Modelling seasonality in residential water demand: the case of Tunisia

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This article proposes to model seasonal patterns of residential water demand using the techniques of seasonal integration and cointegration. The methodology is applied to aggregated quarterly time series data for Tunisia (1980-2007), applying the same increasing, multi-step pricing scheme in the whole country. First, a seasonal cointegration analysis demonstrates the relevance of a pricing policy that increases the size of the lower consumption block in summer. Second, the non seasonal cointegration analysis reveals a relatively high price elasticity for the highest consumption block. Therefore, we also propose to increase the tariff progressivity to promote water saving. This modified pricing scheme will help to achieve goals of environmental protection and social equity.

JEL classification: C22 ; Q25.

I. Introduction

The main purpose of this article is to adequately model seasonal effects in residential water demand functions using tools available in the time series field. We thus use quarterly data from 1980 to 2007 in Tunisia. Indeed, Tunisia is a suitable case study for data aggregated at the national level as the same water tariff scheme is applied in the whole country. Therefore, useful recommendations in terms of pricing policy

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must deal with the country level, without taking into account regional disparities.

In the literature, the role played by seasonality in residential water demand is an important issue that has often been neglected (Martinez-Espineira, 2003). Indeed, the effect of climate fluctuations and modifications in habits imply that water consumption probably follows seasonal fluctuations. Sensitivity to a given change in price may be different across seasons. In the literature, some studies (Martinez-Espineira, 2002, for example) estimate a water demand model for each season while others assume the existence of a purely deterministic seasonal process generated by seasonal dummy variables (Barry *et al*., 2012, for example).

Our study takes an innovative approach by implementing rigorous seasonal unit roots to fully describe the seasonal patterns of the times series included in the water demand function. Martinez-Espineira (2007) was the first to investigate unit root tests using monthly time series describing residential water consumption in Spain. Unfortunately, he failed to detect seasonal unit roots. In our study, empirical results reveal the presence of common stochastic seasonal components, enabling us to develop a cointegration analysis. We then use the selection test developed by Franses (1993) to choose between periodic *versus* seasonal cointegration.

Cointegration and error correction model techniques allow short and long-run price elasticities of residential water demand to be calculated and compared. A short-run analysis of water demand aims to quantify the impact of the factors influencing the duration and the frequency of equipment use. A long-run analysis of water demand emphasizes the role of the determinants responsible for changes in the size and water-efficiency of the stock of appliances. For example, using time series observations from Sevilla (Spain), Martinez-Espineira (2007) derived 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. As the short-run price elasticities were smaller than their long-run counterparts, this and other authors conclude that tariff policies are more efficient in the long-run and recommend subsidies for water efficient appliances. Distinguishing short from long-run analyses is relevant in Tunisia, which is a developing country where the average standard of living is on the rise.

To the best of our knowledge, we are the first to use seasonal cointegration to analyse residential water consumption. We use a rich quarterly data set from the first quarter of 1980 to the fourth quarter of 2007 in Tunisia. The data, obtained from SONEDE the national water distribution company, include aggregated time series for residential water consumption, average price, rainfall, temperature, the number of residential consumers in each block and yearly income values.

In our methodology, the first step is to conduct the Hylleberg *et al.* (HEGY, 1990) seasonal unit roots tests at zero, annual and biannual frequencies. We then study seasonal cointegration using the Engle *et al.* (EGHL, 1993) methodology. Cointegration in the long-run frequency (or at the zero frequency) can be interpreted as indication of a parallel long-run movement in the nonstationary series. Cointegration in the seasonal frequency can be interpreted as evidence of a parallel movement in the seasonal component of the series which exhibit a stochastic seasonal pattern. Finally, we investigate seasonal error correction models.

Applied to the Tunisian time series data, we observe cointegration at both biannual and zero frequencies in the lower consumption block, but only at the zero frequency for users in the upper consumption block. Our results confirm that, in the long-run, users in the upper block are more likely to react sharply to price variations than those in the lower, where a greater proportion of water is devoted to essential uses. We propose that tariffs should be made more progressive to discourage high consumption levels. The results that we obtained from introducing seasonality furthermore suggest that Tunisian authorities should increase the size of the lower block to ensure the satisfaction of households' essential needs in all seasons.

The paper is organized as follows. Section II presents a brief overview of the introduction of seasonality in residential water demand modelling. Section III describes the quarterly data used in our empirical analyses. Section IV develops the methodology we use to estimate short and long-run elasticities of demand including seasonal effects. Section V presents and discusses the results of our empirical analyses. It outlines recommendations for innovative pricing policies to induce household water saving behaviour in higher blocks without affecting the well being of relatively poor people by keeping lower block prices constant.

II. Seasonality in residential water demand modelling

Seasonal fluctuations could be an important source of variation in residential water consumption, especially for outdoor uses. If this is the case, adequate water management policies are required. Seasonal pricing could thus be an effective water conservation policy. It is used already in the USA, notably in Phoenix, Arizona (Yoo *et al.*, 2014, for example). Furthermore, in response to a severe drought, water restrictions on outdoor water use, commonly used in Australia for example, can be imposed.

But such policies cannot be analyzed without improved knowledge of the role played by seasonality in residential water demand models. However, seasonal patterns of demand tend to be under represented in the literature. Since the first study using bi-annual data was conducted (Hanke and De Maré, 1982), few other studies have included seasonal effects in residential demand models.

Griffin and Chang (1991) examine the influence of seasonality by estimating a demand model for each month. Their main results indicate that summer price elasticities exceed winter price elasticities by 30%. Dandy *et al.* (1997) estimate and compare three linear water demand equations (one full year and two seasonal ones, differentiating summer from winter). They obtain seasonal elasticities in the range of -0.29 to -0.45 for winter and -0.69 to -0.86 for summer, and income elasticities in the range of 0.32 to 0.38 for annual consumption, 0.28 to 0.33 for winter, and 0.41 to 0.49 for summer. This agrees with the theory that outdoor uses are more responsive to income levels. Martinez-Espineira (2002) develops a similar methodology. He compares two estimation results from a panel of monthly aggregate data from Spain, using both the whole sample and a sub-sample including summer observations only. Results also show that summer price elasticities are slightly higher than those of other seasons.

A few other studies use a simple modelling of seasonality by adding seasonal dummy variables in their residential water modelling. They implicitly assume the existence of a purely deterministic seasonal process. Ayadi *et al.* (2002) introduce three dummy variables for quarters, excluding the one for the dry season. They find that the coefficient estimates are negative, implying an expected lower consumption in these seasons, but for low consumption levels only. Similarly, using monthly data from June 2004 to June 2009, Barry *et al.* (2012) show that moving from summer to winter results in an estimated reduction of approximately 17% in average daily water use. However, as noted by Franses and McAlleer (1998), the constant seasonality model generally can be dismissed in economic data and there is empirical evidence in favour of seasonal unit roots, i.e. a seasonal variation that is not constant over time.

Two more complex models of seasonality have been proposed. First, Renwick and Green (2000) develop a model including two price equations (marginal price and difference variable equations inspired by Nordin (1976)) and a climate equation. In the latter equation, seasonality is captured through a harmonic model, consisting of a Fourier series of sine and cosines terms of various harmonic frequencies. Using monthly time series data for 8 water agencies in California covering 1989 to 1996, these authors show that demand is 25% more price responsive in the summer months, reflecting the more discretionary nature of outdoor water use. They conclude that price policy will reduce demand more during summer months or in communities with larger irrigated landscape areas.

In the second and more recently developed model, Bell and Griffin (2011) use the technique of periodic cointegration, developed by Boswijk and Franses (1995), to analyse seasonal patterns. Using monthly data from a ten-year panel of 167 US cities, Bell and Griffin (2011) do not reject the hypothesis that all quantity data are periodically integrated. They first implement log-log regressions by month before estimating an error correction model. Their results notably indicate the unlikelihood of a constant demand relationship and support the probability of 12 phase (monthly) seasonality, demonstrating the importance of slow adjustments.

To sum up, these studies indicate that seasonality does influence residential water consumption behaviour. Models that do not include seasonal patterns are therefore misspecified. Yet with the exception of Bell and Griffin (2011), the above mentioned studies do not provide a satisfactory and complete modelling of seasonality. In this article, we aim to fill this gap.

III. Context and data set description

Water resources in Tunisia are characterized by their scarcity, low quality, poor distribution and seasonal fluctuations. Tunisia has a semi-arid climate with alternating dry and rainy seasons. With regard to scarcity, figures published by the World Bank show that available water resources in Tunisia, calculated to be 420 m^3 per quarter per household in 2005 (respectively 500 m^3 in Algeria but 1600 m^3 in France), will fall to 300 m^3 in 2030. Tunisia therefore faces a real water supply crisis that will be accentuated over the next two decades.

Agriculture accounts for the largest part of water consumption, i.e. 80% of total resources (residential consumption accounts for 13% and the tourism sector for the remaining 7%). But residential water consumption must be carefully managed for at least two reasons. First, it concerns the satisfaction of basic needs. Second, residential water demand is increasing due to rapid urbanization. Therefore, as Tunisia cannot resort to unconventional and costly alternatives (such as desalination), water pricing must be considered to ensure the long-term sustainability of water use.

To address these issues, the objective of this article is to investigate the role of seasonality in residential water demand behaviour in Tunisia to know if a seasonal pricing could thus be an effective policy. We thus use a rich and original database covering the period from the first quarter of 1980 to the fourth quarter of 2007, i.e. 28 years. These data are thus particularly well suited to an analysis of the determinants of residential water consumption over the long-term that includes seasonal effects. The data, which were obtained from SONEDE, the national water distribution company, include aggregated quarterly time series for Tunisia on average residential water consumption, average price, the share of customers in each block, rainfall, and temperature. Annual income data derived from budget surveys compiled by the National Statistical Institute also were collected. Price and income are expressed in constant Dinars. As the same water tariff scheme is applied in the whole country, the use of aggregate data at the national level is justified.

Given our aggregated data, we retain a conventional average price specification defined in Table 2. We do not use the theoretically correct price specification (weighted marginal price and Nordin's variable calculated using the proportions of users per block), as the two price specifications give close results (Martinez-Espineira, 2003). See also Gaudin (2006) and Binet *et al.* (2013) for further discussion about water price perception.

SONEDE has applied a five-block pricing structure since the 1970s. The frequency of billing is quarterly. The fixed charges are around 3.3 Dinars (i.e. 0.08% of average income) in the lower block, but 35 Dinars (i.e. 8.9% of average income) in the upper block. The average share of the water billing in households' annual income is estimated to be 11% for a quarterly consumption level equals to 80 m^3.

As this pricing structure has now been effective for forty years, we presume that consumers have had sufficient time to react and adapt their consumption behaviour accordingly. In the following Table 1, we present the size and 2007 tariff rates of the five residential consumption blocks.

Table 1. Residential water blocks and tariffs in Tunisia in 2007

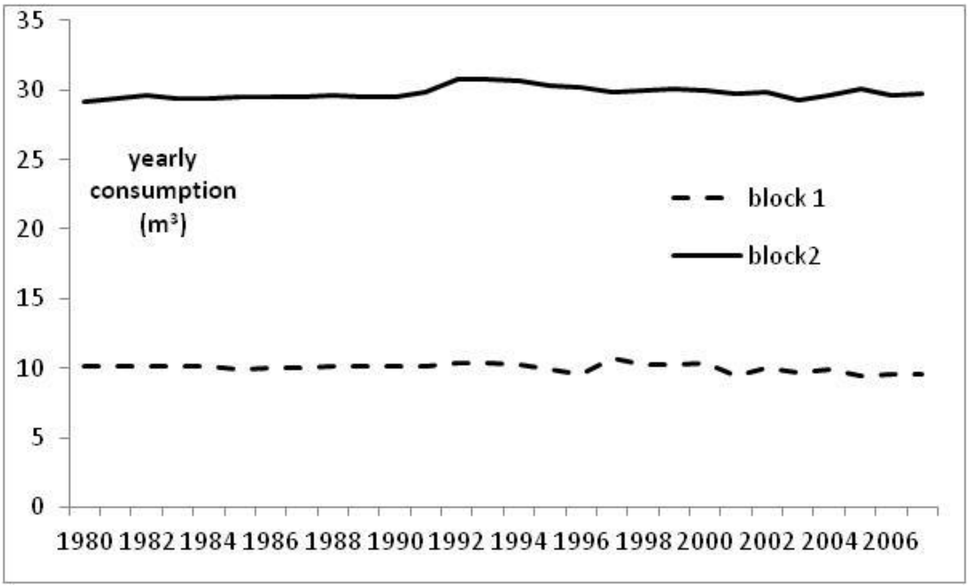
Block size	0-20 m^3	21-40 m^3	41-70 m^3	71-150 m^3	> 150 m^3
Unit price (Dinars)	0.14	0.24	0.3	0.545	0.84

Source: SONEDE

Such an increasing block rate pricing scheme may discourage high consumption levels through price increases in higher blocks without affecting the well-being of low-level consumption customers by applying a social rate for the lowest consumption blocks. Indeed, this tariff scheme is relatively highly progressive as the ratio of the last rate to the first rate is six. In Tunisia, a super-increasing block rate pricing scheme has been used since 2005 to better manage the scarcity of water resources. Under this pricing scheme, a household pays for its entire water consumption at the rate of the last block reached. But the long-run consequences of this pricing scheme cannot be analyzed for the period under study (1980-2007).

Following Ayadi *et al.* (2002), we broke the five consumption blocks into two distinct parts. The lower block groups together the consumers of the first two blocks (0-40 m^3), thus covering basic needs. As shown in Fig. 1, the aggregate consumption levels for this group is steady.

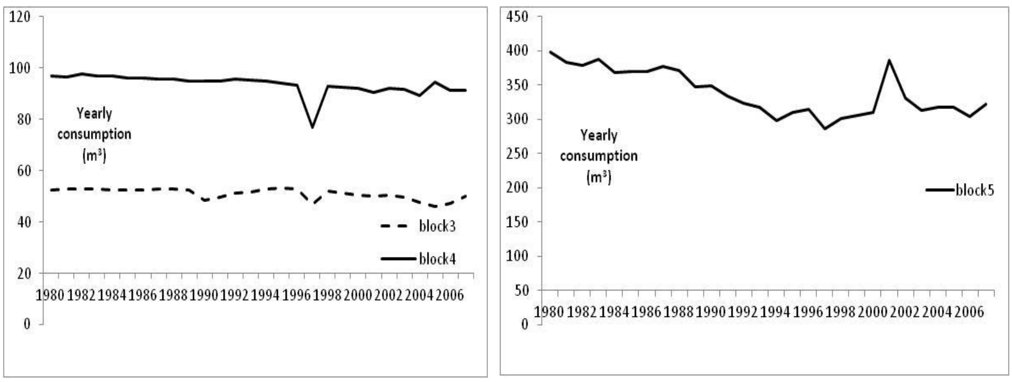
Fig 1. Lower consumption block, yearly average values



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

The upper block groups the three higher blocks (more than 41 m^3 per quarter). It exhibits a decreasing trend, probably due to the tariff progressivity, as shown in the following two figures.

Fig 2. Upper consumption block, yearly average values



One can observe an important dip in residential water consumption in 1997, which was a particularly rainy year. Conversely, the peak observed for block five in 2001 can be explained by a period of severe drought, showing the role played by climate fluctuations in residential water consumption, even using aggregate data.

Compared to the initial five-block scheme, this decomposition into a lower and upper block may improve the quality of our estimation results[1] as it increases price and consumption variability (compared to a distinct cointegration analysis for each block). It also permits to give recommendations to achieve goals of environmental protection and social equity.

A description of the variables in the two consumption blocks and basic descriptive statistics are presented in Table 2.

Table 2. Description of the variables and basic descriptive statistics, 1980.1 to 2007.4

Variable	*Description*	*Mean*	*Max*	*Min*
Average residential water consumption (lower block, m^3)	Sum of consumption in the two first blocks divided by the corresponding number of users.	19.86	40	9.04
Average price (lower block, Dinars)	Total water bill divided by the total volume of water consumed in the two first blocks.	0.39	0.85	0.20

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

Average residential water consumption (upper block, m^3)	Sum of consumption in the three upper blocks divided by the corresponding number of users.	150.61	341.50	54.11
Average price (upper block, Dinars)	Total water bill divided by the total volume of water consumed in the three upper blocks.	0.75	1.33	0.23
Yearly income (Dinars)	Drawn from the expenditure surveys by the National Statistics Institute.	1570	2549.50	1218
Temperature (degrees)	Average quarterly temperature	19.37	33	9
Rainfall (millimetres per quarter)	Average quarterly level of precipitation	172	601	11
Percentage of consumers (lower block)	Share of users	73	85	55
Percentage of consumers (upper block)	Share of users	9	22	4

Source: SONEDE.

On average, the lower block represents 73% of subscribers and 53% of total domestic consumption, whereas the upper block accounts for 47% of total residential consumption. Average consumption in the upper block is eight times higher than the average in the lower block. In Tunisia, the average yearly income is 1570 Dinars, which corresponds to 785 Euros. Average rainfall, which varies from 11 to 601 millimetres per quarter, is expected to have a negative influence on residential water consumption. Conversely, water consumption is expected to increase when temperatures increase.

The proportion of subscribers may be used as a proxy for network expansion, particularly in the lower block. It is thus a variable able to take into account the specific characteristics of a developing country where the distribution network is expanding quickly. This variable also measures the effect of new entrants in each block as a result of seasonal variations in consumption levels. If the average consumption of new entrants in one block is lower than that of existing consumers, a negative coefficient for the percentage of subscribers may be expected.

Table 3 presents some aggregate statistics about quarterly fluctuations of the variables, and thus gives a preliminary description of seasonal patterns over the long-run.

Table 3. Average quarterly values

Variables	*Winter*	*Spring*	*Summer*	*Autumn*
Lower block consumption (m^3)	15	16	24	17
Upper block consumption (m^3)	128	116	182	116
Lower block average price (Dinars)	0.43	0.41	0.34	0.41
Upper block average price (Dinars)	0.72	0.68	0.81	0.68
Percentage of consumers, lower block (%)	76	69	67	69

Percentage of consumers, upper block (%)	7	11	12	11
Rainfall (millimetres per quarter)	187	154	150	154
Temperature (degrees)	12	16	29	12

These figures suggest that most of these variables are seasonal in nature. Indeed, residential consumption and average price in the upper block seem to exhibit two peaks a year (in summer and, in a lower proportion, in winter). Conversely, one can observe one clear peak (in summer) in the temperature distribution. The seasonal component of other variables is not easy to define. To move beyond these preliminary observations, we must analyse the seasonal properties of the quarterly time series.

IV. Empirical methodology

Hylleberg *et al.* (HEGY, 1990) developed the first test procedure to deal with seasonal frequency in quarterly time series (first subsection). In the second subsection, we discuss the two corresponding cointegration techniques available in the literature (i.e. seasonal versus periodic cointegration) and present the test selection between the two methodologies. Finally, we conclude by showing the error correction model.

Tests for seasonal integration

As many economic time series exhibit evolving seasonality, HEGY (1990) developed tests for seasonal unit roots at different frequencies using quarterly data. The properties of seasonally integrated series are quite similar to those of ordinary integrated processes as shocks may permanently change their seasonal patterns.

More precisely, a quarterly time series has seasonal unit roots when the $\Delta_4 = (1 - B^4)$ filter is appropriate, where B is the lag operator. As the polynomial $(1 - B^4)$ can be expressed as $(1-B)(1+B)(1-iB)(1+iB)$, the unit roots are *1*, *-1*, *+i* and *-i*. The first corresponds to zero frequency. The frequency is $\frac{1}{2}$ when the series exhibits one half cycle per quarter or two cycles a year (unit roots at the biannual frequency). The frequency is $\frac{1}{4}$ when one quarter cycle per quarter or one cycle a year (annual frequency) is observed.

For a quarterly time series *y*, the tests are based on the OLS estimation of the following auxiliary regression:

$$(1)\quad \Delta_4 y_t = (1-B^4)x_t = \alpha_1 y_{1,t-1} + \alpha_2 y_{2,t-1} + \alpha_3 y_{3,t-2} + \alpha_4 y_{3,t-1} + \sum_{i=1}^{p} \Delta_4 y_{t-i} + \mu_t + \varepsilon_t$$

Where $y_{it} = \theta_i(B) y_t$ for i=1… 3 with:

$$\begin{cases} \theta_1(B) = (1 + B + B^2 + B^3) \\ \theta_2(B) = -(1 - B + B^2 - B^3) \\ \theta_3(B) = -(1 - B^2) \end{cases}$$

Note that the deterministic component μ_t is added to the regression to include seasonal dummies (*SD*), a linear time trend (*Td*) and a constant term (*I*). The term ε_t is a normally and independently distributed error term. The regression (1) is augmented by additional significant lagged values of the dependent variable to whiten the residuals. The lag length *p* is based on the selection of the latest significant lag.

The regression (1) is estimated by OLS. The test for the unit root at the zero frequency is simply a significance test for $\alpha_1 = 0$, and for biannual frequency $\alpha_2 = 0$. For the complex root, HEGY (1990) suggests the following joint test: $\alpha_3 = \alpha_4 = 0$ or in case of $\alpha_4 = 0$ on the *t* values of α_3 respectively. To determine whether a series has no unit root and is therefore stationary, one must establish that each of the $\alpha_i's$ is different from 0. Using Monte Carlo simulations, HEGY (1990) provides critical values for the different significance tests presented above.

The presence of a common seasonal unit root in all of the variables included in a residential demand function allows the corresponding cointegration model to be analyzed.

Seasonal versus periodic cointegration

As previously mentioned, two earlier studies used cointegration analysis to model residential water demand using monthly data. Bell and Griffin (2011) used periodic cointegration whereas Martinez-Espineira (2007) tried to perform seasonal integration and cointegration analyses. A selection test can be implemented to choose between the two procedures. These aspects are developed in the following two subsections.

Seasonal cointegration. The seasonal cointegration model focuses on two or more series which have common non-stationary components. Two different estimation methods can be used. Lee (1992) developed a Johansen type (i.e. a VAR) approach. However, this approach has some drawbacks (Frances and McAlleer, 1998). If only a few variables are to be analyzed, as it is the case in this study, one can use a bivariate cointegration approach inspired by Engle and Granger (1987).

Engle *et al.* (1993) provide such a two-step procedure to test for cointegration at zero, biannual and annual frequencies. When a system of two variables consisting of y_t and x_t is considered, the following regressions are performed to collect the residual terms u_t , v_t, w_t:

(2) $y_{1t} = \beta_0 + \beta_1\theta_1(B)x_t + u_t$

(3) $y_{2t} = \gamma_0 + \gamma_1\theta_2(B)x_t + v_t$

(4) $y_{3t} = \delta_0 + \delta_1\theta_3(B)x_t + \delta_2\theta_3(B)x_{t-1} + w_t$

Where y_{it} and $\theta_i(B)$ are the same as described in the previous section.

First, the test for no-cointegration at the zero frequency can be performed using an auxiliary regression of Δu_t on u_{t-1} with or without deterministic components and augmented by the lagged values of Δu_t. Second, the test for no-cointegration at the biannual frequency is performed by the auxiliary regression of $(v_t + v_{t-1})$ on $-v_{t-1}$. The null hypothesis of no-cointegration at zero or biannual frequencies is not rejected if the t ratios for the coefficients to u_{t-1} and to v_{t-1} are smaller in absolute value than their critical values tabulated by Engle and Yoo (1987). Third, the null hypothesis of no-cointegration at frequency ¼ implies that both π_3 *and* π_4 are zero in the auxiliary regression of the form, augmented by lagged dependent variables:

$$(5)\ (w_t + w_{t-2}) = \pi_3(-w_{t-2}) + \pi_4(-w_{t-1}) + \varepsilon$$

The t ratios for π_3 and π_4 are used with a joint F test, and the corresponding critical values are tabulated in Engle *et al.* (1993) using Monte Carlo simulations.

Periodic cointegration and selection test. The periodic cointegration model developed by Boswijk and Franses (1995) considers an error-correction model in which the parameters vary across the seasons. More precisely, the periodic cointegration model assumes that the parameters in the cointegration vectors, as well as the adjustment parameters, can vary seasonally.

Franses (1993) proposes a simple selection method between seasonal and periodic cointegration by considering the usual cointegration tests for the annual series containing the observations per quarter. Indeed, the periodic cointegration model assumes that the relations between variables observed in the same season are cointegrated. Seasonal cointegration, however, does not assume such relations. The cointegrating Durbin-Watson and Dickey-Fuller tests (CRDW and DF) are therefore implemented on the residuals from each periodic cointegration relation. Critical values for these tests for n=100 are displayed by Franses and Kloek (1991). When there is non-stationarity for each period, the periodic model is rejected while the seasonal cointegration model may be adequate.

Error Correction Model. Here, we specify and estimate short-run effects through the following seasonal error correction model:

$$(6)\ \Delta_4 y_{1t} = \sum_{i=1}^{q} \alpha_i \Delta_4 y_{1,t-i} + \sum_{i=1}^{q} \beta_i \Delta_4 y_{2,t-i} + \delta_1 u_{t-1} + \delta_2 v_{t-1} + \delta_3 w_{t-2} + \delta_4 w_{t-3} + \varepsilon_t$$

y_{1t} and y_{2t} represent residential water consumption and its determinants, and lagged error terms *u*, *v* and *w* were previously defined. This equation can be estimated by OLS if all the terms are stationary. If the cointegration residuals at different frequencies *u*, *v* or *w* are statistically significant and negative, the size of the related coefficients measures the speed at which the variables adjust to restore the equilibrium in the relation describing residential water consumption. Furthermore, the other coefficient estimates give the short-run elasticities of demand.

V. Empirical results

In this section, after analyzing the unit root properties of all the variables, we perform the selection test to choose between seasonal and periodic cointegration. Given the results obtained, we implement the seasonal cointegration tests. Finally, a seasonal error correction model is estimated in the last subsection.

Tests for seasonal integration

Logged data are used throughout the following analyses. Unit root tests at seasonal frequencies are not implemented for income as the variable is only available on a yearly basis. The Breusch-Godfrey LM test is used to test for residuals autocorrelation. The outcomes of the HEGY tests[2] are shown in Table 4 in the appendix.

First, the results indicate that all the series, with the exception of income and average price in the upper block, are integrated of order one at the frequency zero whatever the deterministic component μ_t considered. The same result is obtained for income and average price in the upper block if we add a trend (*Td*) and a constant term (*I*).

We also find seasonal integration at the biannual frequency of water consumption, prices and the percentage of customers, in both the lower and upper blocks. However, the presence of unit roots at the biannual frequency for the quarterly rainfall and temperature variables is sensitive to inclusion of the deterministic components. Unit roots at the biannual frequency are present in both climate variables when an intercept, a trend and seasonal dummies are included. However, unit roots are only found at the biannual frequency for temperature at the 1% level. This implies that the test has some difficulty in separating a seasonal unit root at the

[2] The procedure hegy.src developed by T. Doan for RATS was used.

frequency $\frac{1}{2}$ from deterministic tendencies.

Finally, seasonal integration at the annual frequency is clearly proved only on the percentage of customers in the two consumption blocks. Conversely, the null hypothesis under the joint test is always rejected if we consider consumption blocks, average price in the upper block, and rainfall, implying that there are no systematic unit roots with an annual frequency.

Selection test between periodic and seasonal cointegration

The CRDW and ADF statistics to test the presence of one unit root in the residuals obtained from the four periodic estimations are shown in Table 5.

Table 5. Selection between periodic and seasonal cointegration

Quarter	*1*		*2*		*3*		*4*	
Block	Lower	Upper	Lower	Upper	Lower	Upper	Lower	Upper
CRDW	1.73*	1.01	0.60	1.16	0.60	1.16	2.08*	2.05*
ADF	-1.97	-2.33	-1.84	-1.30	-1.84	-1.30	-1.89	-2.89

Notes: Critical values for the ADF and CRDW tests are respectively -3.5 and 1.26 at the 5%.   ** if we reject the null hypothesis of no periodic cointegration.*

The ADF tests always indicate the presence of unit roots, implying no periodic cointegration. Conversely, if we consider the CRDW statistics, the hypothesis of stationarity of the residuals is retained in a few cases only (for the lower block in season 1 and for both blocks in season 4). The periodic model consequently is not appropriate for our case study, while the seasonal cointegration model may be adequate. The corresponding results are presented in the next subsection.

Seasonal cointegration tests

We use the procedure developed by Engle *et al.* (1993) to test for seasonal cointegration at the zero and biannual frequencies in the lower and the upper blocks separately. Results are reported in Table 6.

Table 6. Seasonal cointegration tests at the zero and biannual frequencies

Frequency	*Dependent variable*	*t ratio*	*Optimal lag*
	Lower block		
Zero	consumption	-3.74*	6
Biannual	consumption	-4.01*	4
	Upper block		
Zero	consumption	-7.64**	5
Biannual	consumption	-1.84	4

Notes: Critical values are respectively -4.02 and -3.71 at the 5% (**) and 10% (*) levels for N=4 and T=100 according to Engle and Yoo (1987). Optimal lag selection was made using the P-max approach as in HEGY (1990).

Our results show that the null hypothesis of the absence of seasonal cointegration is always rejected in the lower block for both frequencies. This implies long run equilibrium in residential water consumption and its determinants for this block (in levels and among their seasonal components). Indeed, residential consumption is particularly stable in the lower block, where consumers satisfy their essential needs. However, when we consider the upper block, the cointegration relation is statistically significant at the zero frequency only, and not at the biannual frequency (as, in absolute value, the t-ratio is lower than 3.71). In a developing country such as Tunisia, the consumption of high level consumers is therefore less stable than that of low level consumers. We thus observe seasonal instability in residential water consumption.

For further analysis, estimation results from the significant cointegrating equations are shown in Table 7.

Table 7. Cointegrating equations at zero and biannual frequencies

	Frequency	Price	Income	% customers	Rain	Temp.	DW	R^2
Lower	Zero	-0.07*	0.2***	-0.44*	-0.02**	0.6***	0.91	0.36
block		(-1.70)	(6.75)	(-1.76)	(-2.12)	(6,95)		
	Biannual	-1.95***	….	2.5***	-0.03	-0.004	3.84	0.78
		(-12.3)		(3.50)	(-0.57)	(-0.56)		

Upper block	Zero	-0.37*** (-6.06)	0.23*** (3.46)	0.17 (1.53)	0.01 (0.99)	1.77*** (9.96)	1.62	0.72

*Notes: All of the variables are in natural logarithm and DW denotes the Durbin Watson statistic. Figures in brackets are t-statistics of parameter estimates. Significance level for parameter estimates: *** for 1%, ** for 5% and * for 10%.*

In a first step, we analyse the specific results at the biannual frequency in the lower block. Our seasonal modelling gives a long-run price elasticity equal to -1.95, and in absolute value, greater than values obtained without including seasonal effects (i.e. -0.07 and -0.37 respectively for the lower and upper blocks). This means that the seasonal component of residential water consumption in the lower consumption block is very sensitive to the corresponding seasonal price fluctuations, confirming higher price elasticity during the dry season compared to the wet season. Obviously, the value of -1.95 denotes an unusually high elastic demand to price. However, in a developing country such as Tunisia, water expenditure probably constitutes a large proportion of these consumers' total expenditures. Therefore, a seasonal pricing with a higher tariff in summer to achieve water conservation objectives is excluded as the corresponding consumption block only includes low consumption levels. This significant result suggests that long-term, alternative water management policies should be implemented. These could include the promotion of low-water consuming equipment or public assistance for low income households to help them pay their water bills.

The introduction of seasonality reveals another interesting and specific result for the lower block. We observe a positive seasonal effect of the percentage of subscribers on consumption while there is a negative long-run effect at the zero frequency. This positive seasonal effect (with an estimated coefficient equal to 2.5) can be explained by a slide down of certain consumers from the higher consumption block to the lower one in winter. This slide increases the number of consumers in the lower block and increases the corresponding average consumption level because these new subscribers have higher average consumption levels.

Lastly, our results show that some important information for water managers may be lost when information concerning seasonal fluctuations is ignored. As some of the usual lower block consumers are obliged to increase their consumption in summer, and then to switch from the lower to the upper block, we propose increasing the length of the lower block, at least in summer. Such a seasonal tariff policy would guarantee the satisfaction of basic water consumption needs at the lowest price in every season.

In a second step, we analyse results concerning the two cointegrating relations at the zero frequency, which deals with the traditional long-run effects. Here, the variable measuring the proportion of subscribers can take into account the specific characteristics of a developing country in which the distribution network is expanding rapidly. Estimates show that a 1% increase in the number of newly connected households, generally characterized by a low level of equipment and water consumption, will reduce the average consumption of the lower block by 0.44%. This network expansion variable is not statistically significant for the upper block.

The two long-run price elasticities are statistically significant and greater in the upper block (-0.37) than in the lower block (-0.07). These values, similar to the values usually obtained in the literature for developed countries, therefore confirm the greater price sensitivity of consumption which is not fulfilling basic needs. They also confirm the relevance, in terms of water conservation, of the super-progressive water pricing applied in Tunisia. This price sensitivity for the upper block suggests that an increase in the highest tariff can be proposed to reduce excessive water consumption. However, Whitthington (1992) suggests that increasing block tariffs in developing countries may have adverse effects such as reducing the well-being of large families. Therefore, an appropriate pricing policy aiming to reduce only excessive water consumption must rely on a tariff which takes family size into account.

Coefficients on precipitation are statistically significant for the lower block and at the zero frequency only. Indeed, over the long run rainfall has the traditional negative impact on water consumption. Temperature has the expected positive sign, however, the climate effects do not result from seasonal fluctuations. Lastly, the long run income elasticity estimate is around 0.2 in the two blocks. In Tunisia, where droughts are likely to worsen with time and where average income is expected to rise, these results leads us to expect increases in residential water consumption over the long run. A seasonal error correction model (SECM) is considered in the next subsection to analyze short-run effects.

Error correction model

A seasonal Error Correction Model (SECM) allows the speed of adjustment in residential water consumption to be determined. Indeed, the coefficients of the lagged residual terms measure the speed rate at which the consumption corrects short-run deviations in temperature, rainfall, price and income. Since the variables in our study are not cointegrated at the annual frequency, we drop the error correction terms w_{t-2} and w_{t-3}. Furthermore, we derive the short-run elasticities from the estimates of the corresponding variable. Table 8 gives the results for the two SECM models.

Table 8. Estimation of the SECM by OLS

Price	$Cons_{t-1}$	Income	Rain	% custom	Temp.	u_{t-1}	v_{t-1}	R^2	DW
				Lower block					
-0.004	-0.26***	0.08	-0.01	-0.88**	0.09	-0.35***	0.010	0.32	1.96
(-0.03)	(-2.60)	(0.28)	(-0.66)	(-2.20)	(0.85)	(-3.25)	(0.67)		
				Upper block					
-0.01	-0.28**	0.23	-0.04*	-0.01	0.07	-0.19***		0.17	2.00
(-0.10)	(-2.4)	(0.49)	(-1.98)	(-1.20)	(0.53)	(-3.17)			

*Notes: Figures in brackets are t-statistics of parameter estimates. Significance level for parameter estimates: *** for 1%, ** for 5% and * for 10%. DW denotes the Durbin Watson statistic.*

Lagged values of the residuals from the cointegrating relation at the zero frequency, i.e. u_{t-1}, are statistically significant in both blocks. As the estimates are negative, this implies that adjustments will cause the system to gradually converge towards the equilibrium. In the lower consumption block, three quarters (i.e. 1/0.35) are needed for the average water consumption to return to its initial equilibrium level following a shock (five quarters (1/0.19) in the upper block).

Conversely, the coefficient of the biannual cointegration error correction term v_{t-1} is not statistically significant. The results imply that following seasonal shocks, a short-run adjustment of low-level water consumption to price, climate and network fluctuations will not occur.

Finally, the short-run estimates of price and income elasticities are not statistically significant. This confirms that the long-run price mechanism exceeds short-run effects. A single significant result furthermore reveals a short-run negative impact of rainfall on water consumption. Globally, however, such a lack of dynamics is common in a seasonal error correction model where the dynamics usually are fully captured by the error correction terms (see Moosa, 1996 or Ouerfelli, 2008 for example).

VI. Conclusion

The main purpose of this article is to adequately model seasonality in the residential water demand function using quarterly aggregate time series data. In this paper, seasonal modelling is tested explicitly rather than assumed. Thus, using data from Tunisia, our main contribution is twofold.

In a first step, tests for seasonal unit roots show the presence of unit roots at the zero and biannual frequencies for the variables included in the residential water demand function. This is an indication of varying stochastic seasonal patterns. In a second step, we implement a cointegration analysis by testing the relevance of periodic *versus* seasonal cointegration. We then use the procedure developed by Franses (1993) to verify our chosen empirical methodology.

Finally, our main results from the seasonal cointegration tests show that lower block residential water consumption is cointegrated with average price, income and climate variables at the biannual and zero frequencies. However, if we consider the upper block, the cointegration relation is statistically significant at the zero frequency only, and not at the biannual frequency, implying that there is no long-run equilibrium in this submarket. This means that in a developing country like Tunisia, greater levels of consumption, depending on water using equipment and household habits, are less stable. In addition, the seasonal error correction model does confirm the existence of short run adjustment in the water consumption behaviour, but not following seasonal shocks.

Finally, our basic findings are that seasonality can play a significant role in modelling residential water demand. All in all, this study will enable the best water conservation policy to be proposed, including the effect of seasonality.

First, as some lower block consumers are obliged to increase their consumption in summer, and thus to switch from the lower to the upper block, we propose to increase the size of the lower block, at least in summer. Such a seasonal tariff policy would guarantee the satisfaction of basic water consumption at the lower price in every season.

Next, as the long-run price elasticity is positive and significant in the upper block, this leads us to believe that raising prices in the upper block can result in a decrease in the water consumption of well-to-do people. However, the application of such an incentive pricing policy would conduct to a loss in the welfare of the large families if the size of the family and their income level is not taken into account. This modified pricing scheme will help to achieve goals of environmental protection and social equity (see Porcher, 2014 for further discussions about efficiency and equity in water use).

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Appendix

Table 4. HEGY seasonal unit roots results

VARIABLES	Deterministic component	***'t' : α₁***	***'t' : α₂***	***'t' : α₃***	***'t' : α₄***	***'F' :*** $\alpha_3 \cap \alpha_4$
	None	0.63	0.23	-3.25*	2.63*	9.69*
Consumption	I	-1.97	0.21	-3.15*	2.58*	9.16*
Lower block	I,SD	-1.89	-0.70	-3.61**	3.36*	14.18*
	I,Td	-2.72	0.20	-3.16*	2.44*	8.79*
	I, SD, Td	-2.49	-0.70	-3.63**	3.13*	13.36*
	None	-1.02	-0.45	-2.80*	1.72**	5.70*
	I	-2.00	-0.45	-2.76*	1.90**	5.96*
Consumption	I,SD	-2.05	-0.93	-3.77**	1.98**	9.73*
Upper block	I,Td	-3.41	-0.52	-2.60*	2.17**	6.16*
	I,SD,Td	-3.39	-1.00	-3.54**	2.34**	9.86*
Average price	None	-2.02**	0.53	-2.19**	0.97	2.94
Lower block	I	-1.12	0.53	-2.18**	0.96	2.92
	I,SD	-1.08	-0.46	-3.40	2.18**	8.78*

	I,Td	-1.64	0.43	-2.13**	0.91	2.75
	I,SD,Td	-1.60	-0.54	-3.36	2.14**	8.52
Average price	None	-3.68*	-1.05	-4.19*	0.19	8.83*
Upper block	I	-3.21**	-1.06	-4.23*	0.28	9.04*
	I,SD	-3.14**	-1.45	-4.56*	0.07	10.42*
	I,Td	-1.05	-1.06	-4.21*	0.30	8.96*
	I,SD,Td	-1.02	-1.45	-4.53*	0.09	10.32*
Income	None	2.94*		-1.32	-1.33	1.80
	I	3.65*		-1.02	-1.23	1.30
	I,SD	3.21**		-2.40	-2.53**	6.55
	I,Td	0.20		-1.04	-1.21	1.30
	I,SD,Td	-0.04		-2.47	-2.47**	6.61
% subscribers	None	-0.57	-1.75	-1.92**	0.59	2.03
Lower block	I	-1.72	-1.73	-1.93**	0.60	2.05
	I, SD	-1.70	-1.65	-1.40	1.31	1.81
	I,Td	-3.33	-1.63	-2.01**	0.63	2.21
	I,SD,Td	-3.33	-1.56	-1.46	1.40	2.03
% subscribers	None	1.01	-1.53	-1.77	0.43	1.66

Upper block	I	-1.19	-1.53	-1.75	0.45	1.63
	I,SD	-1.18	-1.77	-1.57	0.93	1.65
	I,Td	-2.41	-1.46	-1.78	0.50	1.71
	I,SD,Td	-2.44	-1.70	-1.63	1.06	1.87
Rainfall	None	-0.90	-1.94**	-2.67*	-2.55*	7.53*
	I	-2.08	-3.16*	-4.38*	-1.56	10.77*
	I,SD	-2.00	-3.54**	-5.29*	-1.99**	15.86*
	I,Td	-2.90	-2.02	-3.00*	-2.42*	8.23*
	I,SD,Td	-3.02	-2.17	-3.79**	-2.68**	12.31*
Temperature	None	0.64	-3.44*	-0.87	-0.21	0.40
	I	-2.7	-3.37*	-0.8	-0.16	0.33
	I,SD	-2.58	-3.35**	-3.91*	-0.98	8.34**
	I,Td	-3.31	-3.41**	-3.85*	-0.142	0.37
	I,SD,Td	-3.002	-3.07**	-3.83**	-0.6	7.57**

Notes: Critical values available in HEGY (1990), Tables 1a and 1b, pages 226, 227. (resp. **) if we reject the unit root hypothesis at the 1% level (resp. 5%).*