ON THE URBAN HEAT ISLAND EFFECT DEPENDENCE ON TEMPERATURE TRENDS

INÉS CAMILLONI and VICENTE BARROS
Department of Atmospheric Sciences, University of Buenos Aires, Ciudad Universitaria, Pabellón II, 2° piso, (1428) Buenos Aires, Argentina

Abstract. For U.S., Argentine and Australian cities, yearly mean urban to rural temperature differences (ΔT_{u-r}) and rural temperatures (T_r) are negatively correlated in almost every case, suggesting that urban heat island intensity depends, among other parameters on the temperature itself. This negative correlation is related to the fact that interannual variability of temperature is generally lower in urban environments than in rural areas. This seems to hold true at low frequencies leading to opposite trends in the two variables. Hence, urban stations are prone to have lower trends in absolute value than rural ones.

Therefore, regional data sets including records from urban locations, in addition to urban growth bias may have a second type of urban bias associated with temperature trends. A bulk estimate of this second urban bias trend for the contiguous United States during 1901–1984 indicates that it could be of the same order as the urban growth bias and of opposite sign.

If these results could be extended to global data, it could be expected that the spurious influence of urban growth on global temperature trends during warming periods will be offset by the diminishing of the urban heat island intensity.

1. Introduction

The urban heat island phenomenon has been documented for many cities with varying population, topography and climate regimes (e.g., Chandler, 1962; Oke, 1973; Ackerman, 1985; Moreno García, 1994, among others). Many causes contribute to the urban heat island effect. Mitchell (1961) emphasizes the role of heat capacity and conductivity of building and paving materials: cities can absorb larger amounts of heat than rural soils during the day, which then becomes available at night to partially balance the nocturnal radiation loss. Oke (1982) lists many factors contributing to the urban heat island: increased absorption of short-wave radiation due to canyon geometry, increased long-wave radiation from the sky due to air pollution, decreased long-wave radiation loss because of the reduction of the sky view factor, anthropogenic heat sources, increased sensible heat storage and decreased evapotranspiration due to construction materials and decreased total turbulent heat transport due to wind speed reduction caused by canyon geometry.

Due to the complexity of including all of the factors and the lack of data for this purpose, population is the parameter most frequently chosen to represent the level of urbanization of a city (e.g., Mitchell, 1953; Oke, 1973, 1979, 1982; Colacino and Rovelli, 1983; Karl et al., 1988b). Although it is not a physical quantity, it is the most well documented urban parameter for both long periods of time and cities of different sizes.

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A method of quantifying the urban heat island effect is to compare the urban temperature record with a neighboring rural station. Karl et al. (1988b) demonstrated that the annual urban bias effect is a nonlinear function of population:

$$\Delta T_{u-r} = a(\mathrm{POP})^b, \tag{1}$$

where ΔT_{u-r} is the yearly mean temperature difference (°C) between a station located in an urban area and a rural station, POP is the urban population and a and b are constants. As stated by the authors, Equation (1) adjusts in statistical terms but may have substantial errors in some cases.

It has been suggested (Kukla et al., 1986; Wood, 1988) that a proportion of the long-term warming trends of the last hundred years in many global and hemispheric mean temperature records (Jones et al., 1986a, b; Hansen and Lebedeff, 1987; Jones, 1988) could be partly related to urbanization influences since urban growth enhances the heat island effect. Karl et al. (1988b) analyzed the effect of urbanization in U.S. temperature data derived from the Historical Climate Network (HCN) (Quinlan et al., 1987) for the period 1901–1984 and found that it amounts to 0.06 °C for the annual mean temperature. Jones et al. (1989) identified an urbanization bias in the Jones et al. (1985, 1986a) data for the United States of 0.1 ± 0.05 °C over the period 1901–1984. Generalizing from this, they concluded that the Northern Hemisphere land mass temperature average compiled by Jones et al. (1986a) might contain an urban induced warming trend of, at the maximum, 0.1 °C for the same period, which is about one-fifth of the global land-based temperature increase of 0.5 °C over the same time period. Jones et al. (1990) examined a set of rural station temperature data for European parts of the Soviet Union, eastern Australia and eastern China to assess the urbanization influence in different hemispheric data sets (Jones et al., 1986a, c; Vinnikov et al., 1990). In none of these regions did they find any significant urban influence on regional temperature trends, so they considered that the results of Jones et al. (1989) represent an upper limit to the urban influence on hemispheric temperature trends. While, this argument may hold for those regions, in some others, urbanization might have played a stronger role. In fact, Camilloni and Barros (1995) found that the warming trend derived from the Jones et al. (1991) data for a South American region covering most of subtropical Argentina, defined by the grid points (30° S, 60° W), (35° S, 60° S), (25° S, 70° S) and (30° S, 70° W), has an exaggerated warming of 0.5 °C/100 years for the period 1895–1988. When urban temperature records are corrected, the regional temperature trend is 0.2 °C/100 years, less than half of the trend estimated with the Jones et al. (1991) data. The difference between these trends is due to the lack of identification of Argentine urban series in the Jones et al. (1991) data and strengthens the suggestion of Wigley and Jones (1988) and Karl and Jones (1989) about the importance of detailed regional studies for finding a proper estimate of the magnitude of the urban bias in global land-based temperature trends.

Moreover, it seems that population growth is not the only aspect that must be taken into account in correcting temperature series by urbanization effects.

Barros and Camilloni (1994) found that, although Buenos Aires' Metropolitan Area population had a persistent increase since the beginning of the twentieth century, Buenos Aires' ΔT_{u-r} presents a slow decrease after the 1960s. They found that the correlation between yearly mean rural temperature (T_r) and ΔT_{u-r} is significantly negative, indicating that warmer years were associated with lower ΔT_{u-r}. It seems unlikely that air pollution changes played a significant role in this phenomenon as the mixing layer height over Buenos Aires is higher than 700 m most of the time (Scian and Quinteros, 1975). The explanation is probably related to an easier vertical dissipation of heat caused by a greater frequency of unstable conditions during warmer years (Barros and Camilloni, 1994).

Temperature records corrected by urbanization have been obtained as a difference between the observed urban temperature (T_u) and ΔT_{u-r}, usually estimated as a function of population growth (e.g., Karl et al., 1988b). If the negative correlation observed in Buenos Aires between ΔT_{u-r} and T_r were also observed in most of the cities of the world, this behavior could be contributing to mask the possible global warming presently under way due to an overestimation of ΔT_{u-r} for warming periods. In other words, we would be in the presence of another type of urban bias effect not depending on urban growth but on yearly mean temperature itself. So, the urban heat island correction factors depending on population growth might be leading to an underestimation of regional, hemispheric or global temperature trends due to an overestimation of ΔT_{u-r} during warming periods.

To check if lower ΔT_{u-r} are associated with warmer years in other cities of the world, the correlation between ΔT_{u-r} and T_r for other Argentine locations and for Australian and U.S. cities is explored. The impact of T_r trends in ΔT_{u-r} trends, and consequently in the estimate of regional and hemispheric temperature trends, is also assessed.

2. Data

Identification of urban/rural station pairs is a difficult task when the history of the meteorological stations is not available or there is little direct knowledge of the records. In this study only three countries were considered (United States, Argentina and Australia) because of the difficulty in selecting appropriate station pairs in other parts of the globe.

Nine Argentine urban/rural station pairs were selected according to the information provided by the National Meteorological Service (Figure 1). In Argentina, there are very few urban/rural station pairs available to calculate the urban effect, mainly because there were only few urban stations. Worse yet, in most of the cases when there was an urban record, there was not a simultaneous rural record in a radius of 200 km. So, in those cases, when the rural temperature is not available, it is estimated using the geographic model proposed by Barros and Camilloni (1994) valid for northeastern Argentina. The hypothesis on which the model is based is that in a region with smooth horizontal gradients in surface properties, T_r can be described by a simple function of latitude (lat), longitude (lon) and elevation (h):

$$T_r = T_r \text{ (lat, lon, h)} + e \quad (2)$$

where e is a departure from the model temperature.

The model is polynomial and includes linear, quadratic and product terms of the lat, lon and h variables. For each year of the period 1929–1991, variables are selected through the stepwise regression procedure (Draper and Smith, 1966). According to the results presented by Barros and Camilloni (1994), the yearly mean rural temperature at a location of the studied region can be estimated with a RMSE (root mean square error) of less than 0.5 °C as long as data from at least eight stations of the region are available. Generally, for every year there are much more data than eight records and hence the RMSE is much lower. The reason that this model works quite well in northeastern Argentina is that the geographical features are quite homogeneous, i.e., flat terrain with land mostly tilled or devoted to cattle raising. As an example of the model results, Figure 2 shows the yearly mean temperature series for the period 1950–1991 at Ezeiza, 30 km southwest of the Buenos Aires' downtown area, derived from observations at the international airport and the temperature for the same location derived from the model. There is a general agreement between the observed and model temperatures, but to estimate Buenos Aires urban heat island intensity it is better to use the model as Ezeiza is suspected to have also some minor urban warming. In fact, Ezeiza temperature series has a positive trend of 0.04 °C/year since 1965 while the model shows only a 0.02 °C/year trend since that year which is more consistent with the observed regional trends (Barros and Camilloni, 1994). The geographic model was used to estimate the rural temperature at Buenos Aires, Rosario, Concordia and Goya for the whole period and for 48% of the years at Corrientes (period 1934–1958).

The Australian temperature data were extracted from the World Monthly Surface Station Climatology (NCAR, 1992). Six urban/rural station pairs were considered following Coughlan et al. (1989) for four of the larger Australian cities (Figure 3). Data for the United States stations came from the HCN (Quinlan et al., 1987). Urban/rural station pairs were selected after Kukla et al. (1986) and Karl et al. (1988b). Thirty-one pairs spread over the U.S. territory were chosen, in order to have an approximated homogeneous coverage (Figure 4).

In every pair selected, the altitude difference between the urban and the rural station was less than 150 meters in order to preclude the altitudinal difference as an additional factor in the analysis.

3. Linear Correlation Between ΔT_{u-r} and T_r

Table I shows the correlation coefficient (R) between ΔT_{u-r} and T_r. In almost every case, correlation coefficients are negative. The only exception is Córdoba

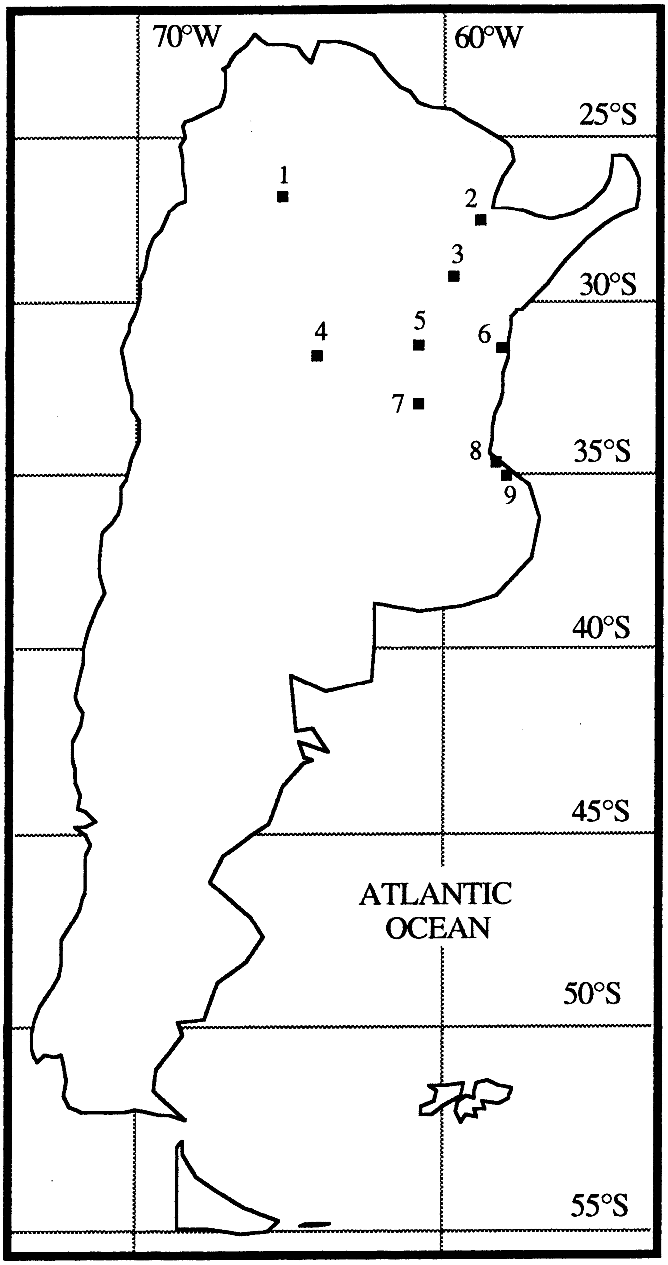

Figure 1. Argentine cities considered in the study of the rural warming influence on the urban heat island effect. (1: Tucumán; 2: Corrientes; 3: Goya; 4: Córdoba; 5: Santa Fe; 6: Concordia; 7: Rosario; 8: Buenos Aires; 9: La Plata).

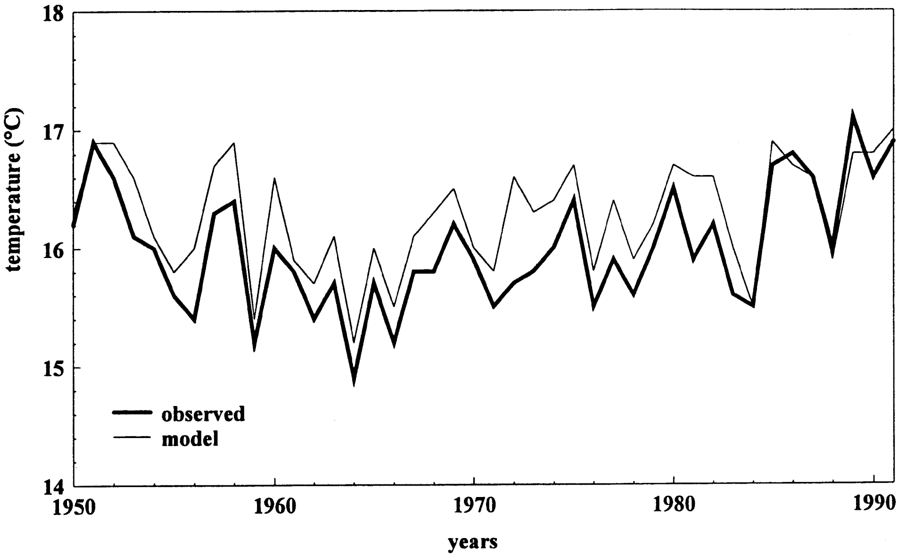

Figure 2. Yearly mean temperature series for Ezeiza.

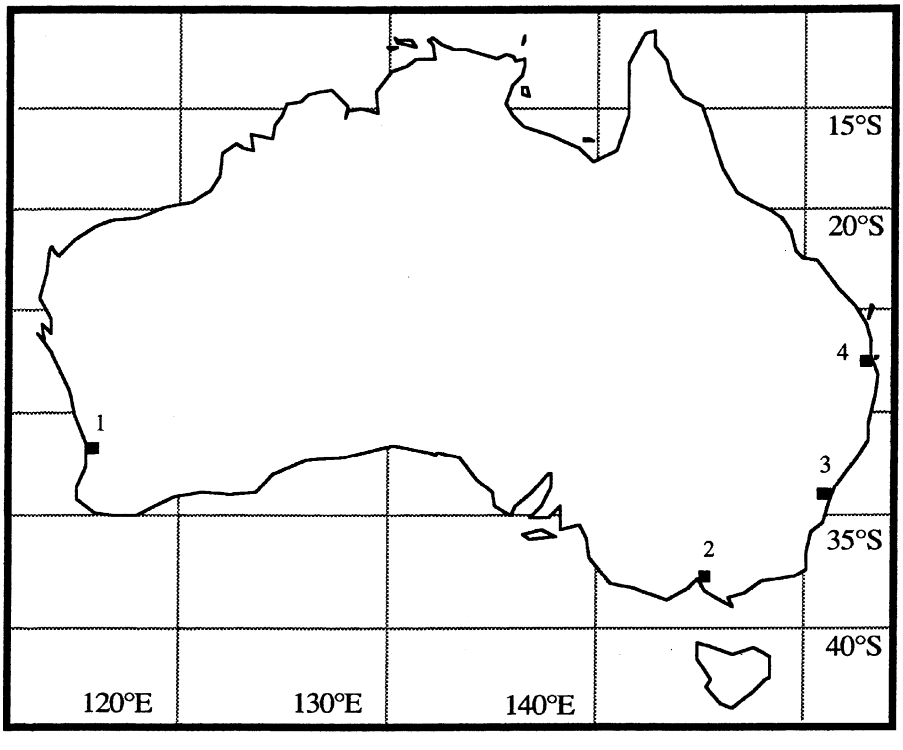

Figure 3. As Figure 1 for Australian cities (1: Perth; 2: Melbourne; 3: Sidney, 4: Brisbane).

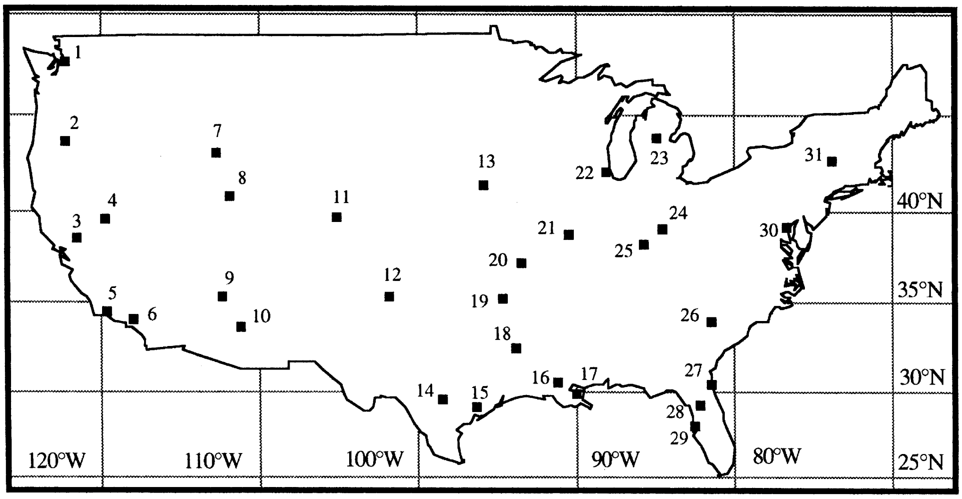

Figure 4. As Figure 1 for American cities. (1: Seattle; 2: Portland; 3: Sacramento; 4: Reno; 5: Ojai; 6: Los Angeles; 7: Aberdeen; 8: Salt Lake City; 9: Williams; 10: Roosvelt; 11: Denver; 12: Amarillo; 13: Omaha; 14: San Antonio; 15: Danevang; 16: Baton Rouge; 17: New Orleans; 18: Shreveport; 19: Fort Smith; 20: Springfield; 21: St. Louis; 22: Chicago; 23: Mount Pleasent; 24: Cincinnati; 25: Louisville; 26: Columbia; 27: Fernandina Beach; 28: Tampa; 29: Ocala; 30: Baltimore; 31: Albany).

(Argentina) and in Tucumán (Argentina) where there is no correlation at all. The negative correlations have high levels of statistical significance (58% of them has at least a 95% confidence level) and suggest an inverse relationship between T_r and ΔT_{u-r}.

Since ΔT_{u-r} is calculated as the difference between urban and rural temperature, it might be argued that the negative correlation should be a simple mathematical consequence of that. This would be true, if urban and rural yearly mean temperatures were completely uncorrelated, but this is not the case, Table I. On the other hand, ΔT_{u-r} is not a mere mathematical difference of temperature: it has a physical identity as it is the result of a number of physical processes which have a different quantitative output in urban and rural environments. Therefore, the negative correlation between ΔT_{u-r} and T_r should not be viewed as a simple mathematical artefact. To further explore this aspect, the linear correlation coefficient (R) between ΔT_{u-r} and T_r can be expressed as:

$$R = (\sigma_u/\sigma_{\Delta T_{u-r}})(R_{ur} - \sigma_r/\sigma_u), \tag{3}$$

where σ_u and σ_r are respectively the standard deviations of the yearly mean urban and rural temperatures, $\sigma_{\Delta T_{u-r}}$ is the standard deviation of the urban to rural temperature differences and R_{ur} is the linear correlation coefficient between the yearly mean urban and rural temperatures. The sign of R depends on the difference ($R_{ur} - \sigma_r/\sigma_u$). In the majority of the cases σ_u is lower than or equal to σ_r, Table I, and so, no matter the value of R_{ur}, R results negative. This is true in nearly 2/3 of the cases indicating that year-to-year temperature variability in an

Table I

Urban/rural station pairs. N is the number of data, σ_u and σ_r are the standard deviations of T_u and T_r respectively, R_{ur} is the linear correlation coefficient between T_u and T_r, and R is the linear correlation coefficient between ΔT_{u-r} and T_r with its respective significance level (α)

City	Urban station	Rural station	Period	N	R_{ur}	σ_u	σ_r	σ_r/σ_u	R	α (%)
ARGENTINA										
Buenos Aires	Buenos Aires Observatory	Buenos Aires[a]	1929–1991	63	0.36	0.46	0.55	1.20	–0.50	99
La Plata	La Plata Observatory	La Plata Airport	1968–1984	17	0.93	0.34	0.36	1.06	–0.30	80
Córdoba	Córdoba Observatory	Córdoba Airport	1953–1991	39	0.97	0.83	0.78	0.94	+0.18	70
Rosario	Rosario	Rosario[a]	1934–1948	15	0.94	0.53	0.68	1.28	–0.70	95
Corrientes	Corrientes	Corrientes[a]	1934–1958	44	0.60	0.48	0.48	1.00	–0.30	95
		Corrientes Airport	1962–1988							
Concordia	Concordia	Concordia[a]	1934–1964	25	0.80	0.46	0.47	1.02	–0.32	90
Tucumán	Tucumán Observatory	Tucumán Airport	1957–1975	18	0.85	0.44	0.37	0.84	0.02	–
Santa Fe	Inmaculada Concepción	Sauce Viejo Airport	1955–1966	11	0.93	0.60	0.60	1.00	–0.15	60
Goya	Goya	Goya[a]	1929–1975	30	0.88	0.50	0.51	1.02	–0.25	90
AUSTRALIA										
Melbourne	Melbourne Reg. Office	Laverton	1961–1981	21	0.91	0.35	0.44	1.26	–0.62	99
Melbourne	Melbourne Reg. Office	East Sale	1961–1981	21	0.78	0.35	0.28	0.80	–0.10	60
Sidney	Sidney	Richmond	1961–1970	10	0.73	0.23	0.32	1.39	–0.72	95
Sidney	Sidney	Nowra	1961–1970	10	0.87	0.23	0.31	1.35	–0.71	95
Brisbane	Brisbane	Amberley	1951–1970	17	0.79	0.40	0.44	1.10	–0.43	95
Perth	Perth	Perth Airport	1961–1970	10	0.95	0.51	0.54	1.06	–0.42	80
UNITED STATES										
Amarillo	Amarillo	Vega	1924–1977	54	0.79	0.83	0.87	1.05	–0.40	99
Baltimore	Baltimore	Woodstock	1910–1984	66	0.51	0.75	0.67	0.89	–0.40	99
Cincinnati	Cinicnnati	Cambridge	1910–1980	70	0.73	0.84	0.96	1.14	–0.51	99
Denver	Denver	Cheesman	1910–1977	68	0.66	0.83	0.76	0.92	–0.32	99
Ft. Smith	Ft. Smith	Ozark	1922–1978	59	0.68	0.73	0.55	0.75	–0.11	75
Louisville	Louisville	Shelbyville	1910–1974	65	0.20	0.74	0.99	1.34	–0.75	99
New Orleans	New Orleans	Houma	1911–1984	45	0.35	0.66	0.69	1.04	–0.59	99
Omaha	Omaha	Logan	1910–1980	67	0.88	0.88	0.96	1.09	–0.41	99
Portland	Portland	Lewiston	1910–1984	75	0.48	0.69	0.71	1.03	–0.53	99
Reno	Reno	Fallon	1910–1984	75	0.44	0.79	0.71	0.90	–0.46	99
Sacramento	Sacramento	Colfax	1910–1968	55	0.64	0.56	0.92	1.64	–0.58	99
San Antonio	San Antonio	Blanco	1910–1984	67	0.79	0.64	0.76	1.19	–0.53	99
Shreveport	Shreveport	Plain Dealing	1910–1980	66	0.82	0.66	0.91	1.38	–0.66	99
Springfield	Springfield	Lockwood	1910–1984	68	0.86	0.79	0.84	1.06	–0.38	99
St. Louis	St. Louis	L.A. Starks	1910–1983	67	0.78	0.83	0.82	0.99	–0.30	99
Tampa	Tampa	St. Leo	1937–1970	34	0.89	0.52	0.50	0.96	–0.16	75
Salt Lake City	Salt Lake City	Morgan	1910–1980	69	0.81	0.72	0.98	1.36	–0.42	99
Los Angeles	Los Angeles	San Bernardino	1910–1976	66	0.84	0.75	0.65	0.84	–0.04	60
Albany	Albany	Tweed	1910–1970	61	0.86	0.84	0.61	0.73	–0.71	99
Chicago	Chicago	Lansing	1910–1980	71	0.89	0.86	0.79	0.92	–0.07	75
Seattle	Seattle	Buckley	1914–1980	66	0.76	0.77	0.67	0.87	–0.14	80
Aberdeen	Aberdeen	Melette	1917–1987	27	0.91	1.04	1.09	1.05	–0.32	90
Baton Rouge	Baton Rouge	Bunkie	1945–1987	18	0.90	0.64	0.60	0.94	–0.11	60
Williams	Williams	Fort Valley	1910–1986	41	0.36	0.77	0.57	0.74	–0.39	99
Roosvelt	Roosvelt	Childs	1926–1987	22	0.85	0.59	0.71	1.20	–0.54	99
Ocala	Ocala	Inverness	1932–1964	11	0.85	0.49	0.51	1.04	–0.36	80
Ojai	Ojai	Sta. Barbara	1910–1958	36	0.83	0.59	0.62	1.05	–0.38	95
Mount Pleasent	Mount Pleasent	Hart	1921–1984	15	0.95	0.67	0.69	1.03	–0.26	80
Fernandina Beach	Fernandina Beach	Federal Point	1910–1987	12	0.86	0.76	0.66	0.87	–0.02	–
Columbia	Columbia	Newberry	1910–1986	46	0.85	0.61	0.53	0.87	–0.06	60
Danevang	Danevang	Halletsville	1932–1987	14	0.94	0.53	0.59	1.11	–0.44	90

[a] Indicate that yearly values were calculated with the geographic model.

urban environment tends to be lower than in its rural surroundings contributing to a negative correlation between T_r and ΔT_{u-r}.

An explanation for this negative correlation may be quantitative changes in meteorological factors associated to changes in mean temperature because the urban heat island depends, for instance, on wind, clouds and near-surface temperature lapse-rate (Sundborg, 1950; Chandler, 1965). Although other meteorological conditions like high winds, rain and snow storms may contribute to dissipate or reduce the urban heat island, the statistical results presented in this paper could be attributed to the relationship between urban heat island intensity and near-surface temperature lapse-rate. Several authors (Ludwig and Kealoha, 1968; Lee, 1975; Godowitch, 1985) found that the near-surface lapse-rate is highly correlated with ΔT_{u-r}. Their results are consistent with the fact that the urban heat island intensity is greater on the minimum than on the mean temperature and is almost zero for the maximum temperature. Urban warming is greater in cases of higher vertical stability usually associated with cooling processes. On the contrary, with unstable conditions and a higher mixing layer, the urban heat is easily dissipated vertically (Holzworth, 1974; Mazzeo and Gassmann, 1990) and so, the urban heat island is minimal or disappears. During warm years an increase in the maximum or in the minimum temperatures, or in both, can be expected. In the first case, an increase of the frequency of the unstable conditions during the day hours is probable. In the second, the frequency and intensity of very stable conditions during night and morning hours would decrease. Therefore, in any case, a reduction of the heat island effect can be expected. Similar reasoning in the opposite sense can explain an increase of the heat island effect during cool years.

4. Urban/Rural Temperature Difference Dependency on Temperature

If the physical explanation given in Section 3 were true, there would be no reasons for the negative correlation between ΔT_{u-r} and T_r not to hold at different time-frequencies. In particular, it would induce opposite trends in both variables. To check if this is the case, all positive and negative trends in the rural records presented in Table I which are significant at 80% level for at least 20 years, were analyzed. The selection of the elapsed period for each trend calculation was objective, starting with the first period of 20 years that either has a positive or a negative trend with 80% significance level and then adding the following years until the resulting trend became non-significant. Not all of the station pairs presented in Table I qualified for this analysis. Some did not because they had simultaneous records shorter than 20 years, and others because the simultaneous records were quite discontinuous with time, as can be deduced from inspection of the first and second columns of Table I.

The significance level used in the selection criterion is certainly low but it is hard to find numerous trends with higher significance levels (e.g., Kukla et al., 1986;

Table II

ΔT_{u-r} and T_r trends for warming and cooling periods. Significant ΔT_{u-r} trends are indicated with ([a]). All T_r trends are significant

City	Warming period			Cooling period		
	Rural trend (°/year)	Urban/rural temperature difference trend (°C/year)	Period	Rural trend (°/year)	Urban/rural temperature difference trend(°C/year)	Period
ARGENTINA						
Buenos Aires	0.03	–0.01	1959–1983	–0.02	+0.04[a]	1929–1950
Corrientes	0.02	+0.01	1963–1982	–0.05	+0.03	1942–1963
Concordia	0.03	+0.01	1935–1954			
Córdoba	0.02	+0.01[a]	1954–1987	–0.05	+0.00	1971–1991
Goya	0.04	+0.01[a]	1934–1954	–0.06	+0.03[a]	1954–1975
AUSTRALIA						
Melbourne – East Sale	0.02	–0.08	1961–1981			
Melbourne – Laverton	0.04	–0.02[a]	1961–1981			
UNITED STATES						
Amarillo	0.06	–0.03[a]	1924–1955	–0.05	+0.04[a]	1955–1976
Baltimore	0.03	+0.01	1933–1955	–0.06	+0.08[a]	1910–1930
Cincinnati	0.06	–0.01	1912–1936	–0.05	+0.03[a]	1931–1968
Denver	0.05	+0.02[a]	1915–1940	–0.06	+0.03[a]	1948–1977
Louisville	0.05	–0.04	1911–1933	–0.09	+0.05[a]	1941–1966
Omaha	0.04	–0.04[a]	1949–1970	–0.06	+0.03[a]	1930–1949
Portland	0.04	–0.03[a]	1912–1946	–0.02	–0.01	1946–1969
Reno	0.04	–0.08[a]	1939–1963	–0.04	+0.07[a]	1963–1984
Sacramento	0.05	–0.01	1914–1935	–0.04	–0.02	1934–1953
San Antonio				–0.06	+0.04[a]	1931–1962
Shreveport	0.02	+0.01[a]	1915–1936	–0.07	+0.04[a]	1951–1980
Springfield	0.05	–0.05[a]	1922–1946	–0.04	+0.03[a]	1950–1974
Tampa				–0.04	+0.01[a]	1951–1970
Salt Lake City	0.07	–0.02[a]	1916–1938	–0.03	+0.03[a]	1956–1976
Los Angeles	0.03	+0.03[a]	1946–1968	–0.04	+0.01	1925–1946
Albany	0.04	–0.02[a]	1912–1937	–0.05	+0.02[a]	1944–1970
Chicago	0.04	–0.03[a]	1934–1954	–0.05	+0.03[a]	1954–1975
Seattle	0.04	+0.01	1950–1970	–0.03	+0.04[a]	1930–1950
Ojai	0.03	–0.01	1910–1930			
Columbia				–0.05	+0.02[a]	1939–1958

Hansen and Lebedeff, 1987). Results presented in Table II show that the 80% confidence level leads to selection of trends higher than or equal to, in absolute value, 0.02°/year except for one case in Australia. This means at least a 0.4 °C difference in 20 years and in many cases more than that since some trends were calculated over longer periods of time.

Buenos Aires rural temperature presents a cooling trend during 1929–1950 and a warming trend during 1959–1983. The other rural stations of Argentina had cooling and warming periods which do not show a coherent regional picture. This reflects the lack of important trends in subtropical Argentina, as reported by Camilloni and Barros (1995). The length of the records allows the analysis of only one case in Australia. In the American case there is a tendency for the warming periods to be concentrated during the first part of the century and the cooling periods after 1930 and before 1970. This is in agreement with regional trends (Karl and Jones, 1989). Table II includes the ΔT_{u-r} trends which are generally opposite in sign to T_r trends and even more when they are significant at 80% level of confidence. In fact, 35 out from 45 pairs i.e., 78% of the cases show opposite signs. In those cases when the ΔT_{u-r} had a significant trend at the 80% level of confidence, this percentage increases to 87%. During cooling periods there is not a single case in which there was a significant cooling trend in ΔT_{u-r}. Over periods of more than 20 years, urban population growth tends to increase the urban heat island, but this increase calculated from Equation (1) (Karl et al., 1988b) would amount to around 0.01°/year for most of the cities of Table II. Therefore, in only one of the cooling cases, i.e., in Tampa, the urban growth could have changed the sign of the ΔT_{u-r} trend. Most of the cases with the same sign in the trends belong to warming periods, i.e., 8 cases, but in 6 of these, the ΔT_{u-r} trend is 0.01°/year or lower (some cases appear with this value in Table II because of round off). This trend could be caused by urban growth. So, in absence of the urban growth effect, most of these cases would probably show trends of opposite sign in both parameters.

ΔT_{u-r} trends shown in Table II were corrected later, subtracting the urban growth bias calculated with Equation (1). Different coefficients were used according to national studies. For American data, following Karl et al. (1988b) the values are $a = 0.00182$ and $b = 0.45$. We used $a = 0.00378$ and $b = 0.38$ for Argentina (Camilloni and Barros, 1995) and $a = 0.0113$ and $b = 0.30$ in the case of Australia (Coughlan et al., 1989). Population data were taken from National Census and adjusted linearly for the periods shown in Table II. Figure 5 shows the scatterplot of ΔT_{u-r} corrected trends against T_r trends. As anticipated, the urban growth correction resulted in only minor changes with respect to the values of Table II. Since only significant rural temperature trends were considered, there are no rural trends with absolute values lower than 0.02°/year. There is a clear dependency of ΔT_{u-r} trends on T_r trends, that when adjusted linearly has a negative slope significant at 99% level. The slope is –0.53 with a standard error of 0.09.

Figures 6a to 6d show two examples of warming periods (Buenos Aires and Amarillo) and two of cooling periods (Melbourne and Amarillo) of T_r and their respective ΔT_{u-r}. These graphs are representative of most cases showing not only the opposite trends but also opposite variations in higher frequencies which also contribute to the negative correlation shown in Table I.

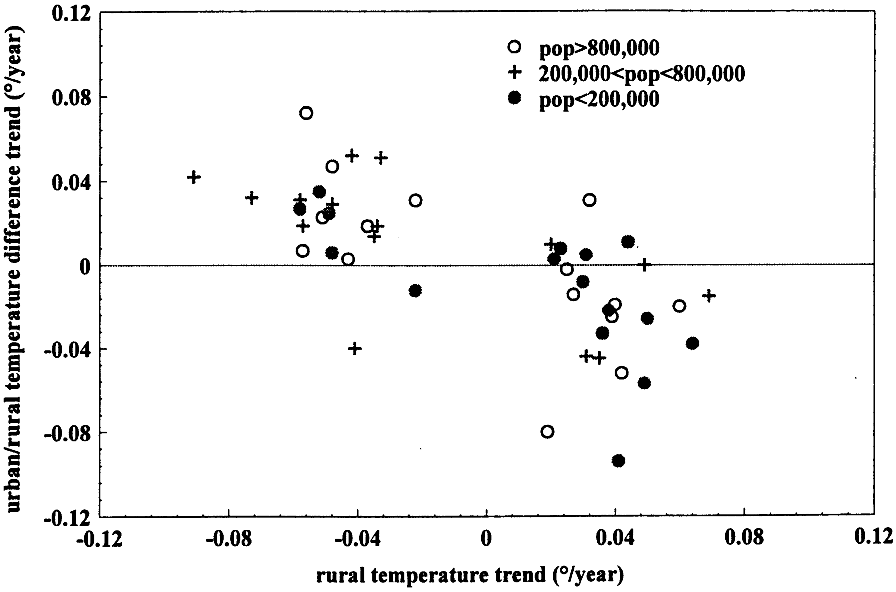

Figure 5. Urban/rural temperature difference trends (°/year) as a function of rural temperature trends (°/year) for different population ranges.

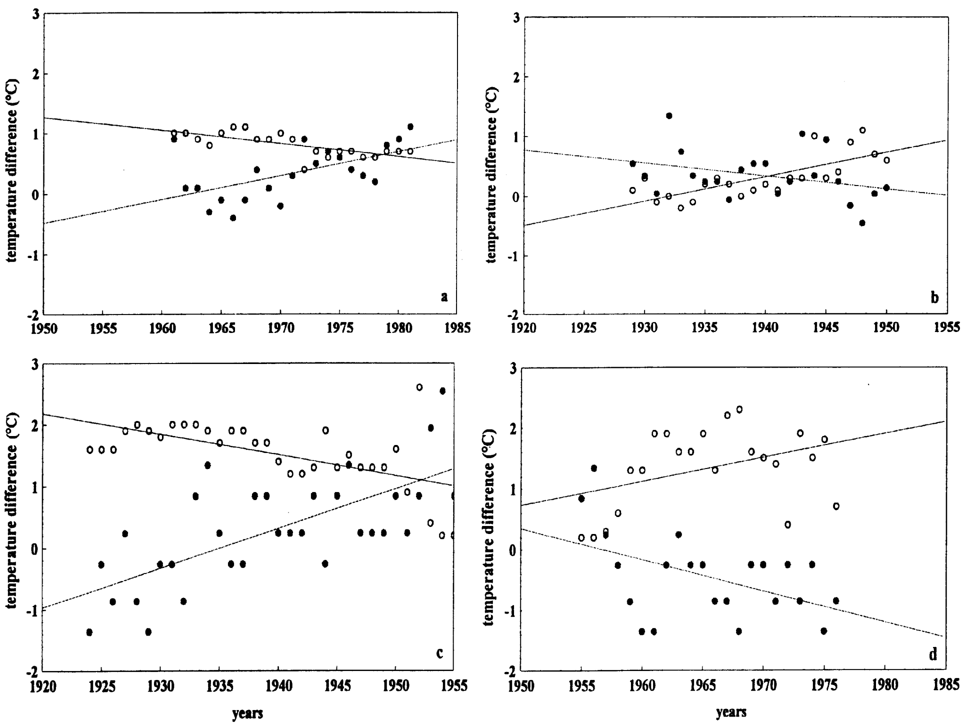

Figure 6. Examples of interannual variation of ΔT_{u-r} (○) and T_r (●) for Melbourne (a), Buenos Aires (b) and Amarillo (c, d). Adjustments to linear trends are shown.

5. Impact of the ΔT_{u-r} Dependency on Temperature on Regional and Global Trends

Results from sections 3 and 4 indicate that urban stations are prone to trends of lower absolute values than rural ones no matter if they are positive or negative. Therefore, other than the well known bias in urban temperature due to urban growth, there seems to be a second bias in urban temperature trends depending on the T_r trends. Regional and global surface temperature analysis included an important fraction of stations which are either urban or could be suspected of having urban influence (Kukla et al., 1986; Wood, 1988). For instance, in the American case, Karl et al. (1988b) using data from the HCN found that the urban effect is statistically detectable at population levels as low as 2,500 and Balling and Idso (1989) note this effect even in communities with lower populations. So, regional and global trends could be affected by this second urban bias which enhances the first one during cooling periods but has opposite sign during warming ones.

Naming ΔT_{u-r} and T_r trends as τ_{ur} and τ_r respectively, the linear adjustment of data depicted in Figure 5 can be written as:

$$\tau_{ur} = \gamma\tau_r \tag{4}$$

where $\gamma = -0.53$ and the calculated intercept is almost zero i.e., 0.002. Both τ_{ur} and τ_r are in $^\circ$C/year. As a result and due to linearity, urban trends (τ_u) can be considered in statistical terms, to be roughly half of τ_r. Of course, Equation (4) is only valid in statistical terms and not for every individual case. Another limitation to Equation (4) is that it was developed from data with trends not higher in absolute value than 0.1°/year and it cannot be valid for indefinitely large incremental values of temperature.

It is convenient to explore if this relationship depends on urban population. In Figure 5, the 45 cases are split in three population ranges so to have approximately a third of the cases in each range. The linear adjustment in each case leads to slope values of –0.59 $\pm$ 0.19 for urban population over 800,000; –0.46 $\pm$ 0.15 for cities between 150,000 and 800,000 inhabitants and –0.53 $\pm$ 0.16 for urban population under 150,000. Although there are few cases in each range to arrive to a firm conclusion, it seems that there is no evidence on urban population dependency.

Notwithstanding, a severe limitation in the analysis leading to Equation (4) is that only one out of the 45 cases is from a city with less than 10,000 inhabitants. In some regions an important fraction of data comes from stations in small communities; for instance in the subset of HCN stations used by Karl et al. (1988a), over 70% were in cities which in 1980 have less than 10,000 inhabitants. Hence, it is very important to learn what should be expected regarding the urban bias due to temperature trends in these small cities. For this reason, another approach was taken to analyze the dependency on population which may include more data from small cities using

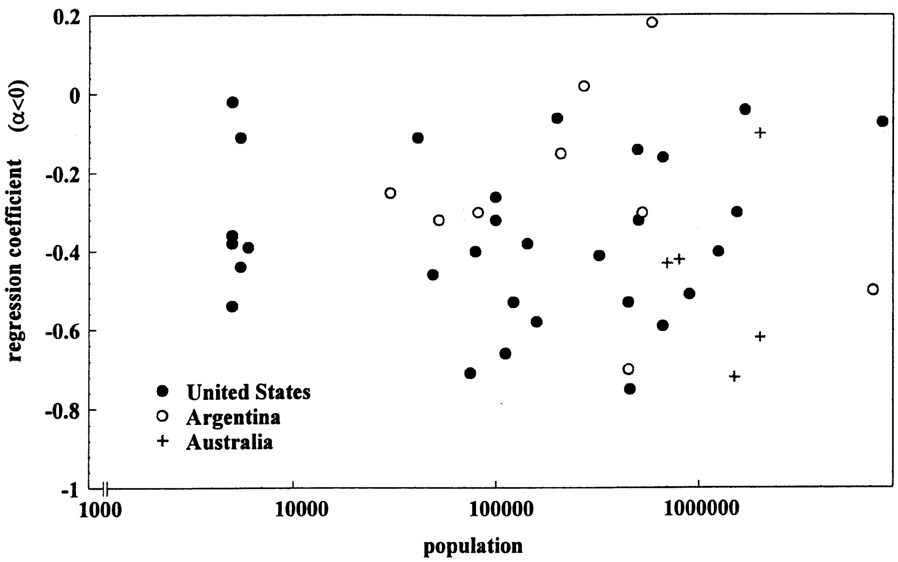

Figure 7. Regression coefficient α for cities listed in Table I.

the same data set. Therefore, linear regressions were calculated between ΔT_{u-r} and T_r:

$$\Delta T_{u-r} = \alpha T_r \tag{5}$$

where α is the regression coefficient and the intercept was forced to be zero to be consistent with Equation (4). If it is assumed that α does not depend on time then it could be considered an estimate of γ. Determination of α was done for all the stations, even for those with discontinuous records using its whole period of time, whether or not they have significant trends.

Figure 7 shows the regression coefficient α as a function of the average population. This average was carried out for each city over the period shown in Table I. From this figure, it can be assumed that this regression coefficient does not depend on urban population. Its average value, –0.37, is not inconsistent with the –0.53 of Equation (4) considering that the sample is not identical. This agreement is the consequence of the linearity in the process of trend calculation.

In this new analysis, seven cases correspond to small urban communities with less than 10,000. Their regression coefficients, α, range from –0.02 to –0.54, being to this respect similar to larger cities. Then, they can be expected to have the same urban bias in temperature trends as larger cities.

The following exercise is only intended to find a bulk approximation to the magnitude of the second urban bias. Jones et al. (1989) derived an area average time series for the Jones et al. (1986c) data for the region 30° N to 50° N, 70° W to 120° W covering most of the United States and parts of Mexico and Canada and some oceanic area referred to as the contiguous United States. The linear warming trend in this series over the period 1901–1984 is 0.31 °C with an estimated urban growth bias of 0.1 ± 0.05 °C. Considering that 0.1 °C is a correct value for the urban growth bias, then the calculated increase without it would be 0.21 °C. Assuming now that 80% of the stations have also the second urban bias in their trends according to Equation (4), i.e., its rural trend would be twice the urban one, the regional increase would be 0.36 °C. This increase is larger than the original 0.31 °C. In other words, this second bias offsets the urban growth effect and in this case seems to be even greater. However, this result should be taken with caution and only as an indication of the need for discussions about the effect of the second urban bias on regional and global temperature trends. A more accurate procedure to be developed should be to correct each data station and afterwards find the regional trend according to the algorithms used for developing regional series (Jones et al., 1986c; Vinnikov et al., 1990).

To what extent this result for the contiguous United States area could be extended to global estimates seems difficult to assess since the presence and magnitude of the second urban bias was estimated using data from three middle latitude countries. The second urban bias in temperature trends will also depend on the relative incidence of the urban bias itself in regional data sets. Regarding this point, there appears to be a variety of situations in different regions of the world. There have been some comments that heat island growth in the contiguous United States has been quite anomalously high with respect to other regions of the world (Karl and Jones, 1989). European Russia and Australian data sets do not seem to present regional urban bias in its trends except for the cooling period 1930–1987 in the European Russia data set (Jones et al., 1990). Therefore, it will be necessary to extend the present study to other regions of the world before arriving to a conclusion regarding the influence of the second type of urban bias in global trends. Nevertheless, it is clear that, at least for some regions, the second bias in urban trends could be of the same magnitude as the urban growth bias enhancing the urban growth bias during cooling periods and compensating it during warming ones.

6. Concluding Remarks

As was stated before, it remains to be confirmed if the results presented here hold for other regions. To do so, proper identification of urban/rural station pairs, a task that can be better done by people with direct knowledge of geography and data sets, is required. The goal of this paper it to encourage this work as well as to warn about a second type of urban bias in temperature trends since, in some cases, this second bias could offset or even outgrow the urban population growth bias.

The sign of the urban trend bias will result from the relative magnitude of both biases which depend in one case on the population growth rate, and in the other, on the magnitude of the temperature trend. So, it cannot be ruled out that regional, hemispheric and/or global trends estimated with data corrected only by urbanization growth will underestimate the warming trends. Given the importance of early detection of the greenhouse effect, this aspect deserves further study.

Finally, knowing that urban trends do not depend only on urban population growth but also on temperature trends, the time rate of change method for estimating urban bias appears as inadequate. It may lead to confusing results because temperature trends impact on urban temperature trends can be another source of error in addition to those already pointed out by Lowry (1977) and Karl et al. (1988).

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