Younes Ben Zaied (France), Houssam Bouzgarrou (Tunisia), Nidhaleddine Ben Cheikh (France), Pascal Nguyen (France)

The nonlinear progressive water pricing policy in Tunisia: equity and efficiency

Abstract

Economic theory and recent empirical evidence show that nonlinear progressive water pricing policies are the most useful tool to reduce water demand in water stressed countries. The originality of our paper is to implement Pedroni (1999) panel cointegration tests, using databases on a breakdown of two consumption blocks (a lower and an upper block) from the Tunisian water regulator over 27 years. The results reveal that increasing block tariffs have been successful in managing scarce water in Tunisia. The authors observe that, in the long-run, proportion of subscribers in the upper water consumption block decreases when price increases, while in the lower block, which is composed essentially of low-income households characterized by inelastic water demand, proportion of subscribers is less elastic to price changes and still unchanged. This paper calls for the implementation of nonlinear progressive pricing to reduce demand by large consumers in order to promote efficiency in use and to promote the access of poor consumers to the resource in order to promote equity.

Keywords: water pricing, two-part tariffs, Tunisia, panel cointegration.
JEL Classification: Q25, Q56.
Received on: 7th of April, 2017.
Accepted on: 3rd of May, 2017.

Introduction

Since water is not traded in markets, prices are not expected to adjust automatically to reflect periods of scarcity as they do for other goods and services. Instead, water pricing is usually regulated by public institutions – city councils, agencies, regulators and other entities. Given the public benefits provided by many aspects of water supply and management, the price-setting public institutions should be able in some way to measure the true economic value of water supply and to use this information to establish economically rational water tariffs. Such an issue is particularly important in water-scarce countries in which the price of water does not reflect scarcity, often because management institutions are reluctant to raise prices.

Basically, the Tunisian state water distribution company has concentrated constantly on adjusting water supply to meet level-price water demand. The cost of supply enhancement continues to rise, as the most accessible sources of water are tapped to capacity or depleted, necessitating tariff changes that subsequently affect the quantity demanded. Econometric estimates of residential demand try to define water management policies that fail to consider the time-path of adjustment risk outpacing consumers' ability to develop new habits or optimize their stocks of water-associated capital, such as landscaping, plumbing fixtures and appliances.

Many issues are raised in terms of equity: in industrialized countries, for example, consumption is more linked to the size of the household than to its financial resources (for a debate about France, see Porcher, 2014), while in developing countries water demand tends to be more closely related to the household financial constraints (see Nauges and Whittington, 2010). Demand-driven solutions are increasingly viewed as a necessary complement to, or even substitute for, supply-oriented policy measures.

The empirical estimation of water price elasticity has been a major issue in applied economics research during the last five decades. Indeed, several literature reviews (Arbuès et al., 2003; Dalhuisen et al., 2003; Worthington and Hoffman, 2008) was demonstrated that households in industrialized countries are not affected by the water tariff progressivity. The water demand literature strand was interested in industrialized countries, while very few studies such as Nauges and Whittington (2010) have considered some developing countries in studying the main determinants of residential water demand. Nauges and Thomas (2003) estimate a dynamic panel data model for a sample of French municipalities and show evidence of short- and long-run price elasticities, respectively, equal to -0.26 and -0.40. Using time-series observations from

Younes Ben Zaied, ESDES, The Business School of UCLy, Lyon, France.
Houssam Bouzgarrou, University of Sousse - ISG, Sousse, Tunisia.
Nidhaleddine Ben Cheikh, ESSCA School of Management, Angers, France.
Pascal Nguyen, ESDES, The Business School of UCLy, Lyon, France.

Seville in Spain, Martinez-Espineira (2007) documents a long-run price elasticity equal to -0.5 from a cointegration model and a short-run price elasticity equal to -0.1 from an error-correction specification.

This paper discusses water pricing options to promote efficiency and equity in use. The nonlinear progressive water pricing policy must satisfy two objectives: first, it should not reduce the small consumer's welfare by imposing relatively low water price to this kind of households. Secondly, the application of appropriate pricing to large consumers should lead to reduction of water consumption and better conservation of this scarce and precious resource by high income households.

Our research focuses on Tunisia, which has very limited water resources. According to the United Nation Development Program[1], the per capita renewable internal freshwater resource in Tunisia, in 2008, was about 406 cubic meters. The amount of renewable fresh water available per inhabitant is more than 50% below the water scarcity standard (1000 cubic meters per capita). However, annual rainfall shortages and the increase of annual average temperature have aggravated the situation. In addition, water supply suffers from several problems, as Tunisian water resources are characterized by bad quality and spatial heterogeneity of its location. On the one hand, the remoteness of water resources from urban areas increases its mobilization cost. On the other hand, the high level of water salinity increases its treatment and distribution cost. Moreover, the limited water supply is unequally distributed across the country and intensively used. This has resulted in serious challenges such as increased degradation and risk of depletion.

In the present paper, we focus on the long-run effect of nonlinear water pricing policy on the distribution of subscribers across consumption blocks. To the best of our knowledge, this is the first attempt to explore the long-run relationship between water demand determinants such as price and climatic factors and the share of subscribers in each block as an indicator of nonlinear pricing policy success. The first step of our study involved analyzing the data and carrying out the necessary tests to see whether the data are stationary. We, then, used the panel cointegration technique, which explicitly integrates the non-stationary character of our panel data, to derive the estimates of the long-run effects of water price and climatic factors with the right properties. Our findings led us to propose relevant policies recommendations.

[1]The Arab Statistics are obtained from the website of UNDP: http://www.arabstats.org/indicator.asp?ind=273.

The remainder of the paper is organized as follows. Section 1 describes the database and presents the model. The econometric method is presented in section 2. The empirical investigation and a discussion of the main results are presented in section 3. Finally, a number of policy recommendations conclude the paper.

1. Data and model

1.1. Data description. To perform the empirical investigation, we use data collected by SONEDE, the national Tunisian water utility. A panel of six heterogeneous regions was constructed from municipality level data. As we propose to investigate the main reasons of consumer movement from one water consumption block to another one, we explore regional quarterly dataset that describes water tariff structure, socioeconomic factors and climatic variables for different Tunisian regions for a long period of time (1980-2007).

We get from this dataset variables related to water consumption such as the average water price (the total bill divided by the consumed volume) and the number of consumers per region in each tier of the tariff (see Table 1). We aggregate the data into five blocks corresponding to those used for the five different tariff rates. The unit price per cubic meter per quarter is defined in Tunisian dinars (one dinar is roughly equal to 0.4 euro) and is reported for the five parts of the tariff. Fixed charges, which depend on the pipe diameter, are presented in Table 2 for each kind of pipe diameter. Our second source of data comes from the national institute of meteorology. We collected average regional rainfall and temperature by quarter, as these variables can impact water consumption, consumer behavior and, thus, scarcity.

Table 1. Tariff of water distribution in Tunisia in 2007, in Tunisian dinars per m^3

Consumption per quarter	0-20 m^3	21-40 m^3	41-70 m^3	71-150 m^3	More than 150 m^3
Dinars per unit	0.14	0.24	0.3	0.545	0.84

Source: SONEDE (2007).

In this paper, our primary focus in on the share of subscribers (the proportion of consumers) in each block. This variable represents the number of subscribers in each consumption block divided by the total number of SONEDE subscribers in Tunisia. Our database covers the period from January 1980 to December 2007 in six regions. Our sample will, then, be composed of 112 quarterly observations in six regions, namely: Greater Tunis (GT), which includes the capital city Tunis with its suburbs, North-East Tunisia (NE), North-West Tunisia (NW), Central-East Tunisia (CE), Central-West Tunisia (CW), and Southern Tunisia (S). We use the database and the decomposition in two. The lower block includes the consumers of the first two brackets (0-40 m^3), while the upper block includes the last three brackets (over 41 m^3). We present in Appendix evidence of the possible decomposition into two blocks. Figure 1 and Figure 2 describe the distribution of consumers across the two levels of the tariffs per region.

Table 2. Fixed charges for water distribution in Tunisia in 2007, in Tunisian dinars per quarter

Pipe diameter	15	20	30	40	60	80	100	150
Dinars per quarter	3.30	5.83	10.74	20.57	53.46	53.46	82.81	220.67

Source: SONEDE (2007).

Looking at the total annual share of subscribers per region in the upper block (Figure 1), we see that there is a decline throughout the period for all regions of the share of consumers in this block. This is especially true for Greater Tunis, which represents the region with the most important proportion of upper block consumers. The decrease of the share of consumers in the upper block is undoubtedly the result of increasing block water tariffs, which have seen a rapid increase during the past years especially for the big consumers. The last tariff (more than 150 m^3) represents about six times the first one (0-20 m^3) and three times the second one (21-40 m^3).

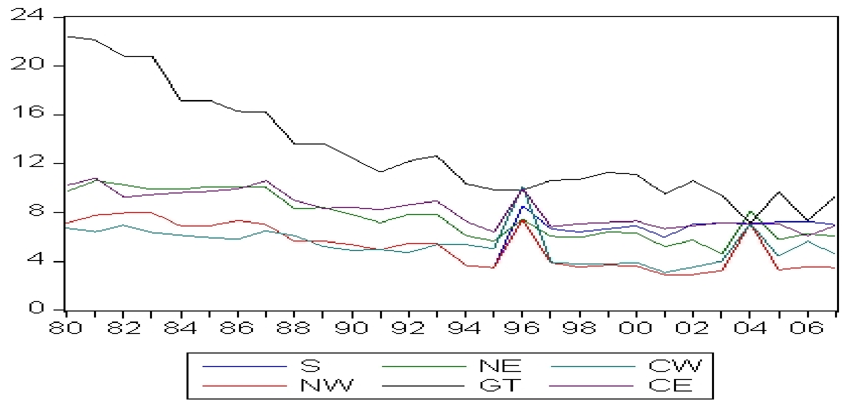

Fig. 1. Proportion of consumers in the upper block (time on x-axis, "80" stands for 1980 and "00" for 2000; % of consumers in the block on the y-axis)

However, the share of subscribers to the network (Figure 2) increases steadily in the lower block. This may predict the long-run consumer behavior. Indeed, consumers who reduce their water use, as a result of water pricing policy or changes in climate conditions, move from the upper block into the lower block. However, lower block consumers move to the upper block following changes in habits (with the concentration of holidays and wedding celebrations in summer) or increases in family size.

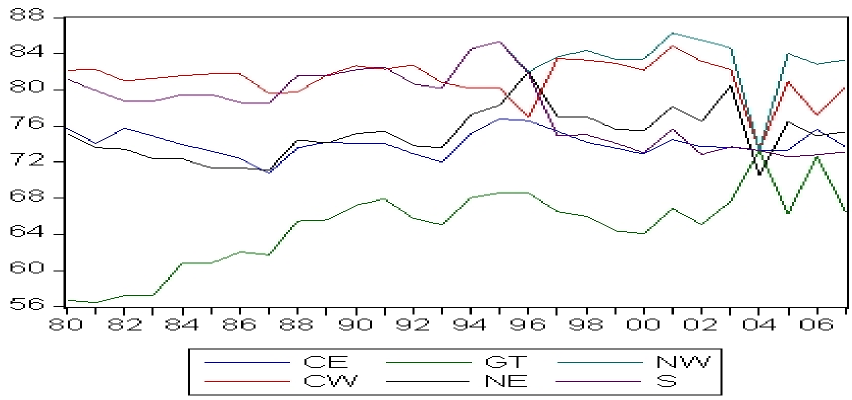

Fig. 2. Proportion of consumers in the lower block (time on x-axis, "80" stands for 1980 and "00" for 2000; % of consumers in the block on the y-axis)

1.2. Theoretical model. To assess the sensitivity of water consumers to nonlinear water tariff changes, the water demand equation is often specified as a regression of water consumption with its main determinants including socioeconomic and climatic factors (see Arbués et al., 2003). The model links household water consumption to its determinants such as price and income, as the main determinants of demand suggested by classical economic theory, followed by socio-economic factors and climatic factors (temperature and rainfall) as control variables. The main originality of this paper is not to empirically assess water price elasticity, but to analyze the movement of consumers between two consumption blocks following increasing water pricing scheme. Therefore, we use the determinants of average regional water consumption to model the proportion of consumers in each consumption block. Such equation can explain the main determinants of consumer's decision to be in the lower or in the upper block.

Figures 1 and 2 suggest that the distribution of consumers between the lower and the upper block is a symmetric one. The share of subscribers is characterized by an upward trend in the lower block and a downward in the upper block. Thus, the main focus of our model is to explain the determinants of household choice between lower and upper consumption block by estimating the long-run relationship between the proportion of consumers in each block and water consumption determinants such as income, price, total number of subscribers in Tunisia and climatic factors (temperature and rainfall). The equation is specified at regional level for each consumption block. Thus, for a period t (quarter) and individual i (region), the demand equation is specified as:

$$LnPr= \alpha_0 + \alpha_1 LnP_{it} + \alpha_2 LnI_{it} + \alpha_3 LnN_{it} + \alpha_4 LnRL_{it} + \alpha_5 LnTM_{it} + \varepsilon_{it}, \quad (1)$$

where Pr, P,I,N, RL and TM denote, respectively, the share of subscribers per quarter per region in each consumption block, average water price (the total bill divided by the volume consumed by the average consumer), average household income, total number of subscribers to the public water network, rainfall and temperature, respectively. ε_{it} is a zero mean error term normally distributed and Ln denotes the logarithmic operator used to linearize the equation so that the coefficient can be interpreted as an elasticity.

2. Empirical approach

From an empirical point of view, the use of panel dataset with long time span should check for the presence of panel unit root. If a panel unit root characterized the variables, the existence of long-run relationship between the variables should also be tested and, then, estimated to conclude for a long-run equilibrium system rather than static and traditional linear regression. To do that, econometric technics have been developed to treat such type of panel data and take care for the presence of non-stationary feature characterizing the panel dataset.

We used a panel dataset for a long period of time (112 quarter). We, then, proceed in three steps to test for unit root and panel cointegration before estimating the long-run equilibrium system using the appropriate estimation method. As we have a microeconomic panel dataset, we applied two first generation panel unit root tests such as Levin, Lin and Chu (2002) and Im, Pesaran and Shin (2003), and one second generation test proposed by Pesaran (2007), which considers cross-sectional dependence between the six regions.

We have also implemented the seven tests proposed by Pedroni (1999) to obtain the long-term relationship between all variables. Then, to estimate the long-run relationship, we use the fully-modified OLS (FMOLS) technique to estimate the cointegration vector for heterogeneous cointegrated panels, which correct the standard OLS bias induced by the endogeneity and serial correlation of explanatory variables.

The first test, which has been developed to test panel unit root, is the one of Levin, Lin and Chu (2002) (LLC (2002) hereafter). This test followed the methodology of the augmented Dickey-Fuller (ADF) unit-root test in time series. LLC (2002) tests the null hypothesis of $\delta = 0$ for all individual i, against the alternative of $\delta \neq 0$ from the following equation:

$$\Delta y_{it} = \delta y_{it-1} + \sum_{l=1}^{P_i} \theta_{ip} \Delta y_{i,t-1} + \alpha_{mi} d_{mt} + u_{it}, \quad (2)$$

where $d_{1t} = \emptyset$, $d_{2t} = \{l\}$, and $d_{3t} = \{1,t\}$ represent the different ADF specifications.

To implement the test, LLC (2002) proceed in three steps, then, the adjusted statistic used to test panel unit root is:

$$t_\delta^* = \frac{t_\delta - N \times std(\delta) \times \mu_{m\tilde{T}}^* \times \hat{\sigma}_{\tilde{\varepsilon}}^{-2} \times \hat{S}_N \times \tilde{T}}{\sigma_{m\tilde{T}}^*} \sim N(0,1)$$

with $\frac{\sqrt{N}}{T} \rightarrow 0$,

where $\hat{S}_N$, $\mu_{m\tilde{T}}^*$ and $\sigma_{m\tilde{T}}^*$ are the average standard deviation ratio calculated in the second step, the mean and standard deviation adjustments simulated by the authors for different order of m and the time series dimension $\tilde{T}$, respectively (see Levin et al., 2002).

For a specific case of LLC (2002), Im, Pesaran and Shin (2003) test (IPS (2003) hereafter) is formulated. Indeed, for m=2 and δ_i varies across cross-sectional units, the IPS (2003) statistic used to test for panel unit root is, then, formulated to test the null hypothesis of

$\delta_i = 0$ for all cross-sections i, against the alternative of $\delta_i < 0$ for $i = N_1+1,...,N$ and $\delta_i = 0$ for $i = N_1+1,...,N$.

With $N_1 \in]0, N[$, such as $\lim_{N\to\infty}(N_1/N) = \delta$, where $0 \le \delta_i \le 1$. If $N_1 = 0$, we find the null hypothesis.

The value of the IPS (2003) test statistic is calculated using the average of individual ADF statistics for all individuals and defined as:

$$\bar{t}_{NT} = \frac{1}{N}\sum_{i=1}^{N} t_{iT}(P_i\beta_i)$$

$t_{iT}(P_i\beta_i)$ is the individual student statistic under the null hypothesis for a given lag order P_i and a vector of ADF coefficients $\beta_i = (\beta_{i,1}, \beta_{i,2}, ..., \beta_{i,p_i})'$.

The IPS (2003) use the standard normal statistic Z.

$$\bar{Z} = \left[\sqrt{N}\frac{(\bar{t}_{NT} - E(t_{iT}))}{\sqrt{var(t_{iT})}} \right] \underset{N\to\infty}{\rightarrow} N(0,1)$$

$E(t_{iT})$ and $var(t_{iT})$ are the mean and variance of each statistic, respectively, and they are generated by simulations and tabulated in the IPS (1997).

As a second-generation panel unit root test, we use Pesaran (2007). The latter considers a simple dynamic linear heterogeneous panel data model:

$$y_{it} = (1-\rho_i)\mu_i + \rho y_{i,t-1} + u_{it}, \quad i = 1,...,N; t = 1,...,T, \tag{3}$$

where the error term u_{it} follow a single common-factor structure:

$$u_{it} = y_i f_t + e_{it}, \tag{4}$$

f_t is an unobserved common factor, y_i is the corresponding factor loading and e_{it} is an idiosyncratic error term independent across i and independent of the common factor. It is convenient to re-write (3) as:

$$\Delta y_{it} = \alpha_i + \beta_i y_{i,t-1} + y_i f_t + e_{it}, \tag{5}$$

where $\alpha_i = (1-\rho_i)\mu_i$, $\beta_i = -(1-\rho_i)$ and $\Delta y_{it} = y_{it} - y_{i,t-1}$. The unit root hypothesis of interest, $\rho_i = 1$, can now be expressed as:

$H_0 : \beta_i = 0, \forall i.$

Against the possibly heterogeneous alternatives:

$$H_1 : \begin{cases} \beta_i < 0 \text{ for } i = 1,2,...N_1 \\ \beta_i = 0 \text{ for } i = N_1+1,...,N \end{cases}$$

with $0 < N_1 \le N$.

In order to take care for the cross-sectional dependence induced by the common factor, Pesaran (2007) suggests to cross-sectionally augmenting the test equation (5) with cross-sectional averages of the first differences and the lagged levels. The cross-sectionally augmented Dickey-Fuller regression is, then, given by

$$\Delta y_{it} = \alpha_i + b_i y_{i,t-1} + c_i \bar{y}_{t-1} + d_i \Delta\bar{y}_t + e_{it}, \tag{6}$$

where $\bar{y}_{t-1} = \sum_{i=1}^{N} y_{i,t-1}$, $\Delta\bar{y}_t = \sum_{i=1}^{N}\Delta y_{it}$ and ε_{it} is the error term. To test the hypothesis $H_{0i} : \beta_i = 0$ for a given i, the t-statistic of b_i in (6) is called cross-sectionally augmented Dickey-Fuller ($CADF_i$). The panel unit root for the hypothesis $H_{0i} : \beta_i = 0$ for all i against the heterogeneous alternative $H_1 : \beta_I < 0$ for some i is given by the cross-sectional average of the $CADF_i$ tests, such that

$$CIPS = \frac{1}{N}\sum_{i=1}^{N} CADF_i, \tag{7}$$

CIPS was inspired from the IPS statistic (Im et al., 2003). The critical values for the test statistics based on stochastic simulations are provided in Pesaran (2007).

To test for long-run relationship between the non-stationary variables integrated in the same order, we use the residual based approach of Pedroni (1999). Following this approach the cointegration rank is a priori known and equal to one. Thus, to test for the null of no cointegration in heterogeneous panels with multiple regressors, Pedroni (1999) considers the following regression:

$$y_{it} = \alpha_i + \delta_i t + \beta_{1i}x_{1,it} + \beta_{2i}x_{2,it} + ... + \beta_{Mi}x_{M,it}\varepsilon_{it}, \tag{8}$$

where $i = 1,...,N$, $t = 1,...,T$ and $m = 1,..,M$. T, N, M refer, respectively, to the time series dimension, the number of cross sectional regions and the number of regression variables, respectively. Pedroni (1999) develops asymptotic and finite sample properties of testing statistics to examine the null hypothesis of non-cointegration in the panel. The tests allow for heterogeneity among individual members of the panel. Four tests statistics are based on the within dimension and three on the between-dimension (see Pedroni, 1999, for more details).

3. Empirical estimation and analysis of the main results

We test for panel unit root using two different first generation tests and one-second generation test. Our empirical investigations will begin by testing stationary using these tests. Testing for stationarity is the first step of panel cointegration procedure. The proof of the same order of integration for all the variables allows testing for panel long-run relationship between the variables integrated in the same order and the long-run relationship can be estimated by FMOLS or DOLS without ambiguity (see Pedroni, 1999).

Panel unit root tests results are shown in Table 3 and Table 4, for both levels and first differences and with two specifications of the deterministic, namely with an intercept only and with an intercept and a linear trend. For the lag-length selection, the Akaike information criterion is used to specify the optimal lag.

From Table 3, which reports calculation results of panel unit root for the two first generation tests, we clearly see that the variables are not stationary. The non-stationary feature characterizes all variables for the two consumption blocks, with and without time trend specification of the model used to test unit root.

However, as shown by the large negative values of LLC (2002) and IPS (2003) statistics for variables in first difference, we conclude that all the variables become stationary after differentiation. Therefore, the variables in first difference are stationary or integrated of order zero, $I(0)$, which means their levels are integrated of order one, $I(1)$.

Table 4 reports the cross-sectionally augmented IPS (Im et al., 2003) panel unit root tests by Pesaran (2007). In the level case, the variables are non-stationary and they become after first difference, which confirm results of previous first generation tests. The variables are integrated for order 1.

Table 3. First generation panel unit roots tests

	Lower block				Upper block			
	LLC		IPS		LLC		IPS	
	Trend	No trend	Trend	No trend	Trend	No trend	Trend	No trend
Pr	-1.11*	-0.93	-2.31*	-3.19*	-5.57*	-1.61	-7.8*	-1.61
P	-4.93*	-0.19	-6.22*	-0.22	1.94	-0.49	4.17	-0.67
TM	-0.97	2.94*	-1.06	1.04	-	-	-	-
N	0.28	1.99	1.2	2.65	-0.48	1.04	-0.01	1.16
RL	-0.84	-1.18	-1.21	-1.91	-	-	-	-
ΔY	-17.4*	-22.7*	-31.6*	-22.8*	-10.16*	-11.2*	-20.6*	-12.34*
ΔP	-14.3*	-12.6*	-23.6*	-26.5*	-16.5*	-17.2*	-27.9*	-22.8*
ΔTM	-13.01*	-17.56*	-28.1*	-29.2*	-	-	-	-
ΔN	-11.16*	-13.6*	-14.3*	-14.9*	-18.03*	-21.7*	-24.1*	-23.7*
ΔRL	-18.57*	-18.1*	-31.2*	-30.56*	-	-	-	-

Note: * denotes the rejection of the null of panel unit root at the 5% level.

We have demonstrated that all the variables have the same integration order. Thus, we can turn to test for panel cointegration relationships between the quarterly share of subscribers in each block *(Pr)* as the dependent variables and its determinants (P, I, N, RL and TM). The seven tests proposed by Pedroni (1999) are implemented and reported in Table 5. The majority of these tests reject the null of no cointegration, which indicates the existence of long-run equilibrium between the variables that are $I(1)$. Indeed, in the long-run, the share or the proportion of households in each block is determined by water pricing structure, climatic variables and the evolution of the quarterly size of the network. We can turn to estimate the long-run relationship for the two consumption blocks. The results are reported in Table 6. These results include the individual and the group FMOLS estimates.

The panel long-run relationship, between the share of subscribers in each block and its determinants, was estimated by the FMOLS method, which has been recognized as the best method for estimating panel cointegration relationships. The results are generally statistically significant and in accordance with the theoretical requirements, as well as the social and economic intuitions.

The empirical estimation allows understanding the cross-correlations induced by the shifting of consumers from one water consumption block to another, an important factor for capturing the effect of price changes. Moreover, we are able to evaluate the success of the progressive water pricing policy as our results confirm that upper block consumers react more significantly than lower block consumers. Indeed, consumers in the lower range use a greater proportion of water consumption for the satisfaction of essential human uses (drinking, cooking and basic hygiene purposes). We are also able to perceive the sliding effect of people moving from one range of consumption to the next and the effect of new entrants to the network as a result of economic development.

Table 4. Pesaran (2007) panel unit root test

Variables	Level				First difference			
	Lower block		Upper block		Lower block		Upper block	
	Intercept	Intercept & trend	Intercept	Intercept & trend	Intercept	Intercept & trend	Intercept	Intercept & trend
$LnPr_{it}$	-0.94 (0.24)	1.09 (0.65)	-0.66 (0.47)	-0.86 (0.56)	-5.23 (0.00)	-7.81 (0.00)	-4.21 (0.00)	-5.76 (0.00)
LnN_{it}	-0.04 (0.39)	-0.86 (0.56)	-0.13 (0.15)	-0.76 (0.45)	-6.18 (0.00)	-8.71 (0.00)	-6.95 (0.00)	-4.32 (0.00)
LnP_{it}	-0.56 (0.67)	-0.71 (0.53)	-0.87 (0.34)	-0.96 (0.23)	-9.56 (0.00)	-8.07 (0.00)	-7.45 (0.00)	-6.78 (0.00)
$LnRI_{it}$	-0.67 (0.23)	-0.92 (0.63)	-0.47 (0.51)	1.34 (0.56)	-5.67 (0.00)	-4.96 (0.00)	-6.87 (0.00)	-5.98 (0.00)
$LnTM_{it}$	-0.98 (0.19)	1.23 (0.14)	-0.88 (0.35)	1.02 (0.54)	-7.25 (0.00)	-7.01 (0.00)	-7.96 (0.00)	-6.45 (0.00)

Note: *p*-values for the null hypothesis of non-stationarity are reported between parentheses. Individual lag lengths are based on Akaike Information Criteria (AIC).

The long-run effect of water pricing policy in the proportion of subscribers in the lower block is positive, which means that consumers who consume under the threshold of 40 m^3 are not discouraged by the tariff rate of the first two brackets. However, as an incentive tool, upper block water price incite households to reduce their consumption by moving to the lower block. The negative sign of the water pricing effect on proportion of subscribers means that in winter, there is a sliding effect of consumers from high consumption level block to the lower one

As consumers switch from a higher block to a lower one in the case of a price increase, since demand is expected to be relatively inelastic to price in the lower block, the long-run coefficient is even positive in this block due to the fast wage increases in Tunisia. Consequently, the nonlinear pricing policy ensures efficiency in water use, as it incites big consumers to preserve water. Moreover, it guarantees equity, as the new entrants and the lower block consumers, who are generally low income households, are not sensitive to it.

In Tunisia, the average yearly income is 4200 USD in 2005, which corresponds to 3000 euros. Income is, as predicted by economic theory, a real determinant of household behavior in water consumption. Table 6 shows that the long-run income effect estimate is around 0.5 in the upper block composed by higher income households. This kind of consumers is characterized by higher outdoor water consumption, because they enjoy in most cases a private garden and pool. Moreover, there is an heterogeneous income effect through the individual estimate.

The impact of income on the share of subscribers is significant only in the Center West and East. The positive impact of income shows that rich consumers are often an upper block's consumers and essentially low-income households compose the lower block. This is confirmed by the negative impact of income on the share of subscribers in the lower block. This kind of consumers becomes an upper block's consumers when income increases, which explains the negative relationship between the share of subscribers in this block and the income. Once again, in a developing country where average income is expected to rise over time, this result leads us to expect increases in water consumption due to increase in the share of subscribers in the upper block.

Table 5. Pedroni (1999) cointegration tests

	Lower block		Upper block	
	Trend	No trend	Trend	No trend
Panel-m	2.1	2.6	3.21	4.8*
Panel-q	-13.54*	-11.54*	-18.8*	-15.1*
Panel-pp	-14.2*	-11.4*	-18.52*	-16.1*
Panel-adf	1.6	0.95	-14.6*	-13.8*
Group-q	-17.2*	-17.5*	-19.5*	-20.3*
Group-pp	-17.8*	-18.3*	-20.4*	-20.5*
Group-adf	1.3	1.5	-16.8*	-16.7*

Note: * denotes the rejection of the null of panel unit root at the 1% level.

The share of subscribers in the upper block increases during the dry seasons and decreases during the wet ones, as we see from the long-run positive and negative effects of temperature and rainfall, respectively. The total number of subscribers to the network has a negative effect on the proportion of consumers in the upper block, and this is certainly due to the fact that the new entrants are generally low-income households with low water consumption. However, the share of subscribers in the lower block, in contrast to the upper one, increases during the wet and decreases during the dry seasons, as can be observed from the long-run negative and positive coefficients of temperature and rainfall, respectively.

Table 6. FMOLS estimation results

Consumption block	Lower block					Upper block				
	LnP	LnN	LnI	LnTM	LnRL	LnP	LnN	LnI	LnTM	LnRL
Center West	0.1*	0.03	-0.1	-0.17*	0.06*	-0.72*	-0.01	0.91	1.45**	-0.11*
	(1.52)	(-0.56)	(-2.1)	(-2.13)	(5.47)	(-2.22)	(-0.08)	(1.8)	(1.66)	(2.05)
Center East	0.01	0.24*	0.1	0.10	0.03*	-0.4*	-0.56*	0.9	0.79**	-0.02
	(0.10)	(-4.63)	(1.07)	(1.07)	(2.95)	(-3.88)	(-4.16)	(4.3)	(1.81)	(-0.4)
North East	0.12*	0.40*	0.5	-0.52*	0.03*	-0.03	-0.63*	0.2	0.26	-0.01
	(2.61)	(-10.3)	(4.5)	(4.53)	(3.96)	(-0.15)	(-3.78)	(0.33)	(0.33)	(-0.7)
North West	0.1	0.17*	0.3	-0.22*	0.01	0.30	-1.28*	0.6	0.64	-0.02
	(1.35)	(-3.7)	(2.3)	(2.95)	(0.65)	(0.85)	(-2.5)	(1.04)	(1.04)	(-0.47)
South	0.21*	0.27*	0.01	0.11	0.01*	-0.70*	0.33*	0.3	0.37	-0.04*
	(3.1)	(-4.7)	(0.06)	(0.16)	(2.26)	(-3.71)	(2.28)	(0.18)	(0.18)	(-1.64)
Great Tunis	0.12	0.42*	0.5	-0.66*	-0.03*	-0.30*	-0.55*	0.2	0.21	-0.01
	(1.2)	(-9.1)	(4.2)	(3.24)	(-3.33)	(-2.62)	(-6.32)	(0.8)	(0.87)	(-0.1)
Panel (without trend)	0.1*	0.24*	0.2*	-0.32*	0.02*	-0.31*	-0.39*	0.5	0.57*	-0.01
	(2.6)	(-13.9)	(4.1)	(5.4)	(4.88)	(-5.02)	(-5.66)	(3.3)	(3.34)	(0.09)
Panel (with trend)	0.53*	0.35*	0.16	-0.2**	0.01*	-0.56*	-0.42*	0.31	0.35*	-0.06*
	(3.14)	(-14.63)	(3.4)	(2.55)	(2.01)	(-7.23)	(-6.12)	(2.6)	(2.03)	(2.43)

Note: *and ** indicates statistical significance at the 5% and 10%, respectively. t-statistics in parenthesis. The variables are in natural logarithms.

Conclusion and policy implications

In this paper, we analyzed data from Tunisia to assess the effect of two-part tariffs on consumption. Our results, using a panel dataset describing six heterogeneous Tunisian regions over a long period, imply that increasing block tariffs are effective to struggle against water scarcity. This multi-regional approach has several implications. First of all, as noticed by Olmstead and Stavins (2009), price elasticity varies geographically and over time, because higher prices result in higher price-elasticities all else equal. Indeed, communities that regularly experience arid conditions and in which water shortage is a relatively more frequent occurrence tend to have higher water prices, on average, than communities in which water is plentiful. The use of regional intra- and inter-regional estimations helps going beyond that point.

Second, we test the conventional wisdom that suggests that increasing block tariffs are "equitable" pricing structures to preserve the resource, since households with low water consumption pay a smaller marginal price than households with high water consumption. However, the distributional impacts of increasing block pricing, can cause a greater consumption reduction for low-income than for high-income households. Mansur and Olmstead (2012) examined the distributional impacts of various demand management policies in a study of 11 North American cities. Using residential water consumption data, they are able to generate demand curve estimates for residential water and to simulate the effects of a price increase that would result in the same aggregate water consumption reduction as the two-day per week outdoor watering restriction. Drought pricing, relative to the prescriptive approach, rises the consumption share of households above both the sample median income and lot size would rise from 35 to 48 percent; the consumption share for households below both median income and median lot size would fall from 23 to 16 percent. A progressive price-based approach to water demand management should, then, be designed in order to re-allocate potential extra-profits based upon income to low-income consumers. Such transfers could, however, weaken the resource preservation goal, as low-income consumers would consume more water.

Studying the impact of rationing policies in Tunisia is well above the scope of this article. However, whether rationing policies could replace price increases remains an open question. The abovementioned study by Mansur and Olmstead (2012) computes the gains of a market-clearing price of an aggregate demand reduction equivalent to a two-day-per-week outdoor watering restriction. They find that they make up $92 per household per summer, i.e., 30 percent of what the average household in the study sample spent each year on water. Although nonprice conservation programs can reduce water consumption, both economic theory and the emerging empirical estimates suggest price increases are most cost-effective when one considers the potential impact on the overall welfare. Given the prevalence of current water scarcity and access problems in the world, even in industrialized countries such as Australia or the United States, further economic research on water market and pricing structures, potential externalities, and distributional impacts should be developed.

References

1. Arbués, F., Garcia Valinas, M. A., and Martinez-Espineira, D. R. (2003). Estimation of residential water demand: a state of the art review. *Journal of Socio-Economics, 32*(1), 81-102.
2. Ayadi, M., Krishnakumar, J., and Matoussi, M. S. (2002). A Panel data analysis of residential water demand in presence of nonlinear progressive tariffs, cahiers du départementd'économétrie, University of Genève, 2002.06.
3. Dalhuisen, J. M., Florax, M. R. de, Groot, H. L. F., and Nijkamp, P. (2003). Price and income elasticities of residential water demand: why empirical estimates differ? *Land Economics, 79,* 292-308.
4. Im, K. S., Pesaran, H., and Shin, Y. (2003). Testing for unit roots in heterogenous panels. *Journal of Econometrics, 115,* 53-74.
5. Levin, A., Lin, C. F., et Chu, J. (2002). Unit root tests in panel data: Asymptotic and finite sample properties. *Journal of Econometrics, 108,* 1-24.
6. Martinez-Espineira, R. (2007). An estimation of residential water demand using cointegration and error correction techniques. *Journal of Applied Economics, 1,* 161-184.
7. Mansur, E. T., and Olmstead, S. (2012). The Value of Scarce Water: Measuring the Inefficiency of Municipal Regulations. *Journal of Urban Economics, 71*(3), 332-346.
8. Nauges, C., and Whittington, D. (2010). Estimation of Water Demand in Developing Countries: An Overview. *World Bank Research Observer, 25*(2), 263-294.
9. Nauges, C., and Thomas, A. (2003). Long-run study of residential water consumption. *Environmental and resource Economics, 26,* 25-43.
10. Olmstead, S. (2010). The Economics of Managing Scarce Water Resource. *Review ofenvironmental economics and policy, 4*(2), 179-198.
11. Olmstead, S., and Stavins, R. N. (2009). Comparing Price and Non-price Approaches to Water Conservation. *Water Resources Research, 45*(4), 1-10.
12. Pedroni, P. (1999). Critical values for cointegration tests in heterogenous panels with multiple regressors. *Oxford Bulletin of Economics and Statistics, 61,* 653-670.
13. Pesaran, H. (2007). A simple panel unit root test in the presence of cross-section dependence. *Journal of Applied Econometrics, 22,* 265-312.
14. Porcher, S. (2014). Efficiency and equity in two-part tariffs: the case of residential water rate. *Applied Economics, 46,* 539-555.
15. Worthington, A. C., and Hoffman, M. (2008). An empirical survey of residential water demand modeling. *Journal of Economic Survey, 22,* 842-871.

Appendix

Following Ayadi et al. (2002), we break the five consumption blocks into two distinctive parts. First, the lower block puts together the consumers of the first two blocks (0-40 m^3), thus, covering basic needs and, therefore, obviously characterized by a steady aggregate consumption level as shown in Figure 1A:

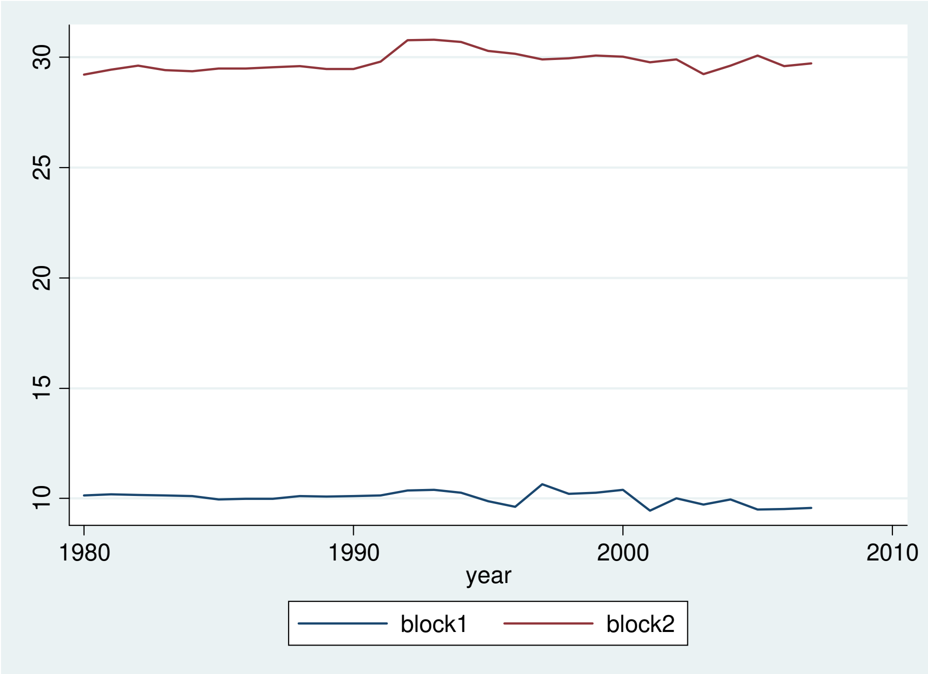

Fig. 1A. Lower consumption block, yearly average values

More precisely, the threshold of 40 m^3 per quarter corresponds to an average daily consumption equals to 88 litres per day and per capita (considering a five persons family). It is, therefore, very much lower than average residential consumption levels in the UE-15 countries, where average daily per capita water use range from 115 litres in Belgium to 265 litres in Spain, EWA (2002).

Next, the upper block gather the three latest blocks (more than 41 m^3 per quarter), which exhibits a decreasing trend, probably explained by the tariff progressivity, as shown in the following two figures:

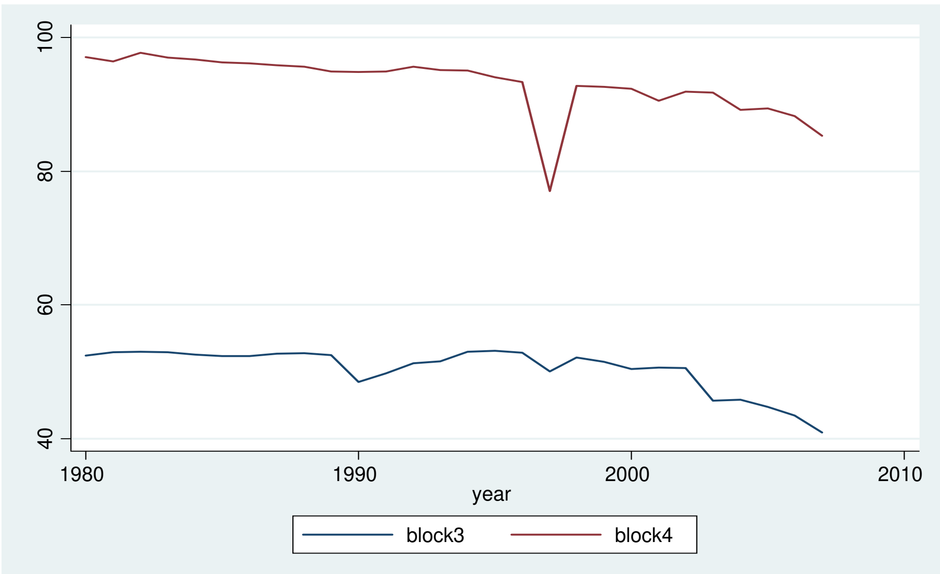

Fig. 2A. Upper consumption block (block 3 and 4), yearly average values

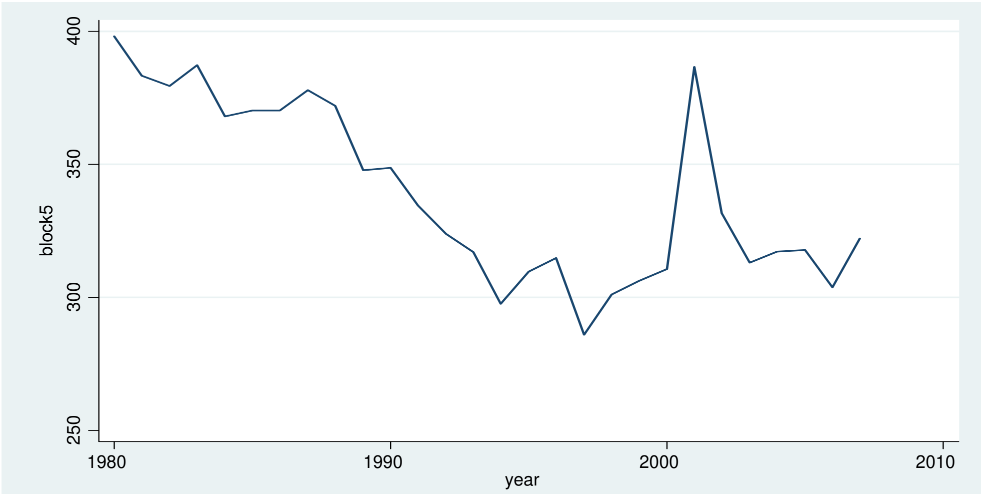

Fig. 3A. Upper consumption block(block 5), yearly average values

We observe an important decrease in residential water consumption in 1997, which was a particularly rainy year. Conversely, the peak observed for block 5 in 2001 can be explained by a period of severe drought, showing the role played by seasonal climate fluctuations to explain residential water consumption. Compared to the initial five blocks scheme, the lower versus upper blocks decomposition may improve the quality of our estimation results[2] as it increases price and consumption variability (compared to a distinct analysis for each block) without distorting the economic analysis, as the increasing block rates feature is preserved in each block so constituted.

[2]Using the same two blocks decomposition, Ayadi et al. (2002) observe that the quality of their results is improved.