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A SEASONAL INTEGRATION AND COINTEGRATION ANALYSIS OF RESIDENTIAL WATER DEMAND IN TUNISIA

Marie-Estelle Binet & Younes Ben Zaïed

University of Rennes 1, CREM-CNRS

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A SEASONAL INTEGRATION AND COINTEGRATION ANALYSIS OF RESIDENTIAL WATER DEMAND IN TUNISIA

Younes BEN ZAIED & Marie Estelle BINET

CREM (UMR6211 CNRS), University of Rennes 1, France

Abstract

The main originality of this article is to empirically incorporate the effect of seasonality in estimating the residential water demand function. We use quarterly times series for the period 1980.1 to 2007.4 from Tunisia and a two consumption blocks decomposition (in a lower and in an upper blocks). As the Error Correction model is poorly significant, we obtain a long-run price elasticity for the upper block equals to -0.39 and greater than the corresponding short run elasticity which is not significant. Therefore, we are able to advocate policies for the upper block both in favor of a water management through pricing and promoting the adoption of water saving equipments. The introduction of seasonality claims for new insights concerning water conservation policies as evidence in favor of seasonal cointegration at biannual frequency is found for the two blocks. Results show that a seasonal pricing policy will not be efficient to reduce the upper block consumption. But, as a part of the consumers switch from the lower to the upper block in summer, we propose to increase the length of the lower block to ensure the satisfaction of households' essential needs in all seasons.

Key-Words: Residential water demand, seasonal cointegration, seasonal error correction model.

1. INTRODUCTION

Residential water demand has been a major issue in environmental economics as proved by the number of recent surveys available in the literature (Arbuès et al (2003), Dalhuisen et al (2003), Worthington and Hoffman (2008)). Most of these researches has been conducted in developed countries, but there also exist some studies for developing countries, see Nauges and Whittington (2010).

Many studies have focused on the implementation of a water management through pricing. In the related literature, a few papers have distinguished the long-run from the short-run effects, using two different empirical methodologies. Nauges and Thomas (2003) have estimated a dynamic panel data model on a sample of French municipalities and have obtained short and long-run price elasticities respectively equal to -0.26 and -0.40. Using times series observations from Seville in Spain, Martinez-Espineira (2007) has derived the long-run price elasticity equals to -0.5 from a cointegration model and the short-run price elasticity equals to -0.1 from an error correction specification. As short-run price elasticities are smaller than their long-run counterparts, authors conclude that consumers might need time to adjust water-using capital stocks or to learn about the effects of their consumption on their bill. If it is the case, tariff policies are more efficient in the long-run.

But, to our knowledge, no studies have integrated seasonal fluctuations to analyze water demand determinants. The modification in habits (with the concentration of holidays in summer) and the effect of climate fluctuations imply that aggregate residential water consumption probably follows seasonal fluctuations. Households are expected to consume more in summer and less in winter. Therefore, the role played by seasonality is one important issue that has been neglected in the literature, probably because of the lack of data. And, seasonal fluctuations could be an important source of variation in residential water consumption and if it the case, adequate water management policies must rely on. To our knowledge, Martinez-Espineira (2007) has been the first and the only one to test seasonal co-integration and error-correction models to distinguish short and long run price elasticities for the case of monthly data. But he did not detect seasonal unit roots.

In the light of these findings, this article aims at expending the existing literature in two ways. First of all, the aim of this paper is to demonstrate that seasonality can play a significant role in modeling residential water demand. The second originality of this article is to analyze the household water consumption in Tunisia. To our knowledge, only one unpublished paper by Ayadi et al (2002) has developed the first demand estimation for Tunisian residential water using quarterly data from 1980.1 to 1996.4. But their methodology is different as they estimate a system of two equations: a demand equation explaining the quantity of water consumed per household in each bracket and a second equation explaining the proportion of households in each bracket.

This article is therefore an original contribution to the empirical residential water modeling as we propose for the first time a seasonal cointegration analysis of residential water demand using a rich quarterly data set for Tunisia. Based on a two consumption blocks decomposition (in a lower and a upper blocks), our data base consists of quarterly values of consumption, average price, rainfall, the number of domestic consumers in each block and yearly values for income.

The first step of our work is to analyze the data and conduct the Hylleberg, Engle, Granger and Yoo (HEGY 1990) seasonal unit roots tests at zero, annual and biannual frequencies. Then, we study seasonal cointegration using the Engle, Granger, Hylleberg and Lee (EGHL, 1993). Cointegration at the long-run frequency can be interpreted as indication of a parallel long-run movement in the nonstationary series whereas cointegration at a seasonal frequency can be interpreted as evidence for a parallel movement in the seasonal component of the series which both exhibit a varying seasonal pattern.

Our basic findings are that we observe cointegration at both biannual and zero frequencies for the lower block, and at biannual frequency only for the upper block. Our results show that pricing variations appear to have effects on water upper block consumption for long-run movements, as the long-run price elasticity is significant. But an appropriate seasonal pricing will not help to conduct to a reduction of water consumption as the corresponding price elasticity is not significant. Other results derived from the introduction of seasonality suggest that authorities should increase the length of the lower block to ensure the satisfaction of households' essential needs in all seasons.

The paper is organized as follows. In the first section we present our data set for Tunisia between 1980.1 to 2007.4. Then, the empirical methodology is developed in section 2. Finally section 3 presents and discuss the main empirical results obtained. Section 4 concludes.

2 DATA SET DESCRIPTION

Water resources in Tunisia are characterized by scarcity, quality problems, bad distribution, as well as time and space volatility. Even if residential water consumption is limited compared to irrigation demand which monopolizes more than 80% of total resources, it must be carefully managed for at least three reasons.

First, the available water resources in Tunisia are calculated to be 420 m^3 per quarter per household in 2005. And, according to the World Bank report, this value will be 300 m^3 in 2030. So Tunisia suffers a real water supply crisis which will be accentuated during the next two decades. Secondly, residential water consumption, which concerns the satisfaction of essential human uses (drinking, cooking and basic hygienic purposes), requires a minimum of regularity, quality (softness, purity etc.) and reliability especially during the dry season, which is not always the case in Tunisia. And, residential water demand is really exponentially increasing as a result of a rapid urban development. Third, Tunisia is committed to manage water uses like other developing countries to boost her frail economy where tourism development requires more water with acceptable quality.

Therefore, if Tunisia don't want to resort to non conventional and costly resources (such as desalination), the only alternative is to rely on appropriate water demand management. Therefore, water pricing must be considered seriously as a useful tool, with the other non-pricing instruments, such as water-saving equipments, awareness, education and participatory management, to keep under control the demand evolution.

We use a rich and original data base covering the period going from the first quarter of 1980 to the fourth quarter 2007. The data, collected by SONEDE (The national water distribution company), includes quarterly observations on average domestic water consumption, average price, network expansion, rainfall and yearly household income observations.

Since Tunisia, as many countries, uses a nonlinear tariff structure in which prices are differentiated for different brackets of consumption, the choice of the price variable (average or marginal prices) is necessary to achieve a good residential water demand specification. Following Ayadi et al (2002), we choose the average price equal to the total bill of the households divided by the volume consumed, as we have semi aggregate data. The average price is a weighted sum of the marginal prices, with the weights being given by the shares of the consumption in each bracket. Therefore, demand specification for residential water usually has to deal with the average price endogeneity. One nice feature of cointegration and error correction models is that estimates are not implemented with instrumental variable estimators but with OLS.

Indeed, SONEDE has built a non linear pricing in all the country using five brackets. But Ayadi et al. (2002) who have conducted their empirical work on the same sample data set have shown that the best choice is to conduct estimations on a two blocks decomposition (a lower and an upper blocks). The lower block will put together the consumers of the first two brackets (0-40 m^3) while the upper block gathers the latest three brackets (more than 41 m^3). And, in a developing country as Tunisia, it is important to estimate one specific residential water demand equation for each block if we want to control water demand as efficiently and fair as possible. Indeed, we have to implement different policies for each block. On one hand, it is important to analyze the effect of price variations on demand for the upper block. On the other hand, marginal price and the length bracket of the lower block should guarantee the satisfaction of the essential needs of the low income households. Therefore, every pricing policy must take care of these different objectives.

Next, annual data on income, derived from budget surveys compiled by the National Statistical Institute, has also been collected. Therefore, the seasonal analysis of integration and cointegration will rely on a residential demand specification without income effect.

Network expansion is an appropriate variable to take into account the specific characteristics of a developing country in which the distribution network is quickly expanding. It measures the effect of new entrants to the network as a result of economic development or seasonal variations in consumption. If the average consumption of new entrants in one block is lower than that of existing consumers, we expect a negative coefficient for the network effect.

Table 1 gives a description of the variables and basic descriptive statistics:

Table 1. Description of the variables and basic descriptive statistics, 1980.1 to 2007.4

Variable	Description	Mean	Max	Min
Residential consumption (lower block, m^3)	Quarterly data for average water consumption equal to the sum of consumption in the two first blocks divided by the corresponding number of households.	19.86	40	9.04
Price (lower block)	Average price equal to total water bill divided by the total volume of water consumed in the two first blocks.	0.39	0.85	0.20
Residential consumption (upper block, m^3)	Quarterly data for average water consumption equal to the sum of consumption in the three latest blocks divided by the corresponding number of households.	150.61	341.50	54.11
Price (upper block)	Average price equal to total water bill divided by the total volume of water consumed in the three latest blocks.	0.75	1.33	0.23
Yearly income, Dinars	Built from the expenditure surveys by National Statistics Institute.	1570	2549.50	1218
Rainfall	Average quarterly level of precipitations (ml/quarter)	172	600.71	10.86
Network expansion (lower block)	Quarterly share of subscribers to the lower block (%)	73	85	55
Network expansion (upper block)	Quarterly share of subscribers to the upper block (%)	9	22	4

All variables were collected by SONEDE.

In average, the lower block represents 73% of subscribers and 53% of total domestic consumption. In average, the upper block accounts for 9% of subscribers and 47% of total domestic consumption. In Tunisia, the average yearly income is 1510 dinars which corresponds to 755 Euros.

Table 2 presents some aggregate statistics about quarterly fluctuations of the variables.

Table 2. Average quarterly values

Variables	Winter	Spring	Summer	Autumn
Lower block consumption	22.50	16.84	24.63	15.48
Upper block consumption	176	116.72	182	128
Lower block average price	0.37	0.41	0.34	0.43
Upper block average price	0.79	0.68	0.81	0.72
Network expansion lower block (%)	79	69	67	76
Network expansion upper block (%)	7	11	12	7
Rainfall	196	154.41	150.52	187.30

These figures confirm that most of the variables are seasonal in nature. Seasonal effects are expected to be more important in the upper block. Indeed, we observe an important intra annual variation both in the volume of consumption and in the number of consumers in the upper block. More generally, we observe, a low level of rainfall in winter. Such seasonal fluctuations show that total demand in a year is not uniformly distributed across seasons. All in all, the seasonal nature of the data can be taken into account using a seasonal integration and cointegration approaches as developed in the next section.

3. EMPIRICAL METHODOLOGY

Co-integration theory, carried out for the first time by Engle and Granger (1987), allows the estimation of long run relationship between non stationary variables and requires the testing for stationarity of the series as a first step. The use of seasonal time series in our study requires an extension of time series unit root tests. Hylleberg, Engle, Granger and Yoo (Hegy 1990) have extended these methods to deal with the seasonal frequency in quarterly times series. Then, Engle, Granger, Hylleberg and Lee (EGHL, 1993) have extended the cointegration techniques to the case where the data have unit roots at both zero and seasonal frequencies. Logged data are used throughout the analysis below.

3.1 Tests for seasonal integration

Testing for unit root has been considered as the first step in econometric time series analysis. We apply the Hylleberg, Engle, Granger, and Yoo (1990) method which allows testing for seasonal unit roots at different frequencies. Indeed this procedure consists in running the following OLS estimation for quaterly times series *y*:

$$(1 - B_4)y_t = \alpha_1\, y_{1,t-1} + \alpha_2\, y_{2,t-1} + \alpha_3\, y_{3,t-2} + \alpha_4\, y_{3,t-1} + \sum_{i=1}^{p} \theta_i \Delta_4 y_{t-i} + \mu_t + \varepsilon_t \qquad (1)$$

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

$\varphi_1(B) = (1 + B + B^2 + B^3)$, $\varphi_2(B) = -(1 - B + B^2 - B^3)$ and

$\varphi_3(B) = -(1 - B^2)$, (B is the lag operator i.e. (1-B_4)y_t = $y_t - y_{t-4}$)

Note that the deterministic component μ_t is added in the regression to include seasonal dummies (SD), linear time trend (Td) and a constant term (I). The term ε_t is a normally and independently distributed error term (i.e.$\varepsilon_t \sim NID(0, \sigma^2)$).

The regression (1) is augmented by additional significant lagged values of the dependent variable to whiten the residuals. The lag length selection is based on the selection of the latest significant lag.

The regression (1) is estimated by OLS and the ratio statistics of the estimated coefficients will be used to test for seasonal unit root at zero frequency, biannual frequency, and annual frequency.

To demonstrate that y_t has no unit root, we should perform the following significance tests: $H_{01:}$ α_1=0, $H_{02:}$ α_2=0, $H_{03:}$ α_3=0, $H_{04:}$ α_4=0, and H_{03+04}: α_3= α_4=0 against the alternative where the coefficient are statistically different from zero. We reject the hypothesis of a seasonal unit root if α_2 and either α_3 or α_4 are statistically different from zero. Using Monte Carlo simulation HEGY (1990) provide critical values for the different significance tests presented above.

3.2 Cointegration tests

Testing for seasonal cointegration, which represents an extension of the Engle and Granger (1987) cointegration theory, allows for cointegration at all possible frequencies. According to Engle, Granger, Hylleberg and Lee (1993), Cointegration at $0, \frac{1}{2}\ and\ \frac{1}{4}$ frequencies is established if the residuals terms u_t, v_t, w_t are stationary:

$$y_{1t} = \beta_0 + \beta_1 \varphi_1(B)x_t + u_t \quad (2)$$

$$y_{2t} = \gamma_0 + \gamma_1 \varphi_2(B)x_t + v_t \quad (3)$$

$$y_{3t} = \delta_0 + \delta_1 \varphi_3(B)x_t + \delta_2 \varphi_3(B)x_{t-1} + w_t \quad (4)$$

Where y_{it} and $\varphi_i(B)$ are the same as described above in the previous section and x_t is a set of the other variables.

More precisely, testing for stationarity is performed on the basis of the auxiliary regressions;

$$\Delta u_t = \pi\ u_{t-1} + \sum_{i=1}^{p} \theta_i \Delta\, u_{t-i} + \mu_{1t} \quad (5)$$

$$v_t + v_{t-1} = \pi\, v_{t-1} + \sum_{i=1}^{p} \theta_i\,(v_{t-i} + v_{t-1-i}) + \mu_{2t} \quad (6)$$

$$w_t + w_{t-2} = \tau_1(-w_{t-2}) + \tau_2(-w_{t-1}) + \sum_{i=1}^{p} \theta_i\,(w_{t-i} + w_{t-2-i}) + \mu_{3t} \quad (7)$$

We reject the seasonal co integration hypothesis at 0 or biannual frequency if the t statistics are smaller in absolute value than the critical value tabulated by Engle and Yoo (1987).

To test for seasonal co integration at $\frac{1}{4}$ frequency ($annual\ frequency$), the t statistics of $\tau_1\ and\ \tau_2$ are used with the joint test F statistics of $\tau_1 \cap \tau_2$, and the critical values are tabulated in Engle et al (1993) using Monte Carlo simulations.

3.3 Error Correction Model

We can turn to specify and estimate short run effects through the following general error correction model;

$$\Delta_4 y_{1t} = \sum_{i=1}^{q} \alpha_i\, \Delta_4 y_{1,t-i} + \sum_{i=0}^{q} \beta_i\, \Delta_4 y_{2,t-i} + \delta_1 u_{t-1} + \delta_2 v_{t-1} + \delta_3 w_{t-2} + \delta_4 w_{t-3} + \varepsilon_t$$

This equation can be estimated by OLS if all the terms are stationary. If the cointegration residuals at different frequencies u_t, v_t, w_t are significant, the size of the related coefficients measures the speed at which the variables adjust to restore the water consumption equilibrium.

4. EMPIRICAL RESULTS

After analyzing unit root properties of all the variables, we study the seasonal cointegration in residential water consumption at all possible frequencies, and for each consumption block. Then, a seasonal error correction model is estimated by OLS for the lower and the upper blocks.

4.1 Tests for seasonal integration

The outcomes of the HEGY tests are reported in table 1 (in appendix). Unit root tests on income are not implemented at the biannual frequency as the variable is on a yearly basis. The Breusch Godfrey LM test is used to test for residuals autocorrelation.

Our results confirm that seasonality in water consumption is the rule and stationarity is an exception. Indeed, we find that all the variables (in logs) are not stationary at zero frequency with robustness to the inclusion of deterministic components, with one exception on the upper bloc average price which appears to be stationary when we include intercept and seasonal dummies.

The main finding is the seasonal integration of water consumption and prices, at the lower and at the upper blocks, and network extension at the biannual frequency. This can be explained by different water consumption patterns at different seasons as it is an indication of varying stochastic seasonal patterns.

The presence of unit roots at the biannual frequency for the quarterly rainfall variable is sensitive to inclusion of the trend. Unit roots at biannual frequency are present when we include a trend whereas no unit roots are found at the biannual frequency when we use intercept or/and seasonal dummies. This implies that the seasonal components of rainfall are deterministic rather than stochastic.

The seasonal integration at annual frequency has been proved only on the income and the two bloc's network extension. The joint test is rejected on most cases implying that there are not unit roots with an annual frequency.

4.2 Seasonal cointegration test

The evidence of the presence of unit roots in the variables at the zero frequency, leads to the examination whether we have a long run relationship at zero frequency between the average water consumption and its usual determinants (average price, income, network expansion and rainfall). If price or rainfall are neutral in the long-run, this could mean that the unit root present in each series should not be a common one. Turning to the case of seasonal cointegration, source of within-year effects including the habits of water consumption as well as seasonality in climate, we estimate the same relation without the income variable.

The procedure developed by Engle et al (1993) is used to test for seasonal cointegration in the lower and the upper block separately. Results are reported in table 3 and 4.

Table 3: Testing for seasonal cointegration at zero and biannual frequencies, lower block (t-statistics are into parenthesis)

	Price	Income	Network	Rainfall	DW	R^2	$t_{\hat{\pi}}$
Zero frequency	-0.15 (-4.81)***	0.4 (44.6)***	-0.12 (-0.46)	-0.02 (-2.96)***	0.54	0.6	-3.83*

biannual frequency	-1.94 (-12.34)***	_______	2.54 (3.60)***	-0.03 (-0.60)	3.82	0.77	-4.12**

All the variables are in natural logarithm, the coefficient significance statistics are in parenthesis, critical values at 5 percent and 10 percent for N=4 and t=100 are respectively -4,02 and -3,71 from Engle and Yoo (1987). DW denotes the Durbin Watson statistic.

Table 4: Testing for seasonal cointegration at zero and biannual frequencies, upper block

	Price	Income	Network	Rainfall	DW	R^2	$t_{\hat{\pi}}$
Zero frequency	-0.39 (-7.45)***	0.59 (13.24)***	-0.09 (-0.82)	0.04 (2.73)***	0.27	0.31	-2.30
biannual frequency	-0.010 (-0.21)	_______	-3.42 (-18.08)***	-0.09 (-1.32)	3.22	0.75	-7.02**

All the variables are in natural logarithm, the coefficient significance statistics are in parenthesis, critical values at 5 percent and 10 percent for N=4 and t=100 are respectively -4,02 and -3,71 from Engle and Yoo (1987).

Results show that the null hypothesis of the absence of seasonal cointegration at zero frequency is not rejected if we consider the upper bloc and is rejected for the lower block, at 10% significance level only. This implies long run equilibrium among residential water consumption and its determinants for the lower block only. Indeed, residential consumption behaviors are stable in the lower block, where consumers satisfy their essential needs. At the opposite, in a developing country as Tunisia, greater levels of consumption, depending on water using equipments and household habits are less stable.

Long-run elasticities, which result from the choice of both the size of the capital stock and its use, are measured by the coefficients of the price variables in these two cointegrating equations. Long-run price elasticities are significant and equal to -0.15 in the lower block and -0.39 in the upper block. In the lower block, most part of the demand satisfies basic needs, and therefore is usually inelastic to price. This part of water can be considered to be an essential good with very low price elasticity. The value obtained for the upper block is similar to values usually obtained in the literature for developed countries. This price sensitivity for the upper block suggests that households do not respond immediately to price variations.

Therefore, an increase in the corresponding marginal prices can be proposed to reduce water consumption in the long-run.

In the long run, rainfall has the traditional negative impact on water consumption for the lower block only. Ayadi et al (2002) have also obtained a positive effect of rainfall in the upper block. They explain it by an increase in average consumption in the wet/cold seasons for the upper block due to a sliding down of some consumers from the upper block (those with a lower average consumption) to the lower block, thus leaving the upper block with a greater proportion of customers with a relatively higher average consumption. The last result is that network expansion is never significant in the long run.

The main originality of our results, compared to those obtained by Martinez Espineira (2007), is that we show that seasonality can influence the chain of causation between water consumption and its determinants. Indeed, results reveal the existence of seasonal cointegration at the biannual frequency for the two blocks. These results suggest that seasonal variations in residential water consumption may be a reflection of seasonal fluctuations in price, income and network expansion, but not in rainfall which is not significant. The results indicate contrasted values of price elasticities according to the consumption block considered. The price elasticity for the lower block, equals to -1.94 denotes a high price elastic water demand. The economic interpretation of such a cointegration relation is that of a different sensitivity to price in summer or in winter, but only for the low level water consumers. Indeed, the corresponding price elasticity is not significant for the upper block.

Next, we observe a positive seasonal effect of network variable on consumption in the lower block and a negative one for the upper block. The positive effect can be explained by a sliding down from a higher consumption bracket to a lower one in winter, reducing the number of consumers in the high bracket, and increasing the average consumption level in the lower block. On the contrary, in summer, the average consumption of new entrants in the upper block is lower than that of existing consumers, and we observe a negative coefficient for the network effect.

So, results show that we could lose some important information for water managers by ignoring information concerning seasonal fluctuations. Indeed, we can propose three policy implications of our results. As a part of usual lower block consumers is constrained to increase their consumption in summer, and then to switch from the lower to the upper block, we propose to increase the length of the lower block, at least in summer and spring. Such a seasonal tariff policy would guarantee the satisfaction of basic water consumption at the lowest price, in every season.

Next, as the long-run price elasticity is positive and significant in the upper block, we can increase the corresponding marginal prices. But, the application of an appropriate seasonal pricing based on a higher price in winter, would not conduct to a significant reduction in water consumption in this consumption block.

Next, our significant results for cointegration analysis, suggest to implement long-run water management policies, such as information campaigns to help people to modify their consumption habits and promotion of low-water consuming equipments.

4.3 Error correction model

The seasonal error correction model (SECM) is the second step of the cointegration procedure. The SECM is useful to determine the speed of adjustment in residential water consumption. Indeed, the coefficients of the residual terms measure the speed rate at which the consumption corrects short-run deviations in rainfall, price or income. Furthermore, short-run price and income elasticities are derived from the estimates of the corresponding coefficient. As cointegration has been found for zero and biannual frequencies for the lower block and at biannual frequency only for the upper block, table 5 gives the results for the two ECM models:

Table 5. Estimation of the ECM by OLS

	$price_{-1}$	$Cons_{-1}$	$Income_{-1}$	$Rain_{-1}$	$Network_{-1}$	$EC_{1,t-1}$	$EC_{2,t-1}$	R^2	DW
Lower block consumption	-0.004 (-0.03)	-0.32 (-2.65)***	0.034 (0.11)	-0.03 (-1.78)*	-0.69 (-1.83)*	-0.21 (-2.60)***	0.010 (0.39)	0.28	2.01
Upper block consumption	-0.09 (-0.66)	-0.32 (-2.84)***	0.06 (0.17)	-0.004 (-0.84)	0.25 (2.03)***	------	0.002 (0.09)	0.17	2.00

T-statistics for regression coefficients are reported in parentheses. All the variables are in Δ logs.

Only the effect of the coefficient of the cointegration error correction term EC_1 at zero frequency is significant for the lower block. As the speed of adjustment is negative, this implies that adjustments will cause the system to gradually converge towards the equilibrium. Conversely, the coefficients of the biannual cointegration error correction terms EC_2 are both insignificant. The results imply that short run adjustment of consumption to price and income and network fluctuations in summer do not occur.

Error correction models give respectively short-run estimates of the price and income elasticities. Short-run elasticities depend only on the intensity of the use of the water using capital stock whereas long-run effects result from the choice of both the size of the capital stock and its use. As usually, we find short-run price elasticities smaller than their long-run counterparts as they are insignificant. Indeed, this confirms that consumers might need time to adjust water-using capital stocks or to learn about the effects of their consumption on their bill. Furthermore, as short-run price adjustment does not occur, public policy should also subsidy water saving equipments.

5. CONCLUSION

The main purpose of this article is to analyze the impact of seasonality in estimating the residential water demand function. Using quarterly data from Tunisia, our contribution is twofold. In a first step, tests for unit roots show the presence of unit roots at the zero and biannual frequencies for all the variables. Findings of this paper reveal that residential water consumption have seasonal components. So, go further into the knowledge of residential water demand determinants, we show that lower block residential water consumption is cointegrated with price and income at the biannual and zero frequencies. But the absence of cointegration at the zero frequency for the upper consumption block implies there is no long run equilibrium in this submarket. In addition, the seasonal error correction model does not confirm the existence of short run adjustment in the water consumption behaviors. All in all, this study will enable to propose the best water conservation policy, including the effect of seasonality.

As a part of usual lower block consumers is constrained to increase their consumption in summer, and then to switch from the lower to the upper block, we propose to increase the length of the lower block, at least in summer and spring. 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, we can increase the corresponding marginal prices. But, the application of an appropriate seasonal pricing based on a higher marginal price in winter, would not conduct to a significant reduction in water consumption in this consumption block.

Then, our significant results for cointegration analysis suggest to implement long-run water management policies, such as information campaigns to help people to modify their consumption habits and promotion of low-water consuming equipments.

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APPENDIX T*able 1: Testing for seasonal integration using HEGY (1990) procedure.*

'*' indicate the rejection of the null hypothesis at the 5% significance level based on the critical values simulated by HEGY [1990] using Monte Carlo simulation.

VARIABLES	REGRESSIONS	't' : α_1 ***0 Frequency***	't' : α_2 ***biannual***	't' : α_3 ***annual***	't' : α_4	'F' :$\alpha_3 \cap \alpha_4$	LM	Lag
Lower block consumption	None	0.63	0.23	-3.25*	2.63	9.69*	3.24	4
	I	-1.97	0.21	-3.15*	2.58	9.16*	3.11 1.70*	4 4
	I,SD	-1.89	-0.70	-3.61*	3.36	14.18*	3.30	4
	I,Td	-2.72	0.20	-3.16*	2.44	8.79*	1.92*	4
	I, SD, Tr	-2.49	-0.70	-3.63*	3.13	13.36*	3.25	6
Upper block consumption	None	-1.02	-0.45	-2.80*	1.72	5.70*	2.86	6
	I	-2.00	-0.45	-2.76*	1.90	5.96*	1.86*	6
	I,SD	-2.05	-0.93	-3.77*	1.98	9.73*	1.71* 0.79*	6 6
	I,Td	-3.41	-0.52	-2.60*	2.17	6.16*	2.42	8
	I,SD,Tr	-3.39	-1.00	-3.54*	2.34	9.86*	2.41	8
Lower block price	None	-2.02*	0.53	-2.19*	0.97*	2.94	1.10*	8
	I	-1.12	0.53	-2.18*	0.96*	2.92	2.62	8
	I,SD	-1.08	-0.46	-3.40	2.18	8.78*	1.32*	8
	I,Td	-1.64	0.43	-2.13*	0.91*	2.75	3.68	7
	I,SD,Tr	-1.60	-0.54	-3.36	2.14	8.52*	3.49	7
Upper block price	None	-3.68*	-1.05	-4.19*	0.19*	8.83*	3.05	7
	I	-3.21*	-1.06	-4.23*	0.28*	9.04*	3.58	7
	I,SD	-3.14*	-1.45	-4.56*	0.07*	10.42*	3.13	7
	I,Td	-1.05	-1.06	-4.21*	0.30*	8.96*	6.91	8
	I,SD,Tr	-1.02	-1.45	-4.53*	0.09*	10.32*	5.06	8
	None	2.94		-1.32	-1.33*	1.80	2.37	8

Income	I	3.65		-1.02	-1.23*	1.30	4.82	8
	I,SD	3.21		-2.40	-2.53	6.55	2.19	8
	I,Td	0.20		-1.04	-1.21*	1.30		
	I,SD,Tr	-0.04		-2.47	-2.47*	6.61	2.08	1
Lower block Network expansion	None	-0.57	-1.75	-1.92	0.59*	2.03	2.00	1
	I	-1.72	-1.73	-1.93*	0.60*	2.05	1.85*	1
	I, SD	-1.70	-1.65	-1.40	1.31*	1.81	2.88 2.72 2.20	1 1 1
	I,Td	-3.33	-1.63	-2.01*	0.63*	2.21	2.08	1
	I,SD,Tr	-3.33	-1.56	-1.46	1.40*	2.03	1.85*	1
Upper block network expansion	None	1.01	-1.53	-1.77	0.43*	1.66	2.06	1
	I	-1.19	-1.53	-1.75	0.45*	1.63	1.83*	1
	I,SD	-1.18	-1.77	-1.57	0.93*	1.65	0.72* 1.22*	4 4
	I,Td	-2.41	-1.46	-1.78	0.50*	1.71	1.41*	4
	I,SD,Tr	-2.44	-1.70	-1.63	1.06*	1.87	0.16*	4
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,Tr	-3.02	-2.17	-3.79*	-2.68	12.31*	0.10*	4