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Managing a scarce resource in a growing Asian economy: Water usage in Hong Kong

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ABSTRACT

An econometric analysis of Hong Kong's monthly per capita water usage for the 25-year period of April 1985 through March 2010 reveals that per capita usage is insensitive to price but dependent upon past usage, per capita income, weather, and seasonal factors, with rising income countering what would otherwise be a downward trend. Given Hong Kong's current inflationary environment and large government budget surplus, these findings affirm the Hong Kong Water Supplies Department's adopted strategy of total water management towards sustainable use of water resources, in lieu of either periodic service interruption or price increases as policy instruments.

1. Introduction

Along with Singapore, South Korea, and Taiwan, Hong Kong has achieved the well-deserved label of an Asian dragon for its commitment to industrialization starting in the 1960s, initially through successful cottage-like textile and clothing industries and plastics manufacturing, and subsequently through the production of low-cost electronics equipment and watches. In 1978, Chinese Premier Deng Xiaoping opened the door to the People's Republic of China, and Hong Kong

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manufacturers took advantage of their ready access to a large and low-wage labor force in neighboring Guangdong Province to engage in "outward processing," whereby goods would be wholly or partially manufactured on the mainland and then shipped across the border for completion and/or ultimate export. The territory's ties to a rapidly growing economic giant, China, which is now the world's second-largest economy, were solidified on July 1, 1997, when the territory became a Special Administrative Region (SAR) of China under the Sino-British Joint Declaration signed in late 1984. Notwithstanding the economic pains caused by the Asian Financial Crisis of 1997–98 and the world-wide economic woes a decade later, the once-again vibrant Hong Kong economy boasts a per capita income of approximately US$31,709 at current prices in 2010 for the seven million people that live within its 1000 square kilometers.[4]

From April 1985 to March 2010, the period on which we focus, real GDP in Hong Kong grew at an annual rate of 3.89 percent, while the population grew at a much lower rate of 1.04 percent. Over the same period, the annual growth rate in total water usage was 1.60 percent, suggesting that Hong Kong's per capita water usage has at least partially been driven by the economic growth that the SAR has enjoyed.[5]

Because of China's rapid growth, we conjecture that Hong Kong's per capita water usage will likely rise. Moreover, there have been stark internal differences in China's rapid growth, with Guangdong Province being at the upper end of the scale (Marti, Puertas, & Fernández, 2011). Indeed, our conjecture is borne out by an econometric analysis that enables us to isolate the impact of real per capita income from the potential impact of price – the water rate – on per capita water usage, and shed important light on the SAR's demand-side water-management options.

Eighty percent of the territory's fresh water now comes from Guangdong's East River; the remainder comes from rainfall stored in local reservoirs (Wong, Zhang, & Chen, 2010). Thus, unlike a one-year period almost 35 years ago when water service was interrupted by as much as 14 h per day for a duration of 288 days (Woo, 1994), mandatory water rationing that results in welfare losses (Grafton & Ward, 2008; Woo & Lo, 1993; Woo, 1994) was unnecessary during the more recent 25-year period, despite the SAR's increasing water usage.

To ensure supply reliability and sustainable use, the Hong Kong Water Supplies Department (HKWSD) in 2008 adopted a total water management (TWM) strategy to implement the integrated management of water demand and supply (HKWSD, 2008). On the demand side, the TWM entails enhanced public education on conservation, the promotion of water-saving devices, and the expanded use of seawater for toilet flushing. On the supply side, the TWM strengthens protection of water resources, including water-leakage control and pressure management, reclamation, and seawater desalination. The TWM, however, does not include the price-induced conservation policy proposed by Woo and Lo (1993) and Woo (1994) for Hong Kong, and more recently by Gabriel, de Azevedo, and Baltar, 2005, Howe (2005), and Ferrara (2008) for other parts of the world. Indeed, Hong Kong's water rates have not changed since February 1995; and consumers do not see monthly water price signals because of the Department's once-every-four-months billing policy.[6]

Using a ten-year sample of monthly data, from April 1973 through March 1984, Woo (1994) estimated that Hong Kong's water-demand price elasticity was −0.47, which was subsequently corroborated by similar findings reported by Espey, Espey, and Shaw (1997), Arbues, Garcia-Valinas, and Martinez-Espineira (2003), and Ferrara (2008). When compared to service interruption, price was seen to be a relatively effective and less costly policy instrument to deter usage (Woo & Lo, 1993; Woo, 1994). This, however, begs the question as to whether, given the new levels of prosperity achieved in Hong Kong and the prospects for even higher levels, current users are even as price sensitive as their price-inelastic counterparts of 30 years ago, or whether alternative demand-management strategies that do not include service interruption and pricing *must* be introduced into the discussion.

To help answer this question, we conduct an econometric analysis of water usage in Hong Kong for the 25-year period of April 1985 through March 2010. Using a sample of 300 monthly observations, the analysis quantifies, in particular, the per capita water-usage effects of real per capita income and the real water rate, along with several other factors, such as weather and seasonality. Most saliently, we show that Hong Kong water users are insensitive to the real rates that they pay for their water, at least insofar as concerns the rate structure maintained for the past quarter of a century. Given Hong Kong's current inflation of 2.6 percent per year[7] and large government budget surplus of HK$71.3 billion (≈US$9.1 billion at the pegged exchange rate of HK$7.8/US$1) for the 2010/2011 fiscal period,[8] these findings do not support a policy of raising rates to induce conservation. Instead, they affirm the HKWSD's adopted TWM strategy towards sustainable use of water resources.

The threefold contribution of this paper is that it (a) offers updated evidence that pricing may not be effective in managing Hong Kong's water usage, (b) confirms the HKWD's expressed urgency when adopting the TWM, since water usage will likely rise over time because of real income growth, and (c) highlights the necessity for updated information on customer price responsiveness when assessing how effective price is as a tool for water demand management.

[4] The US$31,709 figure comes from the U.S. State Department (http://www.state.gov/r/pa/ei/bgn/2747.htm). When adjusted for purchasing power parity, Hong Kong's per capita GDP for 2010 is US$46,900 (https://www.cia.gov/library/publications/the-world-factbook/geos/hk.html).

[5] As noted by an insightful referee, 1.04 percent of the 1.6 percent increase in Hong Kong's total water usage is driven by the population increase. Since 2/3 of the total usage is non-residential, the per capita residential usage may have been decreasing over the 25-year period, reflecting the conservation effort by Hong Kong residents.

[6] Further details are available from the HKWSD's web site: http://www.wsd.gov.hk/en/customer_services_and_water_bills/water_and_sewage_tariff/water_and_sewage_tariff/index.html.

[7] http://www.rttnews.com/ArticleView.aspx?Id=1486746.

[8] http://www.cnbc.com/id/41728197/Hong_Kong_Posts_Budget_Surplus_of_9_1_Billion.

The remainder of the paper proceeds as follows. Section 2 presents our econometric model, which is implemented using the data described in Section 3. Section 4 discusses our regression results. Section 5 concludes.

2. Model

A commonly-used econometric specification for analyzing per capita water usage is a log-linear regression (Arbues et al., 2003; Espey et al., 1997). Woo (1992, 1994) used the log-linear specification to analyze Hong Kong's per capita usage in April 1973–March 1984, after determining that it was not statistically different from the more general Box–Cox specification.

For the analysis herein, we extend the log-linear specification used by Woo (1992, 1994) in two ways. First, we use a time-trend variable to capture the residual effect of factors (e.g., public awareness and conservation) unaccounted for by the regression's other explanatory variables (e.g., price, income and weather). Second, we include the lagged per capita consumption as an additional explanatory variable to capture the possible dependence of current consumption on past consumption.[9] The resulting specification is consistent with the theory of consumer utility maximization, and the coefficient estimate for a continuous explanatory variable, such as income and price, measures the short-run responsiveness of usage with respect to that variable; that is, the income and price elasticity of usage (Woo, Zarnikau, & Kollman, 2012).

The partial-adjustment aspect of the specification is rooted in Nerlove's seminal article (1956, particularly pp. 308–310). In the present context, it is assumed that "in the absence of changes in the [explanatory] variables upon which demand depends, the current [water usage] would change in proportion to the difference between the long-run equilibrium quantity and the current quantity" (p. 308).

Specifically, let Q_t denote the actual per capita water usage, in cubic meters (m^3) per month in month t ($t = 1$ for April 1985, . . ., 300 for March 2010), and let Q_t^* denote its equilibrium level. We assume the long-run equilibrium of the natural log (ln) of water usage is:

$$\ln Q_t^* = \beta^* + \varphi^* t + \Sigma_k \delta_k^* D_{kt} + \beta_Y^* \ln Y_t + \beta_P^* \ln P_{t-3} + \Sigma_j \omega_k^* W_{jt} \tag{1}$$

in term of other explanatory variables to be discussed later.

We further assume the following partial-adjustment process (Kmenta, 1986, p. 529):

$$\ln Q_t - \ln Q_{t-1} = (1-\lambda)(\ln Q_t^* - \ln Q_{t-1}) + \varepsilon_t. \tag{2}$$

Nerlove (1956, p. 309) refers to $0 < \lambda \le 1$ in Eq. (2) as "the elasticity or coefficient of adjustment according to whether quantity is expressed in logarithms or not." The closer is λ to unity and hence $(1-\lambda)$ to zero, the slower the system approaches equilibrium.

We assume ε_t in Eq. (2) is a stationary AR(1) random error such that $\varepsilon_t = \rho\varepsilon_{t-1} + u_t$ with $0 < |\rho| < 1$ and u_t denoting white noise. Thus, overt time a past random shock has a dampening effect on $(\ln Q_t - \ln Q_{t-1})$, the observed monthly percentage change in per capita water usage.

After some simple algebraic manipulations involving Eqs. (1) and (2), the short-run water usage $\ln Q_t$ becomes the dependent variable in the regression model below:

$$\ln Q_t = \beta + \varphi t + \sum\nolimits_k \delta_k D_{kt} + \beta_Y \ln Y_t + \beta_P \ln P_{t-3} + \sum\nolimits_j \omega_j W_{jt} + \lambda \ln Q_{t-1} + \varepsilon_t. \tag{3}$$

The first set of right-hand-side independent variables in Eq. (3) are as follows:

- A monthly trend variable, t, that aims to capture any trend effect not accounted for elsewhere in the regression, such as the introduction of more efficient water-using appliances (e.g. front-loading clothes washers, low-flow shower heads, and low-flush toilets), new construction, building renovation, building turnover, and the HKWSD's improvement of the water distribution system. Inasmuch as some of these factors would tend to reduce water usage whereas others would tend to increase it, we have no prior expectations as to the sign of φ.
- A binary indicator D_{kt} that equals unity if observation t occurs in month k and is zero otherwise (e.g., $D_{1t} = 1$ if t corresponds to January; and 0, otherwise). The binary indicators aim to capture any seasonality in water usage. As with the trend, we have no prior expectations as to the sign of δ_k.
- The natural logarithm of monthly real per capita income (HK\$/month) = (Quarterly GDP/three months)/(monthly population), denoted $\ln Y_t$, where the monthly consumer price index (CPI = 1.0 in April 2005) is used as the deflator. This variable aims to capture the effect of changes in per capita income on water usage. The log-linear format with partial adjustment through $\ln Q_{t-1}$ enables the direct interpretation of β_Y as the short-run income elasticity of water usage, which we expect to be positive. Based on Eqs. (1) and (2), the long-run income elasticity is $\beta_Y^* = \beta_Y/(1-\lambda)$.
- The natural logarithm of the monthly real average water rate (HK\$/m^3) lagged three months. Denoted $\ln P_{t-3}$, it is the annual nominal rate (=HKWSD's fiscal-year revenue/fiscal-year usage) deflated by the CPI. The three-month lag reflects the fact that the HKWSD bills a customer once every four months. The log-linear, partial-adjustment format enables the direct

[9] This dependence was found to be statistically insignificant for the April 1973–March 1984 period (Woo, 1992, p. 2594).

Table 1
Water tariffs used by the Hong Kong Department of Water Supplies for four-monthly billing since February 1995.

Tariff type	Rates (nominal HK$/m^3)	Remarks
Domestic	First 12 m^3: free	The domestic tariff has increasing block rates.
	Next 31 m^3: 4.16	The four-month period is assumed to have 121 days.
	Next 19 m^3: 6.45	When computing a customer's actual bill, the quantities
	Above 62 m^3: 9.05	by block are pro-rated to match the number of days in a customer's billing cycle.
Non-domestic	For trade: 4.58	The non-domestic tariff has fixed rates differentiated by end-use.
	For construction: 7.11	
	For non-ocean-going shipping: 4.58	
	For ocean-going shipping: 10	

Source: Hong Kong Department of Water Supplies, http://www.wsd.gov.hk/en/customer_services_and_water_bills/water_and_sewage_tariff/water_-and_sewage_tariff/index.html.

interpretation of β_P as the short-run price elasticity of water usage, which we expect to be negative. Based on Eqs. (1) and (2), the long-run price elasticity is $\beta_P^* = \beta_P/(1-\lambda)$.

The second set of right-hand-side independent variables aims to characterize the dependence of water usage on the weather:

- W_{1t} = monthly average temperature (°C). As is generally the case, we would expect water usage in Hong Kong to move with the weather, and be likely to increase as the temperature rises ($\omega_1 > 0$).
- W_{2t} = monthly sunshine (total hours of bright sunshine per month). Water usage is likely to increase when sunshine prevails ($\omega_2 > 0$), as hot sunny days tend to raise the demand for water.
- W_{3t} = monthly rain (cm per month). We would expect the cooling effect of rainfall to reduce the demand for water and thus $\omega_3 < 0$.
- W_{4t} = monthly average wind-speed (km per hour), with windy days tending to reduce the demand for water, implying $\omega_4 < 0$.

The last explanatory variable, $\ln Q_{t-1}$, the one-month-lagged $\ln Q_t$, captures the fact that current usage habits for a necessary good tend to be anchored in previous habits. In accordance with Eq. (2), the anchoring effect of past usage via partial adjustment is postulated to dampen over time.

Even though Hong Kong has an increasing block tariff for domestic users, we use the system average price, instead of the domestic marginal rate, to measure the water-rate variable. This is because *domestic* water usage comprises only about one-third of Hong Kong's total usage. Table 1 reveals that non-domestic usage is subject to fixed rates that do not vary with usage volume. Hence, the domestic marginal price is inappropriate for measuring the water-rate variable in our per capita analysis.

To be sure, the system average price may still arguably be endogenous. While the tariffs are administered rates that users pay, the system average price is partly dependent on the domestic users' usage decisions. If the system average price is correlated with the regression's error term, the regression's coefficient estimates are biased and inconsistent. An application of the Hausman test indicates, however, that the average-price variable is uncorrelated with the regression's error term, as was found by Woo (1992, 1994).

The infra-marginal price is not used as an explanatory variable for the following reasons. First, using the infra-marginal price requires disaggregate data that are unavailable for this analysis. Second, ignoring the infra-marginal price should not materially affect the coefficient estimates (Berndt, 1990; Woo, 1992). Finally, the usage effect of an infra-marginal price implicitly assumes good consumer understanding of price information (Gaudin, 2006). This assumption is tenuous for Hong Kong because the HKWSD uses the complicated rate structure shown in Table 1 to bill its customers once every four months.

3. Data

Table 2 presents descriptive statistics of our monthly data. Our main data source is the *Hong Kong Monthly Digest of Statistics*. The chosen sample, which makes full use of the available data, reflects our need for an updated estimate of Hong Kong's price elasticity of water usage. We use the aggregate data series because of our lack of access to monthly usage data by customer class.

While the income and weather data show some variations, the real water rate and its natural-log version have a tight range, since the nominal rate has been fixed since 1995. Hence, any change in the real rate over the last 15 years in our time series is due to changes in the value of the HK$, and principally inflation.

Fig. 1 shows Hong Kong's per capita water usage and real income for the 25-year sample period of April 1985–March 2010. This figure indicates rising per capita usage for the first 10 years, followed by a slowly declining trend. Moreover, the per capita usage exhibits seasonality, with higher usage in May–October than November–April. The figure also indicates a trend of rising real per capita income. The per capita income increase in April 2006–March 2010, however, is not matched by a similar increase in the per capita water usage. Taken together, Fig. 1 suggests that Hong Kong's per capita water usage

Table 2
Descriptive statistics of monthly data for estimating per capita water usage in Hong Kong for the 25-year period of April 1985 through March 2010.

Variable: Description	Mean	Standard deviation	Minimum	Maximum
Q_t: Monthly per capita water use (m^3/month)	11.83	0.91	8.14	13.93
$\ln Q_t$: Natural logarithm of Q_t	2.47	0.08	2.10	2.63
Y_t: Monthly real per capita income (HK$/month)	15,105	2285	9067	20,407
$\ln Y_t$: Natural logarithm of Y_t	9.61	0.16	9.11	9.92
P_t: Monthly average real rate (HK$/$m^3$)	2.59	0.11	2.30	2.83
$\ln P_t$: Natural logarithm of P_t	0.95	0.04	0.83	1.04
W_{1t}: Monthly average temperature (°C)	22.34	4.67	11.3	29.50
W_{2t}: Monthly sunshine (total hours of bright sunshine)	154.04	50.99	12.8	293.60
W_{3t}: Monthly rain (cm/month)	191.24	223.68	0.00	1346.10
W_{4t}: Monthly wind speed (km/h)	23.67	3.89	14.40	35.00

Source: Authors' calculations based on data from the *Hong Kong Monthly Digest of Statistics* and the Hong Kong Department of Water Supplies.

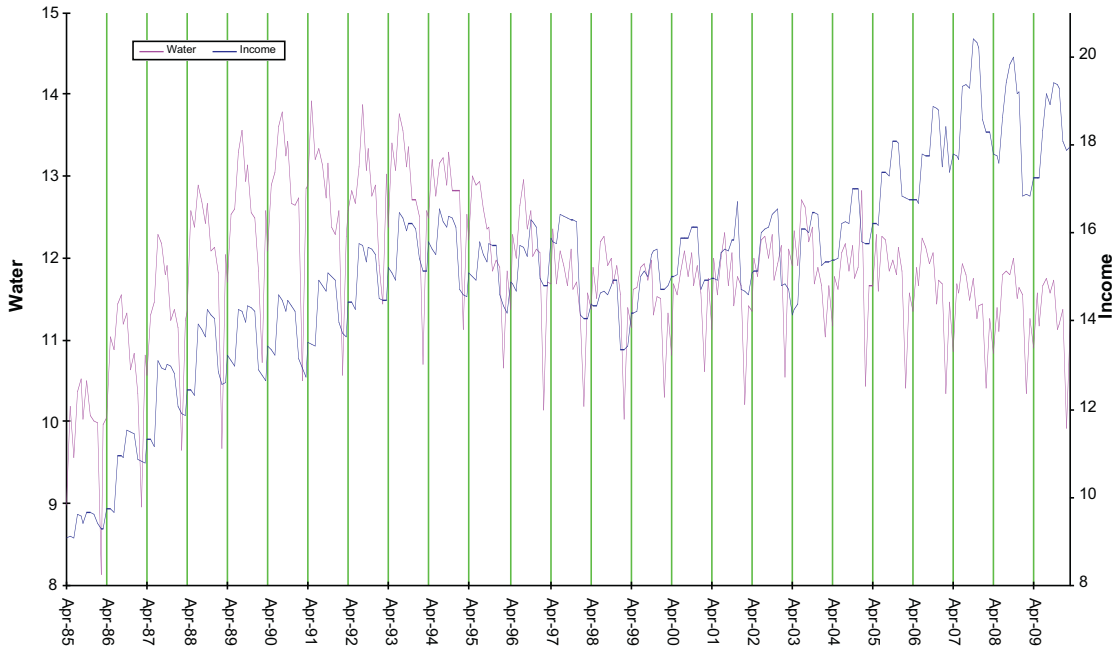

Fig. 1. Hong Kong's per capita water usage (m^3/month) and real income (HK$000/month) for the sample period of April 1985–March 2010 based on data from the *Hong Kong Monthly Digest of Statistics*.

moves with per capita income ($r = 0.27$) and has a recent mildly downward trend caused by non-income factors (e.g., water conservation by Hong Kong residents and resource management by the HKWSD).

4. Results

4.1. Validity of the model

Before presenting our regression results, we examine the empirical validity of Eq. (3). To do so, we use Table 3 that reports the Phillips–Perron unit root test results. This table shows that our main variables, in both their original and logarithmic formats, are stationary.

To further verify the validity of the model, we conduct the Phillips–Perron unit root test of the OLS regression residuals. Reported in Tables 4 and 5, these test results reject ($p < 0.001$) the null hypothesis of non-stationary residuals, thus eliminating any concerns about a spurious regression.

Finally, we apply the Ljung–Box Q statistic, which is a refinement of the Box–Pierce statistic, to test whether the residual series for Eq. (3) is white noise. We find that $Q(6) = 2.38$ and $Q(12) = 21.2$, with respective p-values of 0.8811 and 0.0476. The former statistic suggests that lag 1 to lag 6 of the residuals are not correlated, whereas the latter one tells us that lag 1 to lag 12 of the residuals are not significantly correlated at the 4% level. These results imply that the residual series is close to white noise and does not exhibit any serious autocorrelation.

Table 3
Phillips–Perron unit-root test statistics (*tau* with 3 lags).

Variable	Single mean	Trend
Q_t: Monthly per capita water use (m^3/month)	-8.04^{***}	-8.09^{***}
$\ln Q_t$: Natural log of Q_t	-8.40^{***}	-8.42^{***}
Y_t: Monthly real per capita income (HK$/month)	-2.83^{*}	-4.73^{***}
$\ln Y_t$: Natural log of Y_t	-3.06^{**}	-4.41^{***}
P_t: Monthly average real rate (HK$/$m^3$)	-3.82^{***}	-3.91^{**}
$\ln P_t$: Natural log of P_t	-3.85^{***}	-3.94^{**}
W_{1t}: Monthly average temperature (°C)	-7.41^{***}	-7.52^{***}
W_{2t}: Monthly sunshine (total hours of bright sunshine)	-11.79^{***}	-11.78^{***}
W_{3t}: Monthly rain (cm/month)	-11.54^{***}	-11.52^{***}
W_{4t}: Monthly wind speed (km/h)	-11.27^{***}	-11.59^{***}

Source: Authors' estimates based on data from the *Hong Kong Monthly Digest of Statistics* and the Hong Kong Department of Water Supplies.

* 10%.
** 5%.
*** 1%.

Table 4
Maximum likelihood estimation of the log-linear monthly per capita usage regressions for the 25-year sample period of April 1985–March 2010; "n.a." = "not applicable".

Variable	Model 1: Partial adjustment		Model 2: No partial adjustment	
	Estimate	*p*-Value	Estimate	*p*-Value
Total R^2	0.9268	n.a.	0.9226	n.a.
RMSE	0.0214	n.a.	0.0220	n.a.
AIC	−1420.29	n.a.	−1404.45	n.a.
Phillips–Perron unit-root test statistics for OLS regression residuals (*tau* with 3 lags)	−19.68	<.0010	−5.83	<.0010
Intercept: β	−0.3827	0.0183	−0.5272	0.4205
t	−0.0002	<.0001	−0.0005	0.0013
D_{1t}	−0.0113	0.1281	0.0189	0.0102
D_{2t}	−0.1341	<.0001	−0.1105	<.0001
D_{3t}	0.1073	<.0001	0.0195	0.0549
D_{4t}	−0.0344	0.0003	−0.0199	0.0622
D_{5t}	0.0407	0.0006	0.0353	0.0056
D_{6t}	−0.0299	0.0331	0.0142	0.3284
D_{7t}	0.0199	0.1434	0.0289	0.0231
D_{8t}	−0.0139	0.3193	0.0335	0.0084
D_{9t}	−0.0426	0.0009	0.0088	0.4500
D_{10t}	0.0063	0.5337	0.0226	0.0146
D_{11t}	−0.0519	<.0001	−0.0155	0.0065
$\ln Y_t$	0.0857	0.0002	0.3299	<.0001
$\ln P_{t-3}$	−0.0151	0.5997	−0.1318	0.0985
W_{1t}	0.0010	0.3815	0.0013	0.1780
W_{2t}	0.0001	0.2023	0.0001	0.0081
W_{3t}	−0.0000	0.1402	−0.0000	0.3487
W_{4t}	−0.0015	0.0018	−0.0007	0.0612
$\ln Q_{t-1}$	0.8452	<.0001		
AR(1) parameter: ρ	−0.1982	0.0021	0.8771	<.0001

Source: Authors' estimates based on data from the *Hong Kong Monthly Digest of Statistics* and the Hong Kong Department of Water Supplies.

4.2. Log-linear regression with partial adjustment

Table 4 reports the estimated results for Eq. (3), which we refer to as Model 1. The overall fit of Model 1 is quite good with a total $R^2 \approx 0.93$. The maximum likelihood estimates for the coefficients and their associated *p*-values support the following inferences.

Monthly per capita usage shows a small (−0.0002) and statistically-significant ($p < 0.0001$) declining trend. It also exhibits seasonality, as shown in the estimated coefficients for D_{kt}, six of which are statistically significant ($p < 0.01$). These findings support our first hypothesis of $\varphi \neq 0$ and $\delta_k \neq 0$ for some $k = 1, \ldots, 11$.

The coefficient estimate for $\ln Y_t$ is 0.0857, a statistically significant ($p = 0.0002$) short-run income elasticity, which supports our second hypothesis of $\beta_Y > 0$. The long-run income elasticity is $0.0857/(1 - 0.8452) = 0.55$, suggesting that Hong Kong' per capita usage is income-inelastic.

The estimated coefficient for $\ln P_{t-3}$ is −0.0151, which is not statistically significant, even at $\alpha = 0.10$. This short-run elasticity estimate has the hypothesized negative sign, but does not permit us to reject the alternative of $\beta_P \geq 0$ to our third hypothesis of $\beta_P < 0$. The long-run price elasticity estimate is $-0.0151/(1 - 0.8452) = -0.098$, suggesting that Hong Kong' per capita usage is highly price-inelastic.

Table 5
Maximum likelihood estimation of the log-linear monthly per capita usage regressions without price for the 25-year sample period of April 1985–March 2010; "n.a." = "not applicable".

Variable	Model 1′: Partial adjustment		Model 2′: No partial adjustment	
	Estimate	*p*-Value	Estimate	*p*-Value
Total R^2	0.9267	n.a.	0.9219	n.a.
RMSE	0.0214	n.a.	0.0220	n.a.
AIC	−1422.00	n.a.	−1403.53	n.a.
Phillips–Perron unit-root test statistics for OLS regression residuals (*tau* with 3 lags)	−19.65	<.0010	−5.67	<.0010
Intercept: β	−0.3911	0.0156	−0.5689	0.3812
t	−0.0002	<.0001	−0.0005	0.0008
D_{1t}	−0.0113	0.1282	0.0185	0.0116
D_{2t}	−0.1341	<.0001	−0.1109	<.0001
D_{3t}	0.1070	<.0001	0.0193	0.0578
D_{4t}	−0.0340	0.0003	−0.0190	0.0760
D_{5t}	0.0414	0.0005	0.0373	0.0035
D_{6t}	−0.0288	0.0376	0.0172	0.2346
D_{7t}	0.0205	0.1300	0.0285	0.0254
D_{8t}	−0.0131	0.3435	0.0332	0.0094
D_{9t}	−0.0420	0.0010	0.0087	0.4552
D_{10t}	0.0067	0.5119	0.0227	0.0147
D_{11t}	−0.0517	<.0001	−0.0160	0.0052
$\ln Y_t$	0.0860	0.0002	0.3216	<.0001
W_{1t}	0.0010	0.3958	0.0012	0.1956
W_{2t}	0.0001	0.2169	0.0001	0.0141
W_{3t}	−0.0000	0.1170	−0.0000	0.2198
W_{4t}	−0.0015	0.0020	−0.0008	0.0576
$\ln Q_{t-1}$	0.8421	<.0001		
AR(1) parameter: ρ	−0.1953	0.0023	0.8687	<.0001

Source: Authors' estimates based on data from the *Hong Kong Monthly Digest of Statistics* and the Hong Kong Department of Water Supplies.

In light of its lack of statistical significance, we further modify Model 1 of Eq. (3) by excluding $\ln P_{t-3}$ and re-estimating the parameters. We refer to this modified model as Model 1′ and report the results in Table 5. The result of a partial F test, with a test statistic of 1.516, affirms the inference that the lagged price variable in Model 1 does not contribute materially to the regression – or more affirmatively that water usage is strictly inelastic with respect to price, in the relevant range.

The coefficient estimates attached to the four weather variables ($W_{1t}, \ldots, W_{4t}$) indicate weather-dependence of per capita usage, even though only wind speed (W_{4t}) has a statistically significant ($p = 0.0018$) negative effect (−0.0015). This supports our fourth hypothesis.

The coefficient estimate for $\ln Q_{t-1}$ of approximately 0.85 is statistically different from zero ($p < 0.0001$), suggesting that each percentage increase in the past month's usage would tend to raise the current month's usage by 0.85 percent. This supports our fifth hypothesis of a partial-adjustment process. The coefficient estimate is also significantly different from unity, which also accords with the Nerlove assumption and a stationary adjustment process.

The estimate for the AR(1) parameter ρ is about −0.20 and statistically significant ($p = 0.0021$), suggesting an oscillating and dampening effect of a past random shock on current usage, which supports our sixth hypothesis and lends credence to our use of the AR(1) estimation procedure.

4.3. Additional considerations

We remark *en passant* that when estimating the regression with ordinary least squares, the only two parameter estimates with variance inflation factors of around 10 are those for the trend and per capita income variables. The variance inflation factor (VIF) of a variable, say trend, is $1/(1 - R^2)$ where R^2 is the coefficient of determination when fitting trend to all other independent variables. A VIF = 10 for trend implies that $R^2 = 0.90$, or 90 percent of the total variation of trend is explained by the total variation the regression for which trend is the regressand and the remaining variables are the regressors. Insofar as multicollinearity is a problem, its impact would be to exaggerate the estimated standard errors of the coefficients, which could lead to the erroneous failure to reject our null hypotheses. Nonetheless, the estimates of the regression coefficients are still the best linear unbiased estimates (BLUE) and inferences about the estimates and/or forecasts of the dependent variable are unaffected by multicollinearity. Thus, we keep both the trend and per capita income variables in Model 1. In the present application there are no close calls, so the possibility of failing to reject a null hypothesis that is false need not concern us here.

To determine if Model 1 is a reasonable specification for characterizing the monthly per capita usage data, we have considered a number of alternatives. First, we estimated a log-linear model without partial adjustment.[10] Referred to as Model 2 in Table 4, this model is not preferable to Model 1 because of its higher Akaike Information Criterion (AIC) value

[10] We thank the referee for suggesting this model and its interpretation.

Table 6
Maximum likelihood estimation of the log-linear monthly per capita usage regressions for the 25-year sample period of April 1985–March 2010; "n.a." = "not applicable".

Variable	Model 1″: Partial adjustment		Model 2″: No partial adjustment	
	Estimate	*p*-Value	Estimate	*p*-Value
Total R^2	0.9290	n.a.	0.9272	n.a.
RMSE	0.2115	n.a.	0.02138	n.a.
AIC	−1425.49	n.a.	−1419.24	n.a.
Phillips–Perron unit-root test statistics for OLS regression residuals (*tau* with 3 lags)	−18.21	<.0010	−7.93	<.0010
Intercept: β	−0.8547	0.0003	−2.3872	<.0001
t	−0.0003	<.0001	−0.0008	<.0001
D_{1t}	−0.0058	0.4318	0.0260	<.0001
D_{2t}	−0.1284	<.0001	−0.1017	<.0001
D_{3t}	0.0990	<.0001	0.0279	0.0022
D_{4t}	−0.0343	0.0003	−0.0181	0.0721
D_{5t}	0.0365	0.0021	0.0372	0.0022
D_{6t}	−0.0295	0.0346	0.0175	0.2081
D_{7t}	0.0144	0.2903	0.0244	0.0556
D_{8t}	−0.0142	0.3050	0.0305	0.0172
D_{9t}	−0.0414	0.0012	0.0072	0.5342
D_{10t}	0.0040	0.6911	0.0199	0.0344
D_{11t}	−0.0496	<.0001	−0.0169	0.0036
D_{2003t}	1.1075	0.0037	3.4725	<.0001
$\ln Y_t$	0.1629	<.0001	0.5282	<.0001
$D_{2003t} \times \ln Y_t$	−0.1137	0.0036	−0.3574	<.0001
$\ln P_{t-3}$	−0.0436	0.1834	−0.1438	0.0548
W_{1t}	0.0015	0.1875	0.0016	0.0994
W_{2t}	0.0001	0.0736	0.0001	0.0122
W_{3t}	−0.0000	0.2113	−0.0000	0.1283
W_{4t}	−0.0014	0.0037	−0.0008	0.0505
$\ln Q_{t-1}$	0.7461	<.0001		
AR(1) parameter: ρ	−0.1299	0.0642	0.7564	<.0001

Source: Authors' estimates based on data from the *Hong Kong Monthly Digest of Statistics* and the Hong Kong Department of Water Supplies.

(Akaike, 1974). The coefficient estimates of Model 2, however, corroborate those found for Model 1. In particular, the statistically insignificant ($\alpha = 0.05$) coefficient estimate for $\ln P_{t-3}$ based on Model 2 is about −0.13, which is similar to the long-run elasticity estimate of −0.098 based on Model 1. Thus, it appears that the AR parameter in the disturbance term picks up the lagged consumption's effect instead.

Next, we re-estimated Models 1 and 2 by replacing P_{t-3} with one of the following variables: P_t, P_{t-1}, P_{t-2}, or P_{t-4}. All these regressions produce counter-intuitive positive price-elasticity estimates that are statistically insignificant ($\alpha = 0.10$). Hence, these measurements are not used to represent the price signal given to water users.

Finally, we performed a Chow test, whose results, $F = 3.87$ and $F = 12.68$ for Models 1 and 2, respectively, indicate a statistically significant ($\alpha = 0.001$) structural change around 2003, as suggested by Fig. 1. Hence, we re-estimated the models of Table 4 after including two additional explanatory variables: (a) a binary indicator D_{2003t} that is equal to unity if the monthly observation is after 2002 and is zero otherwise; and (b) an interaction term $D_{2003t} \times \ln Y_t$ that allows the short-run income elasticity for the period after 2002 to be different from the one for the prior period.

Referred to as Model 1″ and Model 2″ in Table 6, the regression results indicate that the SAR's per capita usage had become less income-sensitive after 2002. That is, for the partial-adjustment Model 1″, prior to 2003 the income elasticity is 0.1629 ($p < 0.0001$). From 2003 forward, however, it is $(0.1629 - 0.1137) = 0.0492$ ($p = 0.0356$). Eliminating the partial-adjustment term, the income-elasticity estimates of Model 2″ are 0.5282 ($p < 0.0001$) before 2003 and $(0.5282 - 0.3574) = 0.1708$ ($p = 0.0127$) thereafter. Thus, the income-elasticity estimates reflect what we observe in Fig. 1: notably, per capita income continuing its upward path after 2002, accompanied by a relatively stagnant water-usage series that diverges away from it.

The price-elasticity estimates, however, remain small in size and statistically insignificant. The β_P estimate for Model 1″ is −0.0436 ($p = 0.1834$) and −0.1438 ($p = 0.0548$) for Model 2″.

Taken together, the income and price-elasticity estimates based on Models 1″ and 2″ do not alter our view that water pricing will unlikely be an effective strategy to curb the SAR's rising total water usage caused by its income and population growth.

5. Conclusions

Looming water shortages and the need to manage our water resources are world-wide concerns that extend far beyond the borders of Hong Kong. Seckler, Upali, Molden, Radhika, and Barker (1998, p. 17) were scarcely unique in concluding that "many countries are entering a period of severe water shortage" with "water-deficit countries and regions including not only

West Asia and North Africa but also some of the major breadbaskets of the world such as the Indian Punjab and the central plain of China." The problem is particularly acute in China because of its huge and growing population of 1.4 billion that is engaged in an on-going process of relocation from low-density rural areas to high-density urban areas, reflective of a nation transitioning from an agriculture-based economy to one that increasingly depends upon manufacturing exportable products for its remarkable economic growth (Shalizi, 2006).

Hong Kong is an important component in that nation and its experience and how this bastion of *laissez-faire* economics deals with its water issues should and will command world-wide attention. Perhaps most notably, if the price mechanism is not a principal component of the SAR's demand-management policy, other urban areas and nations might take their cautious cues from Hong Kong and direct their water-management efforts in other directions as well.

But *should* price play a prominent, albeit not the principal, role in the HKWSD's efforts to induce water conservation? Our analysis of the most recent 25 years of the Hong Kong experience shows the primary drivers of water usage in the SAR to be the ability to pay, the water-usage standards to which the populace has grown accustomed, and, subject to the territory's more efficient and educated use of its water, seasonal and weather-related factors. Price was most notable for the absence of *any* non-negligible effect on usage behavior, either in the short- or long-run.

Water management problems worldwide will only be exacerbated by global warming, growing populations, and economic growth. It is therefore incumbent upon policy makers to deal with and anticipate those problems before they occur and before they get out of hand. We commend the HKWSD for having adopted a TWM strategy, in lieu of what would have been an unpopular and ineffective pricing strategy, as a tool for water-usage management, and we would urge policy makers on the mainland and across the globe to take note and act accordingly.

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