage and antibiotics have reduced the incidence of infectious diseases leading to deaths (Ellis 1972). Based on these previous findings, we expect the impact of temperature fluctutations on mortality to have declined during the period from 1921 to 1985, and to be more pronounced in populations with a lower level of wealth or living standard.

Temperature fluctuations might have both an immediate and a delayed effect on mortality, i.e., some people die shortly after they catch a disease, while others suffer from the disease for a while before death occurs. The literature is inconclusive with respect to the lag structure between extreme temperatures and mortality. Lee (1981) found in historical England that the effects of fluctuations in the winter temperatures (December to May) were strongest at lag zero (in the same month), while the summer temperatures (June to November) affected mortality more after a 1-month delay, and the effects of both winter and summer temperature fluctuations tapered off after 3 months. Monthly data have been used in no other time series analysis of the relationship between temperature and mortality. Galloway (1987), however, found on the basis of yearly data that the winter temperature (average of the temperatures in January, February, and March) had an immediate and a 1-year delayed effect, while the summer temperature (average of the temperatures in July, August, and September) had a strong immediate effect and a smaller effect after 1 year in several preindustrial European countries. Hence, we used an exploratory approach in the analysis of the lag structure between fluctuations in the temperature and mortality. Specifically, we investigated whether the lag structure varies across seasons. Finally, the literature suggests that variability in the weather has an even greater effect on mortality than temperature extremes (Howe 1972; Tromp 1980). In this context, it should be noted that temperature fluctuations are generally greater during cold weather, and this notion is taken into consideration in the interpretation of the results obtained.

Operationalization

Mortality

Total counts of deaths by month and by state of residence are published in the vital statistics for the US (Mortality Statistics; Vital Statistics of United States). These counts were converted into monthly crude death rates, CDR_m, using the following formula (Shryock and Siegel 1973):

$$CDR_m = [(D_m/N_m) * 365 * 1000]/P,$$

where D_m is the number of deaths recorded in month m, N_m represents the number of days in month m, and P is the midyear population. This formula corrects the number of deaths recorded in various months for the differing lengths of the months. It thus produces rates that are comparable in scale to annual crude death rates. Yearly data about the midyear population by state are published in Statistical Abstract of the United States.

Temperature

The temperature is measured by the average monthly temperature in degrees Fahrenheit read at one station in each state (Climatological Data; Report of the Chief of the Weather Bureau). The selected stations are located in Boston, Detroit, Seattle, Salt Lake City, Charlotte and Vicksburg in respectively Massachusetts, Michigan, Washington, Utah, North Carolina, and Mississippi. (There might be differences in the level of the temperature within a state, but the monthly fluctuations are very similar. For each state from 1921 to 1955 the temperature read at the selected station and an average measure based on readings from many stations are highly correlated. The average measure is not published after 1955.)

Economy

No monthly data about the economy are available for the entire period from 1921 to 1985 at the state level. However, each census from 1930 to 1980 gives data about annual income per capita, and this measure might be used as an indicator of the general level of wealth within each state (Statistical Abstract of the United States).

Methods

The relationship between monthly crude death rates, temperature, and time period discussed above may be specified as follows:

$$CDR_{t,m} = \prod_{k=0}^{n} TEMP_{t,m-k}^{(\alpha_{m,k}+\beta_{m,k} TIME_{m-k})} \cdot \exp(\chi_m + \delta_{m,1}\, TIME + \delta_{m,2}\, TIME^2 + \varepsilon_{t,m}) \qquad (1),$$

where CDR = crude death rate per 1000 persons, (see above for the computation of CDR); $TEMP$ = temperature in degrees Fahrenheit; $TIME$ = time period; $t=1$ (1921), 2 (1922), ..., $T=65$ (1985); $m=1$ (January), 2 (February), ..., 12 (December); $k=0$, 1, ..., 3; $n=3$.

The crude death rate in month m at time t ($CDR_{t,m}$) is expressed as a function of the temperature ($TEMP_{t,m-k}$) at time t in month m to $m-k$, and k takes the values from 0 to n. The exponent $(\alpha_{m,k}+\beta_{m,k}\, TIME_{m-k})$ allows the effects of temperature to change over time. The crude death rate ($CDR_{t,m}$) is hypothesized to decline over time in accordance with an exponential curve, and to make this specification more flexible the relationship is set to be a second order polynomial of time.

The model specified in Eq. 1 requires monthly data. One relationship is estimated for the crude death rate CDR in each calendar month m. This allows each calendar month's temperature to exert a different effect on mortality, since for example, a high winter temperature might have a different effect than a high summer temperature. It is also necessary to prevent the seasonality of vital events and temperature from influencing the analysis, since otherwise possible spurious correlations might be found. Specifically, we are interested in learning whether an unusually cold winter month leads to unusually high mortality, not whether mortality is typically high in typically cold months. The latter problem has already been examined in detail (Lerner and Anderson 1963). Analyzing each calendar month's mortality separately abstracts from the normal seasonal patterns, and so avoids this problem.

A natural-log transformation transforms the multiplicative model specified in Eq. 1 into an additive model. In order to study short-term fluctuations the series analyzed must be stationary, i.e., have no trend or drift and a constant variance. Otherwise, the estimated relationship between short-term fluctuations in two or more series might be disguised by, for instance, a common trend. Stationarity is obtained by taking first differences of the natural log-transformed series, and Eq. 1 becomes:

$$\Delta \ln CDR_{t,m} = \sum_{k=0}^{n} (\alpha_{m,k}+\beta_{m,k}\, TIME_{m-k})\, \Delta \ln TEMP_{t,m-k} + 2\delta_{m,2}\, TIME + (\delta_{m,1}+\delta_{m,2}) + \Delta\varepsilon_{t,m} \qquad (2).$$

The time period from 1921 to 1985 contains sufficient observations (65) to carry out a time series analysis for each of the six states analyzed. As an alternative, the data can be pooled, and a time series and cross state of residence analysis be performed for all states combined (Pindyck and Rubenfeld 1981). The latter approach enables us to test a number of hypotheses, such as whether the effects of temperature on mortality vary by state of residence because temperature may have a greater impact on mortality in economically poor states. The saturated model based on pooled data leads to exactly the same results as calculating one time series model for each state. However, it is possible that we can constrain some of the parameters in the saturated model to be zero, and in that way obtain a more parsimonious model.

When the data are pooled Eq. 2 can be written as:

$$\Delta \ln CDR_{t,m} = \sum_{s=1}^{S}\sum_{k=0}^{n} (\alpha_{m,k}+\beta_{m,k}\, TIME_{m-k})\, \Delta \ln TEMP_{t,s,m-k} + 2\delta_{m,2}\, TIME + (\delta_{m,1}+\delta_{m,2}) + \Delta\varepsilon_{t,m,s} \qquad (3),$$

where $s=1$ (Massachusetts), 2 (Michigan), 3 (Washington), 4 (Utah), 5 (North Carolina), $S=6$ (Mississippi), and otherwise as defined in Eq. 1.

In Eq. 3 the total number of observations is $ST=6*65=390$. It is plausible that the effects of temperature on mortality vary by state of residence, i.e., the responsiveness of mortality to short-term fluctuations in temperature is dependent on the general level of wealth at the place of residence. (We acknowledge that variations by state of residence may be due to factors other than wealth, and more studies are needed.) As a measure of wealth, income per capita given at each census could have been used in the model directly. Instead, we substituted a variable measuring wealth for a dummy variable D, and ranked the six states analyzed according to their income per capita. The same ranking occurs at each census from 1930 to 1980: the richest state is Massachusetts, then Michigan, Washington, Utah, North Carolina, and the poorest is Mississippi. Introducing dummy variables and assuming a linear relationship the parameters $\alpha_{m,k}$ and $\beta_{m,k}$ can be expressed as follows:

$$\alpha_{m,k} = \phi_{m,k} + \gamma_{m,k} D_{sj} \quad \text{and} \quad \beta_{m,k} = \eta_{m,k} + \iota_{m,k} D_{sj} \qquad (4),$$

where, $m=1, 2, \ldots, 12$; $k=0, 1, \ldots, 3$; $s=1, 2, \ldots, 6$;

$j=1, 2, \ldots, 5$; $D_{sj}=1$ for $s=j$, and $D_{sj}=0$ otherwise.

Substituting 4 into 3 we obtain:

$$\Delta \ln CDR_{t,m} = \sum_{s=1}^{S}\sum_{j=1}^{S-1}\sum_{k=0}^{n} (\phi_{m,k}+\gamma_{m,k} D_{s,j} + (\eta_{m,k}+\iota_{m,k} D_{s,j})\, TIME_{m-k}) \cdot \Delta \ln TEMP_{t,s,m-k} + 2\delta_{m,2}\, TIME + (\delta_{m,1}+\delta_{m,2}) + \Delta\varepsilon_{t,m,s} \qquad (5).$$

In Eq. 5 it is assumed that the intercept $(\delta_{m,1}+\delta_{m,2})$ is constant across states. This assumption might not hold in the pooled model. To allow the intercept to vary across states more dummy variables, D, are used and Eq. 5 becomes:

$$\Delta \ln CDR_{t,m} = \sum_{s=1}^{S}\sum_{j=1}^{S-1}\sum_{k=0}^{n} (\phi_{m,k}+\gamma_{m,k} D_{s,j} + (\eta_{m,k}+\iota_{m,k} D_{s,j})\, TIME_{m-k}) \cdot \Delta \ln TEMP_{t,s,m-k} + 2\delta_{m,2}\, TIME + \lambda_m D_{s,j} + (\delta_{m,1}+\delta_{m,2}) + \Delta\varepsilon_{t,m,s} \qquad (6).$$

The model specified in Eq. 6 is estimated using regression analysis. The estimation is carried out using the software Micro TSP, version 5.0 (Lilien 1986). Serial autocorrelation is corrected for by a second order autoregression.

In interpreting the results, we are interested mainly in studying proportional changes in mortality due to a small change in temperature. Recall, we specified the relationship between mortality and temperature as follows:

$$CDR_{t,m} = \prod_{k=0}^{n} TEMP_{t,m-k}^{(\alpha_{m,k}+\beta_{m,k} TIME_{m-k})} \cdot \exp(\chi_m + \delta_{m,1}\, TIME + \delta_{m,2}\, TIME^2 + \varepsilon_{t,m}) \qquad (1).$$

Table 1. Contrasts between models

Model	Constraint	Contrast	d.f.	χ^2 Month											
				Jan	Feb	Mar	Apr	May	Jun	Jul	Aug	Sep	Oct	Nov	Dec
A	$(\iota)=0$														
B	$(\iota, \lambda)=0$	A vs. B	5	1.6	2.4	0.9	1.3	0.4	0.5	1.5	0.9	2.2	2.1	0.8	1.5
C	$(\iota, \lambda, \delta_2)=0$	A vs. C	6	1.7	7.3	3.1	2.1	1.3	1.2	1.5	1.5	2.7	2.5	0.8	1.5
D	$(\theta, \lambda, \delta_2, \gamma)=0$	A vs. D	26	42.0[b]	73.7[a]	32.6	25.8	33.8	25.1	34.2	15.2	23.0	35.5	32.7	19.1
E	$(\iota, \lambda, \delta_2, \gamma, \eta)=0$	A vs. E	30	49.2[b]	107.2[a]	44.5[b]	31.6	38.4	28.7	38.5	16.7	24.4	38.3	36.9	29.1
F	$(\iota, \lambda, \delta_2, \gamma, \eta, \phi)=0$	A vs. F	34	84.8[a]	136.9[a]	82.9[a]	63.1[a]	46.9[c]	43.3	72.5[a]	47.4[c]	27.9	50.8[b]	64.5[a]	64.5[a]

[a] Significant at the 0.01 level; [b] significant at the 0.05 level; [c] signficant at the 0.10 level

When comparing two models a small likelihood ratio (and small χ^2 relative to its d.f.) indicates that the less restrictive model and the more restrictive one (with more parameters set equal to zero) fit the observations nearly as well; the more restrictive model is preferred

The natural-log transformation of Eq. 1 leads to:

$$\ln CDR_{t,m} = \sum_{k=0}^{n} (\alpha_{m,k} + \beta_{m,k}\, TIME_{m-k}) \ln TEMP_{t,m-k} + (\chi_m + \delta_{m,1}\, TIME + \delta_{m,2}\, TIME^2 + \Delta\, \varepsilon_{t,m}) \quad (7).$$

The effect of a small change in temperature on mortality can be illustrated in different ways, e.g., the effect of a change in one or several months' temperature. As an example, we use the total effect on $CDR_{t,m}$ when all relevant months' temperature change. In this case, the first derivative of $\ln CDR_{t,m}$ with respect to $TEMP_{t,m-k}$ ($k=0$, 1, 2, and 3) becomes:

$$\sum_{k=0}^{n} \partial \ln CDR_{t,m} / \partial\, TEMP_{t,m-k} = \sum_{k=0}^{n} (\alpha_{m,k} + \beta_{m,k}\, TIME_{m-k}) / TEMP_{t,m-k} \quad (8).$$

Substituting Eq. 4 into Eq. 8 and $\overline{TEMP_m}$ for $TEMP_{t,m}$ (the average temperature $\overline{TEMP_m}$ between 1921 and 1985 is used to exclude random variation) we obtain:

$$\sum_{k=0}^{n} \partial \ln CDR_{t,m} / \partial\, TEMP_{t,m-k} = \sum_{s=1}^{S} \sum_{j=1}^{S-1} \sum_{k=0}^{n} (\phi_{m,k} + \gamma_{m,k} D_{s,j} + (\eta_{m,k} + \iota_{m,k} D_{s,j})\, TIME_{m-k}) / \overline{TEMP_{m-k}} \quad (9).$$

The best model is selected by fitting the model of the crude death rate by month to all the covariates analyzed, and on the basis of theoretical considerations constrain parameters to zero. The likelihood ratio statistics are used to eliminate specific covariates.

Results

The analysis of crude death rates by month is structured as follows: (1) the hypotheses tested are listed; (2) the best model for each month is selected; (3) the specific lag structure in each month is determined; and (4) the findings are discussed with respect to whether there is a seasonal pattern, and whether the lag structure varies by season.

Selection of a parsimonious model

The strategy employed in the analysis of monthly crude death rates was to contrast a number of restricted models with the model expressed in Eq. 6. Thus, parameters were constrained to zero and a type of stepwise procedure was used in selecting the best fitting model. A priori the interactions between time period, state of residence, and temperature were constrained to be zero because preliminary analyses of each state of residence separately suggest that the effects of short-term fluctuations in the temperature do not affect mortality in each state differently during the period from 1921 to 1985 (model A); this restriction also makes the model less complex. In model B, proportional changes in the crude death rate in a given month are assumed to be the same in all states, i.e., the intercept is constant across states. It is assumed that there is no time trend in mortality in model C. In model D, the effects of temperature on mortality are hypothesized to be constant across states, and whether the effects of temperature declined over time is tested

Table 2. Determination of temperature lag-structures: lag 0, 1, 2, and 3 months vs. lag 0 and 1 month

Month	Model	χ^2
January	C	30.18[a]
February	C	51.82[a]
March	D	12.32[a]
April	E	1.62
May	E	0.10
June	F	
July	E	5.74[c]
August	E	2.60
September	F	
October	E	0.44
November	E	5.64[c]
December	E	2.40

[a] Significant at the 0.01 level; [b] significant at the 0.05 level; [c] significant at the 0.10 level

In all models, d.f. = 2

Table 3. The effects of temperature and the interactions between temperature and time period and temperature and state of residence on the crude death rate by month, 1921–1985

Independent variable	Dependent variable: Crude death rate Month											
	Jan	Feb	Mar	Apr	May	Jun	Jul	Aug	Sep	Oct	Nov	Dec
Temperature												
Lagged 0 month	−0.011[a]	−0.013[a]	−0.008[a]	−0.003[a]	−0.001	0.002[b]	0.007[a]	0.006[a]	−0.002[c]	−0.003[a]	−0.004[a]	−0.005[a]
1 month	0.001	−0.011[a]	−0.003[a]	−0.003[a]	−0.002[a]	−0.002[c]	−0.002[b]	−0.001	0.000	−0.001	0.000	−0.002
2 months	−0.003	−0.004	−0.001	−0.000	0.000	−0.002[c]	0.001	−0.001	−0.000	0.001	0.002[b]	0.001
3 months	−0.016[a]	−0.010[b]	0.001	−0.001	0.000	−0.000	−0.002[b]	0.001	−0.000	−0.000	−0.000	−0.002
Temperature and state of residence[d]												
Lagged 0 month												
Massachusetts	0.004	0.019[a]										
Michigan	0.009[c]	0.019[a]										
Washington	0.002	0.007										
Utah	0.009[c]	0.018[a]										
North Carolina	−0.000	−0.001										
Lagged 1 month												
Massachusetts	0.010[b]	0.007										
Michigan	0.008	0.007										
Washington	0.002	0.009[c]										
Utah	0.001	0.009[c]										
North Carolina	0.001	0.003										
Lagged 2 months												
Massachusetts	−0.010[c]	0.012[b]										
Michigan	−0.000	0.008										
Washington	0.007	0.002										
Utah	0.004	0.007										
North Carolina	−0'.000	−0.000										
Lagged 3 months												
Massachusetts	0.018[a]	0.004										
Michigan	0.016[a]	0.006										
Washington	0.013[c]	0.019[a]										
Utah	0.016[a]	0.011[c]										
North Carolina	0.007	−0.001										
Temperature and time period												
1921												
Lagged 0 month	0.000[a]	0.000[b]	0.000[b]									
1 month	−0.000	0.000[b]	0.000									
2 months	0.000	0.000[c]	0.000									
3 months	0.000[b]	0.000[a]	0.000									
1940												
Lagged 0 month	0.003[a]	0.003[b]	0.002[b]									
1 month	−0.002	0.003[b]	0.001									
2 months	0.000	0.003[c]	0.001									
3 months	0.004[b]	0.005[a]	0.001									

Table 3 (continued)

Independent variable	Dependent variable: Crude death rate											
	Month											
	Jan	Feb	Mar	Apr	May	Jun	Jul	Aug	Sep	Oct	Nov	Dec
1960												
Lagged 0 month	0.006[a]	0.005[b]	0.004[b]									
1 month	−0.004	0.006[b]	0.002									
2 months	0.001	0.005[c]	0.001									
3 months	0.008[b]	0.009[a]	0.002									
1985												
Lagged 0 month	0.013[a]	0.011[b]	0.009[b]									
1 month	−0.009	0.013[b]	0.004									
2 months	0.002	0.011[c]	0.002									
3 months	0.017[b]	0.022[a]	0.004									

[a] Significant at the 0.01 level; [b] significant at the 0.05 level; [c] significant at the 0.10 level

The effects are estimated in model C for January and February, in model D for March, and for the rest of the year in model E. The effects' estimates have been scaled to show the proportional change in a month's crude death rate resulting from a change of one degree Fahrenheit in temperature. Some effects are so small, that the coefficients are less than 0.000

[d] The effects' estimates for January and February are relative to Mississippi, the omitted category

in model E. Finally, in model F the effects of temperature are eliminated and proportional changes in the crude death rate by month are modelled by a constant.

Proportional changes in the crude death rate do not vary across state of residence during the period from 1921 to 1985 in any one month of the year; models B and C fit the data nearly as well as model A (Table 1). (In Table 1 and Table 2 an[a] after a coefficient indicates that the difference in fit between two models is significantly different from zero at the 0.01 level based on a χ^2 test; a[b] indicates the 0.05 level; and a[c] the 0.10 significance level.) The effects of temperature on mortality differ across state of residence in January and February, and in these months as well as in March, fluctuations in the temperature have a time dependent effect on mortality. Furthermore, short-term fluctuations in the temperature affect proportional changes in the crude death rate directly in every month except June and September, i.e., there is a significantly reduced fit in moving from model E to F relative to model A. (The same results were found when the following models were contrasted: A vs. B, B vs. C, C vs D, D vs. E, and E vs. F.)

In all the models estimated, the impact of temperature is modelled with a 0, 1, 2, and 3-month delay, but few of the temperature parameters beyond a 1-month delay are significantly different from zero (based on a two-tailed t-test at the 5% level). Therefore, we tested whether models with temperature lagged only 0 and 1 month provided as good a fit as models with temperature lagged up to 3 months (Table 2). This exploratory work suggested that fluctuations in the temperature had a delayed effect up to 3 months in January, February, and March as well as in July and November, although in the latter 2 months the less restrictive models (including delayed effects up to 3 months) fit the data better only at the 0.10 level of statistical significance.

Seasonal pattern between mortality and weather

We are interested primarily in determining the seasonal pattern between temperature and the crude death rate and the underlying lag structure. To facilitate this analysis the effects estimated in the best model including temperature lagged up to 3 months are analyzed for each month (Table 3). (In Table 3, an[a] after a coefficient indicates that the coefficient is significantly different from zero at the 0.01 level based on a two-tailed t-test; a[b] indicates the 0.05 level; and a[c] the 0.10 level of significance.)

In order to understand the impact of temperature on mortality by month the cumulated temperature effects are discussed. A one degree Fahrenheit drop in the temperature throughout the year is associated with a proportional increase in the crude death rate in each month with the exception of July and August (Fig. 1). In July and August cooler weather reduces mortality. Also, temperature fluctuations do not affect mortality significantly in either June or September, according to the analysis above (Table 3). In both January and February, proportional changes in the crude death rates are more affected

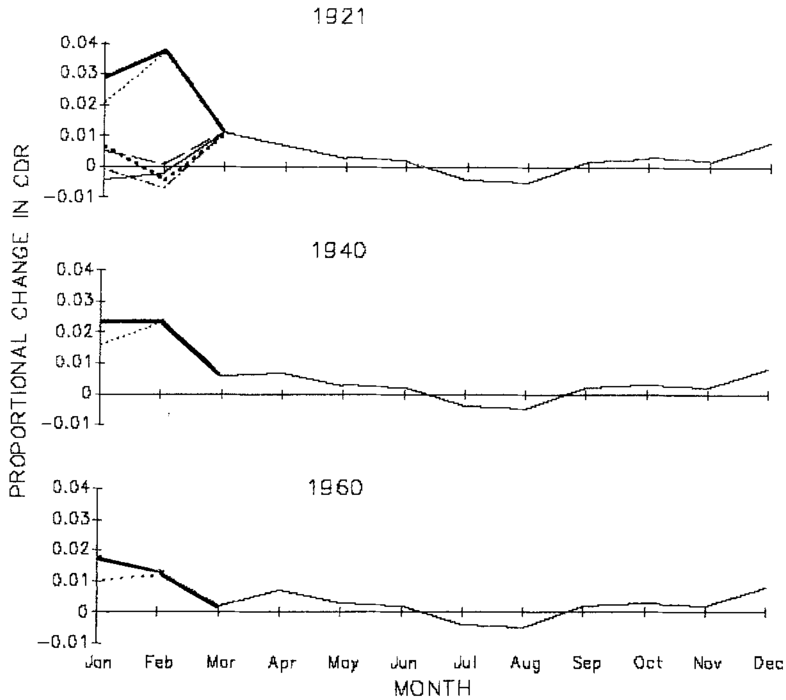

Fig. 1. The lagged summed effect of a one degree Fahrenheit decline in temperature on the crude death rate by month in 1921, 1940, and 1960. Note: The verticale scale measures the proportional change in the monthly crude death rate from its mean level following a decline of one degree Fahrenheit in the temperature in each month of the year. In January and February the lagged summed effect of temperature on the crude death rate varies by state of residence, and from January to March it varies by time period. Source: Table 3. ··· North Carolina; — Mississippi; — All states

by temperature fluctuations in North Carolina and Mississippi than in the other four states examined: Massachusetts, Michigan, Washington, and Utah. North Carolina and Mississippi are two southern states, and of the six states analyzed these two states rank the lowest with respect to income per capita. (The same ranking occurred at each census from 1930 to 1980: Mississippi ranked poorest, followed respectively by North Carolina, Utah, Washington, and Michigan, and Massachusetts was the richest state.) However it is conceivable that the variation by state of residence may be explained by other factors than income per capita.

The finding that in January and February the lagged summed temperature effects are greater in the poorer states examined, and from January to March the impact of temperature on mortality declined significantly during the period studied, suggests underlying socioeconomic factors are intervening, i.e., lethal effects of extreme temperatures are ameliorated by better clothing, housing, and medical care. However, it is surprising that the effects of temperature on mortality declined significantly from 1921 to 1985 only in the months from January to March. It was expected also that the effects of unusually hot weather in the summer had declined. Refrigerators, air conditioning, and generally better sanitation in combination with medical innovations like antibiotics were hypothesized to have reduced, among others, intestinal and infectious diseases leading to lower mortality. Here, it should be pointed out that in North Carolina and Mississippi unusually cold weather had a stronger effect on mortality in January and February than unusually hot weather in July and August, and this circumstance might explain why only effects of unusually cold temperatures declined during the period studied. Finally, the weather is on average warmer in Mississippi and North Carolina than in the other four states examined. Consequently, in Mississippi and North Carolina people may be less prepared for very cold weather, and therefore a cold spell has more lethal effects in these southern states.

To discern better the seasonal relationship between temperature and mortality the proportional change in the crude death rate due to a one degree Fahrenheit drop in temperature in a given month is summed over the four succeeding months (Fig. 2). A drop in the January to June temperatures is associated generally with higher mortality, while the inverse relationship prevails between mortality and the temperatures in July and August. A slightly more heterogeneous pattern is found in the fall months from September to December, although in general a lower temperature is associated with higher mortality. In line with these findings we suggest dividing the pattern between temperature fluctuations and mortality into three seasonal patterns; a winter pattern including the months form January to August (lower temperatures are associated with higher mortality), a summer pattern containing the months of July and August (lower temperatures are associated with lower mortality) and a fall pattern from September to December (lower temperatures are associated generally with higher mortality).

The seasonal pattern between temperature and motality in 20th century US is similar to the summer and winter patterns in historical England estimated by Lee (1981). However, the summer pattern is more pronounced in historical England: lower temperatures are associated with lower mortality from June to November, and the mortality response is greater to a given change in the temperature. In historical England the strongest temperature effects are found in July and September, and in these months a one degree Fahrenheit decline has a cumulative effect of approximately a 5% decline

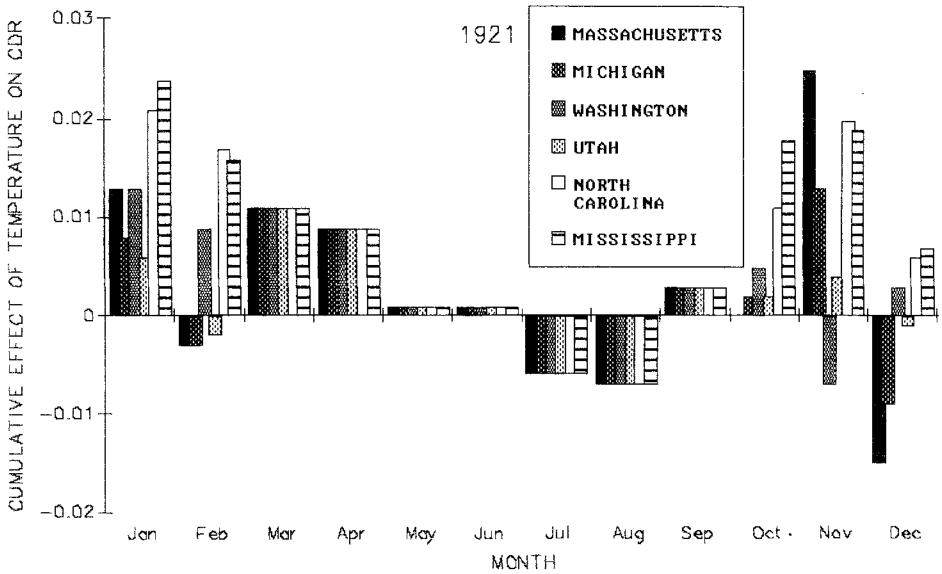

Fig. 2. The cumulative effect of a one degree Fahrenheit decline in each month's temperature on the crude death rate by month in 1921. Note: The effects' estimates are cumulated over lags of 0 to 3 months and express the cumulative effect of a decline in temperature of one degree Fahrenheit. From October to February the cumulative effect of temperature on the crude death rate varies by state of residence. Source: Table 3

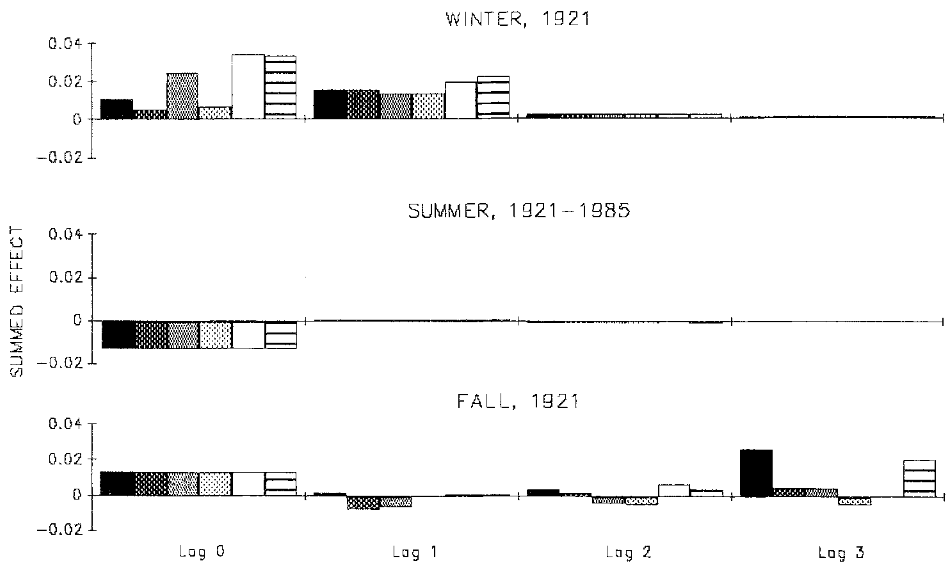

Fig. 3. Lag structure in the winter, summer and fall patterns of the effect of a one degree Fahrenheit decline in temperature on the crude death rate. Note: Winter refers to the months of January to June; summer, July and August; fall, September to December. In the winter and fall patterns the effects of temperature on the crude death rate vary by state of residence. The vertical scales measure the sum of all the effects' estimates at each indicated lag. Source: Table 3

in deaths. In contrast, a one degree Fahrenheit decline in the July temperature in the six American states analyzed is associated with a 0.6% decline in the crude death rate. On the other hand, a one degree Fahrenheit decline in the January temperature is linked to, for example, a 2.4% increase in the crude death rate in Mississippi in 1921, and to about a 2% increase in deaths in historical England.

Lag structure between mortality and weather

Fluctuations in the summer temperature (the effects' estimates in July and August) have an immediate effect on mortality, but almost no delayed effect (Fig. 3). In contrast, in the winter (January to June) the summed temperature effect is substantial at both lag 0 and 1 month. In the fall (September to December) temperature fluctuations affect mortality mainly at lag 0 month, although there is a delayed effect at lag 3 months in Massachusetts and Mississippi. This finding is in line with models with temperature lagged up to 3 months compared to models with temperature lagged up to 1 month which fit the data significantly better at the 0.01 level only in the months of January, February, and March (Table 2). The lag structure found in historical England (Lee 1981) is slightly different: the winter (December to May) temperature effect is greatest at lag zero, while the summer effect (June to November) peaks after a 1 month delay, and both the winter and summer temperatures have almost no effect on mortality after a 3-month delay. In order to explain differentials in the lag structure across seasons in 20th century USA (as well as in historical England), further studies of the underlying disease pattern might be fruitful.

Conclusions

The analysis of monthly crude death rates demonstrated that unusually cold weather in the months from January to June, and September to December, and warmer than usual weather in July and August are associated with higher mortality in all six states analyzed from 1921 to 1985. An analysis of the lag structure between monthly fluctuations in mortality and temperature revealed that extreme temperatures in July and August affect mortality greatest in the same month, while from January to June unusual temperatures have a pronounced effect on mortality both in the same month and in the next month. From September to December short-term temperature fluctuations have mainly an immediate effect. A similar analysis could be carried out for each of the fifty states, but the additional work does not seem warranted, and we propose to generalize our findings from these six states to the entire United States.

The impact of unusually cold weather on mortality declined from 1921 to 1985 in the months of January, February, and March. For instance, on average a one degree Fahrenheit drop in the temperature from November to February in Mississippi was associated with a decline in the February crude death rate of 3.8% (60 deaths) in 1921, 2.4% (55 deaths) in 1940, and 1.3% (28 deaths) in 1960, while by 1985 the relationship between mortality in February and temperature vanished. These figures should be seen in light of the fact that the temperature does vary from year to year, e.g., the standard deviation of the monthly temperatures from November to February ranges from 3.3 to 5.0 in Mississippi, 1921–1985. The trend in the relationship between mortality and temperature supports the hypothesis that modern housing, clothing, and medicine may have contributed to a decrease in the fatal effects of very cold weather. This notion that economic factors ameliorate the lethal impact of extreme temperatures is further supported by the result that, in January and February, mortality is more responsive to unusual temperatures in the two poorer southern states analyzed (Mississippi and North Carolina) compared to Massachusetts, Michigan, Washington, and Utah.

In January and February, North Carolina and Mississippi have warmer weather than the other four states examined (higher average temperatures), but the weather is generally more variable (greater standard deviations, with the exception that the temperature is more variable in Utah). (The average January temperature is 29.19°, 25.49°, 40.46°, 28.88°, 41.67°, and 47.62° F and the standard deviation is 4.09, 4.41, 3.99, 5.38, 4.58, and 4.96 in Massachusetts, Michigan, Washington, Utah, North Carolina, and Mississippi respectively. The average February temperature is 30.31°, 27.12°, 43.39°, 34.42°, 44.16°, and 50.87° F and the standard deviation is 3.83, 4.35, 3.04, 5.04, 4.11, and 4.37 in the respective states.) Hence, there is some evidence for the notion that variability in the temperature is a greater mortality risk factor than extreme temperatures, at least in the winter when unusually low temperatures are associated with higher mortality. It would be interesting to examine further the effect of fluctuations in the temperature variance on mortality.

Overall mortality in 20th century USA is significantly affected by unusually cold winter weather and hot summer weather, but the association is relatively weak by 1985. Several factors do limit the strength of our conclusions. For example, we have not determined if temperature extremes have a greater impact on the very young or old, since crude death rates do not measure differentials in mortality by age or the effects of changes in the age structure on mortality. In addition, from 1921 to 1985 the causes of death have shifted from mainly acute diseases (infectious and respiratory diseases) to chronic diseases (cancer and cardiovascular diseases), which may obscure the gross effects attributale to weather extremes. Furthermore, only the effects of temperature fluctuations have been examined, although other components of the weather, such as atmospheric pressure fronts, number of hours of sunshine per day, or humidity may be important either in combination with the mean temperature or separately. Also, it may be more appropriate to analyze daily or weekly weather fluctuations than fluctuations by month.

This study was constrained by the limited American mortality data published for the period 1921 to 1985. However, significant results were obtained, and all the findings are in accordance with prior theoretical expectations. Therefore, it should be worthwhile to extend this somewhat general time series study to more specific cross-section analyses of data from recent years. In particular, we need to determine whether there are greater mortality responses to unusually cold winter weather and hot summer weather in specific states or regions of the US, and whether certain population groups are more susceptible, e.g., disadvantaged people. This analysis draws added policy relevance for the elderly from the fact that substantial funds have been set aside in the US federal budget, by the Low Income Energy Assistance Act, for expenditure to offset the effects on the elderly of climatic variation. Finally, this line of work is of general interest to determine the importance of temperature extremes relative to other underlying determinants of mortality.

Acknowledgements. This research was supported by National Institute of Aging Training Grant HD 07275-04. An earlier version of this paper was presented at the Annual Meeting of the American Sociological Association, Atlanta, August 24–28, 1988. The author thanks Patrick R. Galloway, Ron D. Lee, and Robert E. Strecker for many helpful suggestions.

References

Bull GM, Morton J (1975) Relationships of temperature with death rates from all causes and from certain respiratory and arteriosclerotic diseases in different age groups. Age Ageing 4:232–246

Bull GM, Morton J (1978) Environment, temperature and death rates. Age Ageing 7:210–224

Clark RP, Edholm OG (1985) Man and his thermal environment. Edward Arnold, London

Climatological Data. The National Oceanic and Atmospheric Administration

Eckstein Z, Schultz TP, Wolpin TP (1985) Short-run fluctuations in fertility and mortality in pre-industrial Sweden. Eur Econ Rev 26:295–317

Ellis FP (1972) Mortality from heat illness and heat-aggravated illness in the United States. Environm Res 5:1–58

Galloway PR (1987) Population, prices, and weather in preindustrial Europe. PhD thesis, University of California at Berkeley
Howe GM (1972) Man, environment and disease in Britain. Barnes and Noble, New York
Keatinge WR, Coleshaw SRK, Cotter F, Mattock M, Murphy M, Chelliah R (1984) Increases in platelet and red cell counts, blood viscosity, and arterial pressure during mild surface cooling: Factors in mortality from coronorary and cerebral thrombosis in winter. Br Med J 289:1405–1408
Knochel JP (1985) Disorders due to heat and cold. In: Wyngaarden JB, Smith LH (eds) Cecil textbook of medicine. Saunders, Philadelphia
Lee R (1981) Short-term variation: Vital rates, prices, and weather. In: Wrigley EA, Schofield RS (eds) The population history of England 1541–1871. Harvard University Press, Cambridge, pp 356–401
Lerner M, Anderson OW (1963) Health progress in the United States. 1900–1960: A report of health information foundation. University of Chicago Press, Chicago
Lilien DM (1986) Micro TSP, version 5.0
Mortality Statistics. Department of Commerce Bureau of the Census
Pindyck RS, Rubinfeld DL (1981) Econometric models and economic forecasts. McGraw-Hill, New York
Report of the Chief of the Weather Bureau. US Department of Agriculture Weather Bureau
Schryock HS, Siegel JS (1973) The methods and materials of demography. US Department of Commerce Bureau of the Census, p 396
Statistical Abstract of the Unted States. US Department of Commerce Bureau of the Census
Tromp SW (1980) Biometeorology. Heyden, London
Visher SS (1954) Climatic atlas of the United States. Cambridge
Vital Statistics of United States. US Department of Helath, Education and Welfare. Public Health Service