Journal of Applied Economics. Vol X, No. 1 (May 2007), 161-184

AN ESTIMATION OF RESIDENTIAL WATER DEMAND USING CO-INTEGRATION AND ERROR CORRECTION TECHNIQUES

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Submitted July 2004; accepted February 2006

In this paper short- and long-run price elasticities of residential water demand are estimated using co-integration and error-correction methods. Unit root tests reveal that water use series and series of other variables affecting use are non-stationary. However, a long-run co-integrating relationship is found in the water demand model, which makes it possible to obtain a partial correction term and to estimate an error correction model. Using monthly time-series observations from Seville, Spain, we find that the price-elasticity of demand is estimated as around -0.1 in the short run and -0.5 in the long run. These results are robust to the use of different specifications.

JEL classification codes: C22, D12, Q25

Key words: seasonal unit roots, residential water demand, price elasticity, co-integration, error correction

I. Introduction

While it is generally agreed that residential water users' short-run and long-run reactions to price changes might be substantially different, long-run water demand elasticities have been rarely estimated for European users. The main purpose of this paper is to calculate and compare short-run and long-run price elasticities of residential water demand using data from Seville, Spain. A secondary objective is to assess the usefulness of the techniques of co-integration (see Engle and Granger 1987, and Johansen 1988, among others) and error correction (Hendry et al. 1984) in the analysis of these price-elasticities of water demand.

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While this econometric methodology has proved very useful to estimate the demand for other types of transformed natural resources, such as gasoline and electricity, to the author's knowledge no previous published work has applied it to the study of water demand. The analysis in this paper is similar to these previous studies in that it deals with a resource whose price could affect purchase patterns of capital stock (water-using appliances) and whose consumption might respond partly to habit.

Residential water demand has been extensively analyzed during the last decades. Water demand appears as inelastic but not perfectly inelastic. Most applied studies focus on areas of the USA (e.g. Nieswiadomy and Molina 1989; Renwick and Green 2000; Taylor et al. 2004). Some exceptions which use European data are Hansen (1996), Höglund (1999), Nauges and Thomas (2000), and Martínez-Espiñeira (2002). Due to differences in how water is used and how it is priced, there are sharp geographic variations in price elasticities of demand, especially between Europe and North America (Arbués et al. 2003; Dalhuisen et al. 2003). Therefore, it is important that European policy makers be provided with the results from analyses based on European data that consider long-run versus short-run usage responses to price changes.

Short-run elasticities are usually found smaller than their long-run counterparts, suggesting that consumers might need time to adjust water-using capital stocks and to learn about the effects of use on their bills. Adjustment to consumption is normally assumed to be a fixed ratio of the total desired or equilibrium adjustment. Short-run elasticity is then modeled as a choice of utilization rate of the water-using capital stock, while the long-run involves the choice of both the size of this capital stock and the intensity of its use. Agthe et al. (1986), Moncur (1987), Lyman (1992), and Dandy et al. (1997) are examples of this approach, while more sophisticated econometric techniques have recently been applied (Nauges and Thomas 2001) and alternative models use information on water related technology to explicitly analyze endogenous technical change (Renwick and Archibald 1998). Agthe and Billings (1980) find, using a linear flow adjustment model, the short-run elasticity value (-2.226) much higher than the long-run value (-0.672), suggesting that there could be a short-run overreaction to price changes and that alternative techniques of time-series analysis might help solve these inconsistencies.

However, none of these studies used co-integration and/or error correction techniques to estimate short-run and long-run price effects. These methods have been extensively applied since Engle and Granger's seminal 1987 paper to, for example, the estimation of energy and gasoline demand (Bentzen 1994; Beenstock et al. 1999. Using co-integration analysis to estimate demand functions avoids problems of spurious relationships that bias the results and provides a convenient and rigorous way to discern between short-run and long-run effects of pricing policies. In this study, measures of short- and long-run price elasticity are estimated, as well as the speed of adjustment towards long-run values. The elasticities estimated suggest, in line with previous studies, that household water demand is inelastic with respect to its own price, but not perfectly so. The results show remarkable consistency between the different techniques used to analyze the dynamics of the relationships.

II. Data and methodology

A. Data

EMASESA, the private company in charge of supplying water and sewage collection service in Seville provided data for the period 1991-1999 on tariffs, number of domestic accounts, and total domestic use. The tariff consists of a fixed quota and an increasing three-block rate. It includes a water supply fee, a sewage

Table 1. Tariff evolution for water, sewage and treatment (1992 euro equivalents, excluding VAT)

Year	Fixed	PBL1*	PBL2*	PBL3*	Canon	TEC**	Fixed sew.	Sewage	Treat. sew.	Canon sew.
1991	1.063	0.139	0.214	0.386	0.000	0.000	0.000	0.075	0.130	0.000
1992	1.010	0.138	0.212	0.384	0.000	0.000	0.000	0.082	0.123	0.000
1993	1.133	0.132	0.206	0.378	0.000	0.020	0.000	0.089	0.120	0.000
1994	1.187	0.126	0.204	0.398	0.016	0.019	0.270	0.088	0.118	0.016
1995	1.246	0.125	0.213	0.421	0.016	0.021	0.285	0.092	0.124	0.016
1996	1.443	0.126	0.252	0.505	0.015	0.093	0.505	0.111	0.131	0.015
1997	1.524	0.130	0.260	0.609	0.015	0.093	0.550	0.125	0.140	0.015
1998	1.540	0.131	0.263	0.616	0.050	0.000	0.555	0.126	0.141	0.040
1999	1.533	0.131	0.262	0.614	0.048	0.000	0.553	0.126	0.141	0.039

Notes: * PBL*i* = water price in block *i*, for *i*=1,2,3. **TEC = Temporary extra charge. The last four columns refer to the sewage component of the bill.

collection fee, and a treatment fee, and, from 1994 on, a wastewater infrastructure fee (*canon*) collected on behalf of the government. Finally, from 1993 to 1997, a temporary extra fee was charged for the company to recover from the impact of a drought. Table 1 shows the evolution of the tariff between 1991 and 1999. Prices are expressed in constant pesetas (ESP) of 1992, translated into euro equivalents (1EU = 166.386 ESP).

Residential water use represents about 74% of the demand for drinking water in Seville and its metropolitan area. Sevillan households use 53% of the water in the toilet, in the kitchen, and for washing clothes. These components could be significantly affected by the efficiency of water-using equipment and the frequency of its renewal. An extra 39% is used in showers, which could be determined by both habits and the characteristics of water-use equipment. Outdoor use is minimal (EMASESA 2000, p. 7).

Seville suffered a serious draught between 1992 and 1995, during which important savings were achieved through measures such as media campaigns, municipal edicts and the ban of certain uses, water restrictions, and consumption control inspections. A detailed description of the measures implemented to reduce demand can be found in EMASESA (1997), García-Valiñas (2002), and Martínez-Espiñeira (2003), which is an extended version of the present article.

The original data were manipulated into the variables briefly described in Table 2 (where the subscript t refers to month t). The reader is referred to Martínez-Espiñeira (2003) for a detailed description of the variables, data sources and data manipulations. Descriptive statistics are shown in Table 3.

Table 2. Description of variables

Variable	Description	Units
ABONS	Number of domestic accounts.	accounts
Q	Average per capita domestic water use.	m3/capita month
P	Marginal price of water, based on the Taylor-Nordin specification (Taylor 1975; Nordin 1976). Following Billings (1982), an instrumental marginal price and difference are derived from an artificial linearization of the tariff structure by regressing the constructed bill amounts associated with each integer value of potential (between 1 m3 and 25 m3) monthly water use per account on these water use values. This instrumental marginal price is the slope of that estimated regression. The estimated intercept provides an estimate for the Taylor-Nordin difference.*	1992 euro equivalents/m^3
VI	Virtual income: difference between average salaries (*Wt*), a proxy for income, and *D*, the instrument for the Nordin-difference (Nordin 1976) variable. It is the intercept of the estimated linear function used to derive *P*. It only exerts an income effect caused by the nonlinearity of the tariff structure (Billings 1982), so, theoretically, its coefficient would have the same magnitude and opposite sign as the one of income. The values for *P* and *D* were calculated using the tariff schedules applied in each period.	1992 euro equivalents
INF	Binary variable with value 1 if water conservation information campaigns were being applied during the drought.	
RES	Daily hours of supply restrictions applied during the drought (EMASESA 1997). The number of hours of restriction a day is weighted by the number of days in the month to which that number applied.	hours/day
BAN	Binary variable with value 1 when temporary outdoor-use bans were applied during the drought.	
SUM	Binary variable with value 1 in May, June, July, and August.	
TEMP	Average of the daily maximum temperatures each month.	C/10
RAIN	Current level of precipitation.	mm/month

Note: * The modified Nordin-Taylor specification was chosen for this study in light of the results obtained with different specifications by Martínez-Espiñeira (2002) in other Spanish regions and because it was successfully applied by Martínez-Espiñeira and Nauges (2004) in a previous study in Seville, which also makes it possible to compare results based on the same price specification but different econometric models. Monetary values were deflated using the official provincial-wise retail price index.

Table 3. Summary statistics

Variable	Min	Max	Mean	Std. dev
ABONS	109,082	201,385	148,310	27,671
POP	683,028	719,588	702,529	9684
Q	5.054	8.201	6.352	0.648
P	0.472	0.7	0.571	0.08
W	1128	1410	1235	71,379
D	0.522	1.413	0.999	0.369
RES	0	12	1.4	2.99
BAN	0	1	0.273	0.445
SUM	0	1	0.333	0.474
TEMP	152	385	255.259	68.965
RAIN	0	3105	421.926	605.559
INF	0	1	0.319	0.466

Note: For all variables, N = 108.

B. Econometric methods

The techniques of co-integration (see Engle and Granger 1987) and error correction (see Hendry et al. 1984, among others) were used to investigate the dynamics of household water consumption and to measure the short-run and long-run effects of the price of water on household demand. Let us consider the simple form of a dynamic model:

$$y_t = \mu + \gamma_1 y_{t-1} + \beta_0 x_t + \beta_1 x_{t-1} + e_t, \quad t = 1, \ldots, T, \tag{1}$$

where y_t and x_t could represent, respectively, consumption and price at time t. The error term (e_t) is assumed independently and identically distributed. We will assume in the following that x_t is a one-dimensional vector for ease of exposition. μ, γ_1, β_0, β_1 are unknown parameters. In Model 1 the short-run and long-run effects of x on y are measured respectively by β_0 and $(\beta_0 + \beta_1)/(1 - \gamma_1)$.

Re-arranging terms yields the usual error-correction mechanism (ECM):

$$\Delta y_t = \mu + \beta_0 \Delta x_t - (1 - \gamma_1)(y_{t-1} - \theta x_{t-1}) + e_t, \quad t = 1, \ldots, T, \tag{2}$$

where Δ represents the difference operator (e.g. $\Delta y_t = y_t - y_{t-1}$) and $\theta = (\beta_0 + \beta_1)/(1 - \gamma_1)$. Thus, the estimation of the ECM gives a direct measure of the short-run and long-run effects of x on y through the estimates of β_0 and θ. The term $(y_{t-1} - \theta x_{t-1})$ in Model 2 can be seen as a partial correction for the extent to which y_{t-1} deviated from the equilibrium value corresponding to x_{t-1}.

That is, this representation assumes that any short-run shock to y that pushes it off the long-run equilibrium growth rate will gradually be corrected and an equilibrium rate will be restored. The expression $(y_{t-1} - \theta x_{t-1})$ is the residual of the long-run equilibrium relationship between x and y. Therefore, this error correction term will be included in the model if there exists a long-run equilibrium relationship between x and y or, in other words, if both series are co-integrated in the sense of Granger (see Engle and Granger 1987). If the series are co-integrated they will, in the long run, tend to grow at similar rates, because their data generating processes may be following the same stochastic trend, or may share an underlying common factor.

For any $0 < \gamma_1 < 1$, the absolute value of θ will exceed the absolute value of β_0 as long as β_0 and β_1 have the same sign. Under such likely conditions, the long-run adjustment to a change in the price of water will be stronger than the short-run adjustment. Although it would be a rare occurrence in practice, households might instead overreact in the short run to changes in prices. This effect has been observed in applications focusing on other commodities (Fouquet 1995).

The econometric analysis will proceed in two steps. In the first step, we test if x and y are co-integrated series, by testing for the stationarity of the series. If the series are integrated of the same order, a co-integrating vector might be then found such that a linear combination of the non-stationary variables obtained with that vector is itself stationary. If this proves to be the case, the estimation of the Granger co-integration relationship will give a measure of the long-run effect of x on y. In a second step, the co-integration residuals are used as an error correction term in the ECM above and the short-run effect and the speed of adjustment can be estimated.

Basic tests for the order of integration

A time series is I(i) (integrated of order i) if it becomes stationary after differencing it i-times. A non-stationary series can be represented by an autoregressive process of order *p*, so unit-root tests for a variable y_t usually rely on transformed equations of the form:

$$\Delta y_t = \mu + \lambda t + (\gamma - 1)y_{t-1} + \sum_{i=1}^{p} \gamma_j \ \Delta y_{t-i} + e_t . \tag{3}$$

This Augmented Dickey Fuller (ADF, Dickey and Fuller 1981) allows for an AR(*p*) process that may include a nonzero overall mean for the series and a trend variable (*t*). The special case where *p=1* corresponds to the Dickey-Fuller (DF) test. Its test statistics would be invalidated if the residuals of the reduced-form equation $\Delta y_t = \mu + \lambda t + (\gamma - 1)\ y_{t-1} + e_t$ were autocorrelated. In order to test the null hypothesis of nonstationarity, the t-statistic of the estimate of (γ-1) is compared with the corresponding critical values, calculated by Dickey and Fuller (1979 and 1981). A key consideration is how many lags of *y* to include in equation (3) and whether to include a constant and a trend variable. The choice can be based on the adjusted R^2, the Akaike Information Criterion (AIC), the Schwartz (1978) Bayesian Information Criterion, or the Schwert (1989) Criterion. These rules might lead to conflicting recommendations, so, for consistency, the sequential-t test proposed by Ng and Perron (1995) was used.

If the null of a unit root cannot be rejected, a second ADF test checks whether the series are integrated of order one or more than one, by testing for the null hypothesis of a unit root in the residual series of a regression in which the series has been differenced once. If the null of unit root is now rejected, the series is deemed I(1) or integrated of order one.

Further unit root tests

In small samples, the Dickey-Fuller tests above can lack power to reject the null hypothesis of non-stationarity (Baum 2001; Eguía and Echevarría 2004). Several complementary tests can be used that alleviate this problem. For example, the Dickey-Fuller Generalised Least Squares (DFGLS) approach proposed by Elliott et al. (1996) is likely to be more robust than the first-generation tests (Baum 2001). An alternative to the DF-type of tests above is the KPSS test (Kwiatkowski, Phillips, Schmidt and Shin 1992), which uses the perhaps more natural null hypothesis of stationarity rather than the DF style null hypothesis of I(1) or nonstationarity in levels. A KPSS may be applied together with a DF-style test, hoping that their verdicts will be consistent.

Finally, DF-style tests potentially lead to confusing structural breaks with evidence of nonstationarity, while recent unit root tests allow for structural instability in an otherwise deterministic model (Perron 1989, 1990; Banerjee et al. 1992; Perron and Vogelsang 1992; Zivot and Andrews 1992). It is therefore wise to conduct the latter when the former do not reject the null of nonstationarity. For example, Zivot and Andrews (1992) propose a method that allows for a single structural break in the intercept and/or the trend of the series, as determined by a grid search over possible breakpoints. Contrary to the case of Perron (1990)'s test, a priori knowledge of the location of the break is not needed. Subsequently, the procedure conducts a DF style unit root test conditional on the series inclusive of the estimated optimal breaks. Clemente et al. (1998)'s tests allow for two events within the observed history of a time series. Following the taxonomy of structural breaks in Perron and Vogelsang (1992), either additive outliers (the AO model, which captures a sudden change in a series) or innovational outliers (the IO model, which allows for a gradual shift in the mean of the series) can be considered.

Seasonal unit root tests

The tests described above for the stationarity of the series are not sufficient when the data exhibit a seasonal character, since seasonal unit roots must be investigated. A number of seasonal unit root tests have been proposed for the case of monthly data (Franses 1991; Beaulieu and Miron 1993) as an extension to the one suggested by Hylleberg et al. (1990). Seasonal unit root tests often exhibit poor power performance in small samples and that power deteriorates as the number of unit roots under examination increases. For example, in a simple test regression with no deterministic variables, the HEGY (Hylleberg et al. 1990) test procedure in the quarterly context requires the estimation of four parameters, whereas in a monthly context this number increases to twelve. In addition, the algebra underlying monthly seasonal unit root tests is more involved than in the quarterly case and the associated computational burden non-negligible. To circumvent these problems, the analysis of seasonal unit roots draws on the results found by Rodrigues and Franses (2003). These authors find out which unit roots affecting monthly data can also be detected by applying tests on quarterly data and, in particular, they show that 'with regard to the zero frequency unit root, there is a direct relationship between the monthly and quarterly root'. This means that the problem of non-stationarity of the series can be highly simplified by collapsing the monthly data into quarterly data (obtaining *N*/3 quarterly observations on all relevant variables by summing the monthly values or averaging them, depending on the nature of the variable) and then using the original HEGY test. If all the null hypotheses of any type of seasonal roots can be rejected based on the quarterly test, the monthly series can be also deemed free of seasonal unit roots.

As a test for a seasonal unit root in the $\{y_t, t=1, \ldots, T\}$ series, HEGY propose to apply OLS on the following model:

$$y_t - y_{t-4} = \pi_0 + \pi_1 z_{1,t-1} + \pi_2 z_{2,t-1} + \pi_3 z_{3,t-2} + \pi_4 z_{3,t-1} + e_t\ , \qquad (4)$$

where $z_{1t} = (1+L+L^2+L^3)y_t$, $z_{2t} = -(1-L+L^2-L^3)y_t$, $z_{3t} = -(1-L^2)y_t$, and L is the lag operator. To find that y_t has no unit root at all and is therefore stationary, we must establish that each of the $\pi_i (i = 1, \ldots 4)$ is different from zero. Moreover, we will reject the hypothesis of a seasonal unit root if π_2 and either π_3 or π_4 are different from zero, which therefore requires the rejection of both a test for π_2 and a joint test for π_3 and π_4. HEGY derive critical values for the tests corresponding to each of the following null hypothesis: H_{01}: $\pi_1 = 0$, H_{02}: $\pi_2 = 0$, H_{03}: $\pi_3 = 0$, H_{04}: $\pi_4 = 0$, H_{03+04}: $\pi_3 = 0$ and $\pi_4 = 0$. The tests statistics are based on Student-statistics (t-stat) for the first four tests and on a Fisher-statistic (F-stat) for the last one.

Co-integration

If these unit root tests suggest that the series are integrated of the same order, their long-run relationship is then investigated applying OLS on the simple model:

$$y_t = \theta x_t + u_t. \qquad (5)$$

The series *x* and *y* are said to be co-integrated if there exists a linear combination of those non-stationary series that is itself stationary. This means that their linear combination yields a stationary deviation (the residual series $\hat{u}_t$ is itself stationary). Following Engle and Granger (1987)'s approach, a unit root test[1] is applied whereby the resulting t-statistic is compared with the critical values provided by Engle and Yoo (1987).[2] The null hypothesis in this case is that of non-cointegration, so rejecting a unit root in the residuals in a DF type of test will constitute evidence of a co-integrating relationship among the variables. If the series are proved to be co-integrated, the estimate of θ in equation (5) provides a measure of the long-run effect of x on y.

Therefore, the long-run estimates of the price-elasticities are calculated using the estimated coefficients of the price variables in this equation. Additionally $\hat{u}_t$ can be used as an error correction term in the ECM:

$$\Delta y_t = \mu + \beta_0 \Delta x_t - (1 - \gamma_1)\hat{u}_{t-1} + e_t, \qquad t = 1, \ldots, T, \tag{6}$$

from which the estimations of β_0 and $(1-\gamma_1)$ would represent estimates of the short-run effect and the speed of adjustment towards the long-run values respectively. Short-run price elasticities are then derived from the estimates of price variables in this model.

III. Results

A. Unit root tests

First, the order of integration of all relevant series was determined, using a test of seasonal integration and the ADF test (see Section II.B). Table 4 summarizes the seasonal tests applied on the series collapsed into quarterly data. The non-rejection of H_{01}, together with the rejection of both H_{02} and the joint hypothesis H_{03+04}, suggests the presence of a unit root at the zero frequency and no seasonal unit roots. Since there is a correspondence between the quarterly and the monthly root at the zero frequency, not detecting seasonal unit roots at the quarterly level is enough to consider that the series is affected only by unit roots at the zero level

[1] The ADF test applied in this instance does not contain neither a trend nor a constant term, since the OLS residuals will be mean zero with a constant included in the co-integration regression.

[2] The critical values calculated by Dickey and Fuller are not appropriate, since the t-statistic's distribution is affected by the number of variables in the co-integration regression (Engle and Yoo 1987).

and no testing at the monthly level is necessary. The table shows that the null of seasonal unit roots can be rejected for all series,[3] with the not very surprising exception of *TEMP*.

Table 4. Quarterly seasonal unit root tests

Variable	Test	HEGY model specification[a]							
		SEAS		TREND		STREND		CONST	
		Statistic[b]	Lags[c]	Statistic[b]	Lags[c]	Statistic[b]	Lags[c]	Statistic[b]	Lags[c]
Q	H_{01}	-1.947	8	-2.379	0	-1.199	8	-2.929 *	0
	H_{02}	-3.998 **		-3.684 **		-3.867 **		-3.848 **	
	H_{03+04}	17.857 **		6.176 **		16.407 **		6.485 **	
P	H_{01}	-0.033	1,4,5	-1.136	0	-1.528	0	-1.108	0
	H_{02}	-4.116 **		-2.708 **		-2.795 *		-2.708 **	
	H_{03+04}	12.143 **		11.666		13.407 **		11.701 **	
P^2	H_{01}	0.044	1,4,5	-1.237	0	-4.016 **	4	-1.124	0
	H_{02}	-4.455 **		-2.750 **		-3.475 **		-2.743 **	
	H_{03+04}	14.219 **		12.233 **		20.187 **		12.194 **	
VI	H_{01}	1.181	0	-0.787	1	0.311	0	-0.791	1,5
	H_{02}	-3.172 **		-0.418		-3.051		-0.829	
	H_{03+04}	6.711 **		0.332		6.604		0.744	
RES	H_{01}	-2.970 *	6	-2.643	3,5,7	-3.549 *	0	-2.732 *	5,6
	H_{02}	-3.177 **		-4.419 **		-4.299 **		-3.291 **	
	H_{03+04}	15.982 **		2.785 *		7.027 **		13.050 **	
TEMP	H_{01}	-2.419	0	-3.401 *	2,3,8	-2.399	0	-3.8	2,3,6,7,8
	H_{02}	-2.133		-1.281		-2.111		-1.057	
	H_{03+04}	8.117 **		0.040		7.850 **		0.28	
RAIN	H_{01}	-2.062	4	-1.888	0	-2.775	2,4,8	-1.779	0
	H_{02}	-1.674 *		-1.674 *		-3.193 **		-1.677 *	
	H_{03+04}	2.437 *		2.437 *		7.753 **		2.400 *	

Notes: (a) Test specifications: SEAS (Seasonal dummies + constant), TREND (Constant + trend), STREND (Seasonal dummies + constant + trend), CONST (constant only). (b) HEGY estimates, ** and * denote a statistic significant at the 5% and 10%. The first two rows show t-stats and the third one shows the F-stat. (c) Lag length and lags of the fourth difference of the time-series to be included in the auxiliary regression.

[3] The test also permitted to reject the null hypothesis of seasonal unit roots in the *INF* series, although the estimates are not shown, since this variable is not used in most of the main final water demand models.

After detecting with the seasonal approach the presence of only unit roots at the zero frequency, the order of integration of the series was further tested using Dickey-Fuller-type tests. Two auxiliary DF regressions, with and without a trend, were used and the optimal lag was chosen by an automatic sequential t-test. The results, shown in Table 5, reveal that the trend component is not relevant in most cases and that most variables proved to be I(1). The hypothesis tests permit the rejection of the null of non-stationarity of the differenced series at the 99% level of confidence. Once again, there are some doubts about the climate variables. *TEMP* appears to be stationary, but the seasonal unit root tests did not reject the hypothesis of seasonal roots, so this variable should be considered with caution, since it might be I(0, 1). In the case of *RAIN*, we also see that the series might actually be stationary in levels also at all frequencies, I(0). Since a definite claim cannot be made that the climatic variables are nonstationary, their introduction in the co-integration relationship and error-correction model cannot be understood

Table 5. Augmented Dickey-Fuller unit root tests

Variable	No trend			With trend			
	t-stat[a]	Lags[b]	B-stat[c]	t-stat[a]	Lags[b]	B-stat[c]	Trend t-ratio
Q	- 2.527	4	0.64	- 2.922	4	0.64	- 2.00 ***, 2.01 *** [d]
ΔQ	- 6.276***	5	0.73	- 6.966 ***	8	1	2.48 **
P	- 1.165	0	0.26	- 1.888 **	0	0.26	1.5
ΔP	- 3.801 ***	7	0.22	- 10.550 ***	0	0.25	-0.26
P^2	-1.202	0	0.23	- 1.847	0	0.27	1.42
ΔP^2	- 10.569 ***	0	0.23	- 10.525 ***	0	0.23	-0.28
VI	0.333	8	1.47 **	- 3.970 **	8	1.46 **	- 4.15 ***, 4.18 *** [d]
ΔVI	-12.709 ***	8	1.29 *	- 13.800 ***	8	1.07	3.35 ***
RES	- 2.547	1	0.99	-3.369 *	2	0.47	- 1.23
ΔRES	- 5.472 ***	1	0.48	- 5.454 ***	1	0.48	- 0.28
$TEMP$	- 8.921 ***	8	1.45 **	- 9.016 ***	8	1.44 **	1.17
$\Delta TEMP$	- 12.192 ***	8	1.43 **	- 12.041 ***	8	1.43 **	- 0.22
$RAIN$	- 6.849 ***	0	0.89	- 5.100 ***	4	0.37	0.43
$\Delta RAIN$	- 5.946 ***	0	1.13	- 12.273 ***	0	1.14	0.09

Notes: (a) t-ratio of estimates ***, **and *denote a t-ratio significant at the 1%, 5% and 10%. (b) The number of lags (with a maximum of 8) to be included was selected using the Ng-Perron sequential-t test. (c) Bartlett's (B) statistic of a cumulative periodogram white-noise test (H0: error is white noise). (d) This series exhibited a quadratic trend rather than a linear trend.

in the same way as the remaining variables. For this reason, an alternative model that accounts for seasonality using only the binary variable *SUM* is also reported.[4]

The augmentation of the basic DF regression with extra lags (in Section II.B) was motivated by the need to generate independent and identically distributed (iid) errors. The number of lags was selected by a Ng-Perron (1995) sequential-t test. A maximum of eight lags was considered, given the small sample size. A cumulative periodogram white-noise test was used to check that the error was white noise in the unit root test regressions.[5] An alternative solution is the Phillips and Perron (1988) test (PP), which uses the same models as DF but, instead of lagged variables, employs a non-parametric correction (Newey and West 1987) for serial correlation. The critical values for both the Dickey Fuller and Phillips Perron tests have the same distributions (critical levels are reproduced in Hamilton 1994). In principle, the PP tests should be more powerful than the ADF ones, so the unit root tests were conducted using both. The results of the PP test are not reported, but available upon request.

B. Further unit root tests

As described in Section II.B, the conclusions of the Dickey-Fuller type tests can be suspect in small samples and when the possibility of structural breaks has to be considered. A DFGLS (Elliott et al. 1996) test was used to confirm the nonstationarity of the main series in the model. Since it is more powerful than the basic Dickey Fuller tests, this test will more easily reject the null of nonstationarity. A KPSS test (see Baum 2000) was also conducted to test the alternative hypothesis of stationarity of the series. The results of both types of tests, in Table 6, show no relevant differences with respect to those described in Section III.A.

Similarly, structural breaks in the model's series could affect the conclusions of

[4]A model (not reported but available upon request) based on a co-integration equation that did not include neither *SUM* nor the climatic variables (*TEMP* and *RAIN*) resulted in very similar price elasticities (see Section III.5) but slightly worse fit than the models that include seasonal variables. This, of course, confirms the intuition that the seasonal variables do not substantially affect the long-run dynamics of water demand.

[5] In the case of *TEMP* and *VI* there were some doubts about whether the error had been whitened. However, we do not suspect that this results in a bias of the unit root tests coefficients substantial enough to affect their conclusions.

Table 6. DFGLS and KPSS unit root tests

Variable	DFGLS (no trend)		DFGLS (trend)		KPSS (no trend)		KPSS (trend)	
	t-stat	Lags	t-stat	Lags	t-stat	Lags	t-stat	Lags
Q	-2.511 **	6	-2.565	6	0.620 **	3	0.427 ***	3
Δ*Q*	-2.082 **	4	-5.287 ***	3	0.045	3	0.038	3
P	-0.559	1	-1.737	1	2.320 ***	3	0.274 ***	3
Δ*P*	-7.234 ***	1	-7.297 ***	1	0.106	3	0.099	3
P^2	-0.619	1	-1.715	1	2.250 ***	3	0.270 ***	3
ΔP^2	-7.210 ***	1	-7.260 ***	1	0.107	3	0.099	3
VI	1.151	8	-1.612	8	0.630 **	3	0.209 **	3
Δ*VI*	-12.183 ***	8	-12.269	8	0.035	3	0.028	3
RES	-2.932 ***	6	-3.134 ***	6	0.387 *	3	0.129 *	3
Δ*RES*	-4.303 ***	7	-4.317 ***	7	0.049	3	0.038	3
TEMP	-0.938	8	-2.594	7	0.019	3	0.014	3
Δ*TEMP*	-7.190 ***	8	-8.960 ***	8	0.339	3	0.016	3
RAIN	-6.253 ***	1	-6.548 ***	1	0.111	3	0.059	3
Δ*RAIN*	-0.963	6	-2.554	5	0.020	3	0.018	3

Note: *** ,**and *denote a t-ratio significant at the 1%, 5% and 10%.

the unit-root tests, above all because of the effects of the drought and the conservation measures adopted by the water utility. For this reason Clemente et al. (1998)'s tests, which allow for either one or two structural breaks in the time series (see Baum 2005), were used to obtain further confirmation of the nonstationarity of the main series considered. The results (not reported but available upon request and in Martínez-Espiñeira 2003) confirm that most series exhibit evidence of structural breaks. However, even when these are considered, there is not enough evidence to reject the null of nonstationarity. Since the additional unit root testing supports the notion that, with the likely exception of the climatic variables, the series used in the demand model are nonstationary, the next step is to analyze the co-integrating relationships among those series.

C. Co-integration regression analysis

All the series in first-differences are stationary, so the next step is to check that there is a long-run equilibrium relationship between the variables. This requires an extension of the linear relationship between water consumption and a series of variables that the economic theory suggests appropriate. The model given by equation (5) was extended into two alternative models (time subscripts have been dropped to simplify the exposition):

$$Q = \alpha + P + P^2 + RES + VI + BAN + SUM + u \tag{7}$$

including the binary variable *SUM* instead of the climatic variables (see Section II) and:

$$Q = \alpha' + P + P^2 + RES + VI + BAN + TEMP + RAIN + u'. \tag{8}$$

Table 7 shows the estimated OLS coefficients and t-statistics in these estimations. The ADF test shows that the hypothesis that the residuals in Model 7 are non-stationary can be rejected. The relevant t-ratio is -5.625 in the usual test of a unit root and must be compared with the critical values provided by Engle and Yoo (1987), which depend on the dimension of the time-series and on the number of variables included in the model.[6] The DW statistic is also higher than the R^2, which suggests the existence of a co-integration relationship. The long-run price-elasticity calculated at the means of price and quantity according to Model 7 is -0.491. All the variables present the expected signs and are highly significant.

The ADF test shows that the hypothesis that the residuals in Model 8 are nonstationary can be rejected too. The relevant t-ratio is -5.364.[7] The DW statistic is again higher than the R^2. The corresponding long run price-elasticity according to Model 8 is -0.494, basically the same as the one obtained with Model 7. Once again, all variables have the expected signs and are highly significant. The exception

[6] This value permits the rejection of the null of no co-integration at a 99% confidence level, but it is achieved when the auxiliary regression includes no lags. Five lags are selected by Ng Perron's sequential t-ratio and the Akaike Information Criterion test, yielding a t-statistic of -3.053 and one lag is selected by the Hannan-Quinn (1979) criterion, yielding a t-statistic of -3.804.

[7] This value permits the rejection of the null of no co-integration at a 99% confidence level, but it is achieved when the auxiliary regression includes no lags. If the auxiliary regression is run with the optimal number of lags (three) the t-ratio is -4.192.

Table 7. Cointegration models and error correction equations

Variable	Model 7	ECM7	Model 8	ECM8
P	-78.629 ***		-76.950 ***	
P^2	64.070 ***		62.569 ***	
RES	-0.066 **		-0.067 **	
VI	0.003* **		0.002 ***	
BAN	-0.451 **		-0.473 **	
SUM	0.317 ***			
$RAIN$			0.000	
$TEMP$			0.002 ***	
ΔP_t		-37.437		-32.609
$\hat{u}_{t-1}$		-0.218 *		-0.249 **
ΔP^2_t		31.264		27.572
ΔRES_t		-0.076 **		-0.098 ***
ΔSUM_t		0.317 ***		
ΔP_{t-1}		-71.006* **		-79.555 ***
ΔP^2_{t-1}		54.729 **		61.881 ***
ΔQ_{t-1}		-0.273 **		-0.276 **
$\Delta TEMP_t$				0.004 ***
$\Delta RAIN_t$				0.000
constant	26.489 ***	0.019	26.197 ***	0.019
L l	-47.07	-7.233	-50.1	-3.197
N	108	106	108	106
Adj-R^2	0.644	0.386	0.619	0.425
Durbin-Watson	0.919		0.856	
Jarque-Bera χ (2)	2.044		2.059	
ARCH-LM order(1) $\chi^2(1)$	2.913 *		0.001	
Breusch-Godfrey LM χ^2 (1)	16.917 ***		1.876	
White's general test $\chi^2(44)$	44.025		75.721 **	
Cook-Weisberg χ^2 (1)	3.26*		0.78	

Note: ***, ** and * denote a t-ratio significant at the 1%, 5% and 10%.

is *RAIN*, which presents a positive sign, while we would normally expect more precipitation to reduce water use. However, it cannot be rejected that its coefficient is null. According to the ADF tests, the null of no co-integration can only be rejected if the lag length of the auxiliary regression is not optimally chosen. However, the value of the DW test and economic intuition suggest that a long-run relationship would govern the variables concerned.

There are more than two nonstationary series in the model, so there could exist more than one co-integrating relationship. For this reason and in order to obtain more evidence on the existence of a co-integrating regression, the Johansen-Juselius maximum likelihood method for co-integration (see Johansen 1988; Johansen and Juselius 1990; and Osterwald-Lenum 1992, for details) was used to determine the number of co-integrating relationships. The results, not reported but available upon request, show the null hypothesis of more than one co-integrating relationship is rejected at the 1% level of significance in all cases, except in the case of the trace test for Model 7, which rejects the null of non-cointegration only at about the 15%.[8] Likelihood-ratio and Wald test statistics for the exclusion of variables from that co-integrating relationship were also conducted, and all variables included in the co-integration tests were found relevant. Therefore, the Johansen tests support the assumption of co-integration for both models.

D. Error correction models

Since most of the evidence points towards the stationarity of the residuals of the co-integrating regressions, their residuals can be introduced as error correction terms in two error correction models. The x_t variables in equation (6) are substituted by first differences and lagged differences of the co-integrating variables. The first error correction specification, ECM7 includes a summer variable, whereas the second model, ECM8, includes *TEMP* and *RAIN* (*TEMP* might suffer problems of seasonal unit roots and it is dubious that *TEMP* and *RAIN* are nonstationary, so this second model should be considered with caution). Table 7 reports the results of these OLS estimations. These include lagged values of the differences of some variables.[9] *VI* was left out of the ECM models, since it showed problems of

[8] Martínez-Espiñeira (2003) reports the results of these tests.

[9] Interestingly the lagged value of *RES* suggests that at the beginning of an increased restriction period the savings are larger than after allowing for some adjustment by the consumers (perhaps filling up water containers during the non-restricted hours). The lagged value of *SUM* also suggests that being at the beginning of the summer period also increases use (perhaps some seasonal uses, such as filling up swimming pools, are more intense at the beginning of the season).

multicollinearity with the price variables and its introduction made them non-significant. It is reasonable to assume that changes in income tend to affect water use only in the long run, most likely through impacts on the composition of the capital stock. *BAN* was found non-significant too and it was removed from the ECM regressions. The speed of adjustment towards equilibrium given by predicted $\hat{u}_{t-1}$ in Table 7, corresponding to (y_{t-1}-θx_{t-1}) in the notation used in equation (2), is -0.218 in ECM7 and -0.249 in ECM8. It can be seen that these error correction terms have the expected negative sign and are both significant, which further supports the acceptance of the co-integration hypothesis.

The Ramsey RESET-test (using powers of the fitted values of Q_t) shows that the null hypothesis that Models ECM7 and ECM8 have no omitted variables cannot be rejected. Table 7 includes a battery of diagnostic tests used to check that the residuals are normally distributed and are neither autocorrelated nor heteroskedastic in the error correction equations. These include a Jarque-Bera (1980) test for normality of the residuals; White's (1980) general test statistic and the Cook-Weisberg (1983) - Breusch-Pagan (1979) test, which uses fitted values of Q_t, for heteroskedasticity; a Lagrange multiplier test for autoregressive conditional heteroskedasticity (ARCH), based on Engle (1982); and a Breusch (1978)-Godfrey (1978) LM test. They all present acceptable values, with the exception of the Breusch-Godfrey LM test, which leads to the rejection of the null of non-autocorrelation in ECM7. An alternative model with extra lagged values of the price variables solves this problem and yields a short-run elasticity of -0.073, as reported below. The results of this additional augmented regression do not differ significantly from the ones reported and are available upon request.

E. Price elasticities

The computation of short-run price elasticities (ξ_{SR}) using the average price and water consumption, yields the following results. Using ECM7 and the co-integration regression in Model 7, ξ_{SR} = -0.159 (while the augmented model used to correct for autocorrelation would yield ξ_{SR} = -0.073) and the ξ_{LR} = -0.494. Similarly, ECM8 and Model 8 yield ξ_{SR} = -0.101 and ξ_{LR} = -0.491. These estimates of price-elasticities confirm that residential water demand is inelastic to its price, but not perfectly so. Almost all the papers published on residential water demand agree on this result. Additionally these results confirm the intuition that long-run elasticities are higher (in absolute value) than short-run ones (Dandy et al. 1997; Nauges and

Thomas 2003; Martínez-Espiñeira and Nauges 2004), being also higher than most of the measures that have been obtained in other European countries (Arbués et al. 2003). The use of the co-integration approach to model the demand for water yields rather sensible results and helps to distinguish between the short-run effects and the long-run effects of pricing policies.

F. Wickens-Breusch one-step approach

The Engle-Granger procedure described above enjoys important attractive asymptotic properties but it also suffers weaknesses. In finite samples, the parameter estimates are biased. The extent of this bias can be severe, and will depend on omitted dynamics and failure of the assumption of weak exogeneity among other things. The reasonable sample size and the fact that the estimates agree with economic theory and previous research based on alternative econometric techniques suggest in principle that this might be a minor problem in this case. Another problem, however, is that one cannot test the long run parameters. For these reasons, an additional regression was run using the one-step Wickens-Breusch (1988) approach. The full results of this analysis are available upon request and in Martínez-Espiñeira (2003). The associated price-elasticities, at the means of price and quantity are ξ_{SR} = -0.08 and ξ_{LR} = -0.405 in the model that uses *SUM* and ξ_{SR} = -0.113 and ξ_{LR} = -0.514 in the model that uses *TEMP* and *RAIN*. These elasticities are very close to the ones calculated with the Engle-Granger approach, which suggests that they can be accepted with more confidence. They also fall reasonably close to the values previously obtained using alternative econometric techniques on data from the same city. García-Valiñas (2002) estimated a price-elasticity of -0.25 for the first block of consumption and -0.77 for the second block, while García-Valiñas (2005) estimated the price-elasticity as -0.49. Martínez-Espiñeira and Nauges (2004) found also that the short-run elasticity fell around the value of -0.10, using a Stone-Geary demand specification.

IV. Conclusions and suggestions for further research

This study is innovative in two aspects. This is the first time that co-integration and error correction techniques are used to study water consumption. Moreover, the estimation of residential water demand using time-series monthly data is still rather uncommon in Europe. The application of these techniques to monthly data to the case of Seville leads to satisfactory results. The fit of the Granger co-integration relationship between water use and the variables that should be expected to influence it in the long run and of the error correction models is quite good. The dynamic properties of the series were analyzed using different approaches and two alternative specifications for the water demand functions were used. However, the results in terms of price elasticities, most of all in the short run, are remarkably close. This robustness to specification and testing procedures leads to confidently accept the main results.

The estimates of the price effects obtained are less than one in absolute value, which confirms the inelasticity of household demand with respect to the price of water. As predicted by the theory, the long-run price elasticities are greater, in absolute value, than their short-run counterparts.

The measure of the impact of pricing policies on the behavior of households depending on the changes that these policies introduce in the tariff structure is still an open research area. The long-run effects of water pricing on water use should be investigated using other datasets, involving different regions, and, if possible, longer time-series or panel data. Ideally, studies should be conducted at the individual level, with observations linked to the ownership and frequency of renewal of capital stock.

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