DOES PRICE STRUCTURE MATTER? HOUSEHOLD WATER DEMAND UNDER INCREASING-BLOCK AND UNIFORM PRICES

Sheila M. Olmstead
Yale University, School of Forestry and Environmental Studies

W. Michael Hanemann
University of California, Berkeley
Department of Agricultural and Resource Economics, and Goldman School of Public Policy

Robert N Stavins
John F. Kennedy School of Government, Harvard University
and Resources for the Future

March 12, 2003

Please Contact:
Professor Sheila M. Olmstead
School of Forestry and Environmental Studies
Yale University
360 Prospect Street
New Haven, Connecticut 06511
Phone: 203-432-6247
Fax: 203-432-3809
E-Mail: sheila.olmstead@yale.edu

ABSTRACT

We carry out an empirical analysis of the influence of the price of water and the structure of water prices on residential water demand. We adapt a model from the labor economics literature – the Hausman model of labor supply under progressive income taxation – to estimate water demand under increasing-block prices. We apply this structural model to the most price-diverse, detailed, household-level water demand data yet available to estimate the price elasticity of residential water demand. Our results indicate that the sensitivity of residential water demand to price is quite low, in contrast with results of previous studies using similar models to account for the piecewise-linear budget constraint of block prices. We also find, however, that price elasticity is higher and demand is lower among households facing block prices than among households facing uniform marginal prices. The impact of the price structure on demand appears to be greater than the impact of marginal price itself.

DOES PRICE STRUCTURE MATTER? HOUSEHOLD WATER DEMAND UNDER INCREASING-BLOCK AND UNIFORM PRICES

Sheila M. Olmstead, W. Michael Hanemann, and Robert N. Stavins[*]

1. INTRODUCTION

In large parts of the United States, scarce water supplies are a serious resource and environmental concern. In some cases, water is being used at rates that may exceed those dictated by efficiency criteria, particularly when externalities and the existence of water supply subsidies are taken into account. As a result, some policymakers have promoted the use of demand management techniques, including requirements for the adoption of specific technologies and restrictions on particular uses. A natural question for economists to ask is whether price would be a more cost-effective instrument to facilitate water demand management.[1]

Much of the water demand literature is rooted in engineering rather than economics, and there is widespread belief among water managers that consumers do not respond to price signals. This, in conjunction with low price elasticity estimates in the literature, may have contributed to the predominant use of command-and-control instruments in water demand management. Yet, empirical evidence of the effectiveness of such policies is mixed, at best. Some studies have failed to identify statistically significant impacts (Schultz et al. 1997), while others have identified only small negative impacts on residential water demand (Mayer et al. 1998, Kiefer et al. 1993,

[*]Olmstead is an Assistant Professor of Environmental Economics at the Yale School of Forestry and Environmental Studies. Hanemann is the Chancellor's Professor in the Department of Agricultural and Resource Economics and in the Goldman School of Public Policy, University of California, Berkeley. Stavins is the Albert Pratt Professor of Business and Government, Chair of the Environment and Natural Resources Faculty Group, and Director of the Environmental Economics Program, John F. Kennedy School of Government, Harvard University, and a University Fellow of Resources for the Future. This work was supported in part by a grant from the National Science Foundation, Economics Program. The authors are grateful to Jeffrey Liebman, Tony Gómez-Ibáñez, Frank Wolak, and Richard Zeckhauser for their comments on an earlier draft, and to Julie Hewitt, Ellen Pint, and Bob Triest for sharing methodological advice. The authors, alone, are responsible for any errors.

[1]While the comparison of market-based and command-and-control approaches to environmental policy has a long history (Pigou 1920, Crocker 1966, Dales 1968, Montgomery 1972, Baumol and Oates 1988), we know of few studies that have addressed water conservation policies in this framework (Michelsen et al. 1998, Corral 1997). This is surprising, given that energy policies have frequently been examined in this context (Jaffe and Stavins 1994a, 1994b, Stelzer 1980, Schipper 1979, Greene 1990, Hirst and Goeltz 1984, Wirl and Orasch 1998, Wirl 1997).

Renwick and Green 2000).[2]

A review of the literature suggests that water demand is sensitive to price, but that the magnitude of that sensitivity is small at current prices. In a meta-analysis of 124 estimates generated between 1963 and 1993, accounting for the precision of estimates, Espey et al. (1997) obtain a mean price elasticity of -0.51, a short-run median estimate of -0.38, and a long-run median estimate of -0.64. While there are some outliers, 90 percent of estimates during this time period were between 0 and -0.75 (Espey et al. 1997).[3]

But one recent study that modeled directly the piecewise-linear budget constraint created by block pricing, drawing upon an approach developed by Hausman for estimating labor supply effects of progressive income taxation, obtained the largest price elasticity estimates in the literature, ranging from -1.57 to -1.63 (Hewitt and Hanemann 1995).[4] The result caused some to consider whether previous studies had simply underestimated price elasticity through incorrect modeling.[5]

We examine the possibility that price structure, as well as the magnitude of marginal price, may significantly affect demand. Our data comprise 1,082 households in 11 urban areas in the United States and Canada, served by 16 water utilities. There are 26 price structures in the data; eight two-tier increasing-block structures, ten four-tier increasing block structures, and eight uniform structures. Previous water demand studies have not exploited such substantial cross-sectional variation in the price incentives faced by households to estimate price elasticities.

Our results suggest that the high elasticities reported in previous applications of the Hausman model to water demand were unlikely due to the model itself. Through simulations

---

[2]One comprehensive study of aggregate water district consumption in California found small but significant reductions in total water use attributable to landscape education programs and watering restrictions, but no effect due to conservation education programs, low-flow fixture distribution, or the presentation of drought and conservation information on customer bills (Corral 1997).

[3]Hanemann (1997a) lists 99 water studies for urban areas in North America between 1951 and 1991. A simple average of those estimates, using midpoints of ranges where ranges are reported, is -0.47. With three outliers excluded, the average is -0.44. This rough average does not account for the precision of estimates, but we include it here as an additional point of reference.

[4]Using a related specification, Pint (1999) obtains estimates ranging from -0.04 to -1.24, depending upon season. Prior to Hewitt and Hanemann (1995), only three studies had estimated price elasticities in the elastic range (Howe and Lineweaver 1967, Danielson 1979, Deller et al. 1986).

[5]Over the past 40 years, economists have improved water demand estimation through: correction for the endogeneity of price and quantity under block rate pricing; proper specification of marginal price and implicit income effects due to changes in infra-marginal rates, and the use of time series rather than cross-sectional data (Howe and Lineweaver 1967, Billings 1982, Foster and Beattie 1979, Chicoine and Ramamurthy 1986, Nieswiadomy and Molina 1989). While early water demand studies used average price as an explanatory variable, Gibbs (1978) first argued for the use of marginal price.

based upon our econometric model, we obtain a full-sample price elasticity estimate of -0.33, well within the range of the previous literature on water demand. The presence of block pricing appears to affect both water demand and price elasticity. We cannot rule out the possibility that the difference in price elasticities is due to some factor other than true consumer response to different price structures, such as differences in characteristics omitted from the water demand model. Nor have we ruled out the possibility that water rate structures are themselves endogenous (communities more sensitive to price changes might also be those that adopt increasing-block structures). In the future, determining whether water rate structure influences price elasticity will have important implications for water demand management, as well as for other areas where piecewise-linear budget constraints are the norm.

The paper is organized as follows. In section 2, we review the economic theory of block pricing. Section 3 reviews possible explanations for variation in price elasticity of residential water demand. Section 4 examines econometric techniques for estimating water demand functions. Section 5 describes the data, and Section 6 describes our model of water demand under increasing-block and uniform marginal water prices. In Section 7, we describe our results, and in Section 8 we conclude.

2. BLOCK PRICING AND ECONOMIC THEORY

Urban residential water service pricing typically takes one of three forms in the United States: (1) constant or uniform rates; (2) increasing block rates; or (3) decreasing block rates. Each of these price structures is typically accompanied by a fixed water service fee. Under constant or uniform rates, households are charged a single volumetric marginal price at all levels of consumption. Increasing block structures charge higher marginal prices for higher quantities consumed, resulting in a water supply function that resembles a staircase ascending from left to right (Figure 1); decreasing block structures are stacked in the opposite direction.

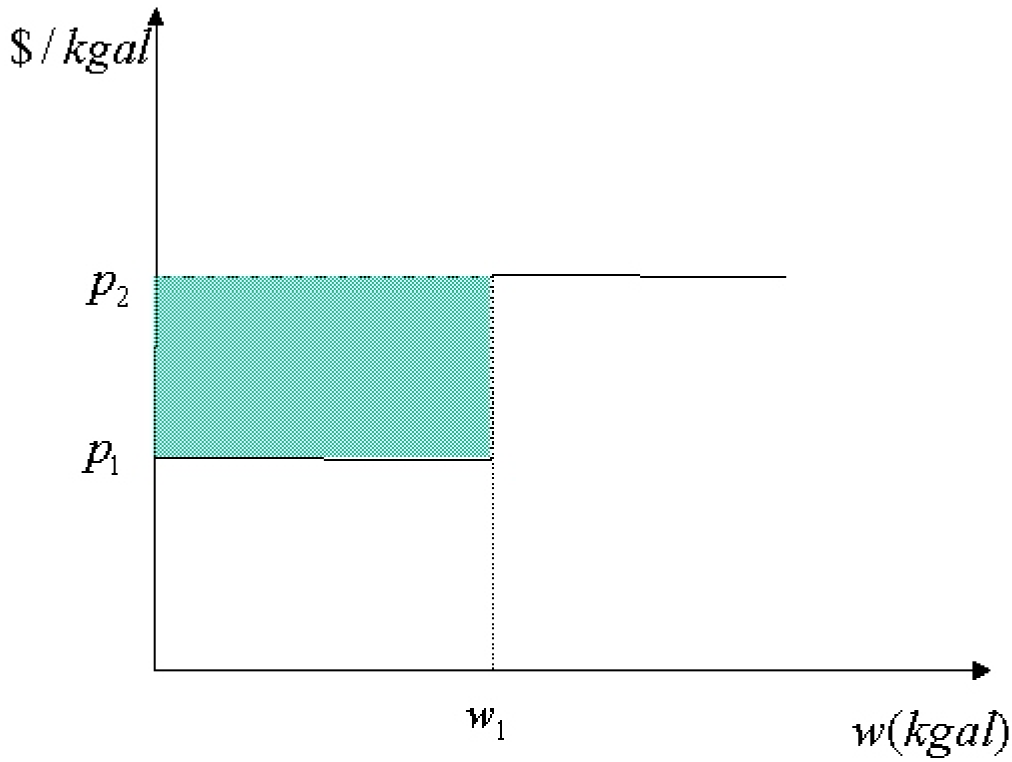
2.1 Piecewise linear budget constraints

Under block pricing, consumers face a piecewise-linear budget constraint.[6] Figure 2 depicts the budget constraint in a simple two-tier increasing block price system. The consumer has three reasonable consumption choices: consume on the interior of segment one, on the interior of segment two, or at the kink point – the quantity at which the marginal price increase occurs.[7]

---

[6]Such budget constraints arise in many economic settings, including government tax and transfer programs, water and electricity pricing, and volume discounts.

[7]Under decreasing-block structures, budget sets are non-convex, enabling multiple tangencies between consumer indifference curves and the budget constraint. While slightly more complex, the theory and empirics of the non-convex case are straightforward extensions of the convex case described here. The decreasing block case has been developed generally (Moffit 1986, 1990); for the case of water supply (Hewitt 1993, Hewitt and Hanemann 1995); and in the literature on labor supply and taxation (Burtless and Hausman 1978, Hausman 1985). None of the

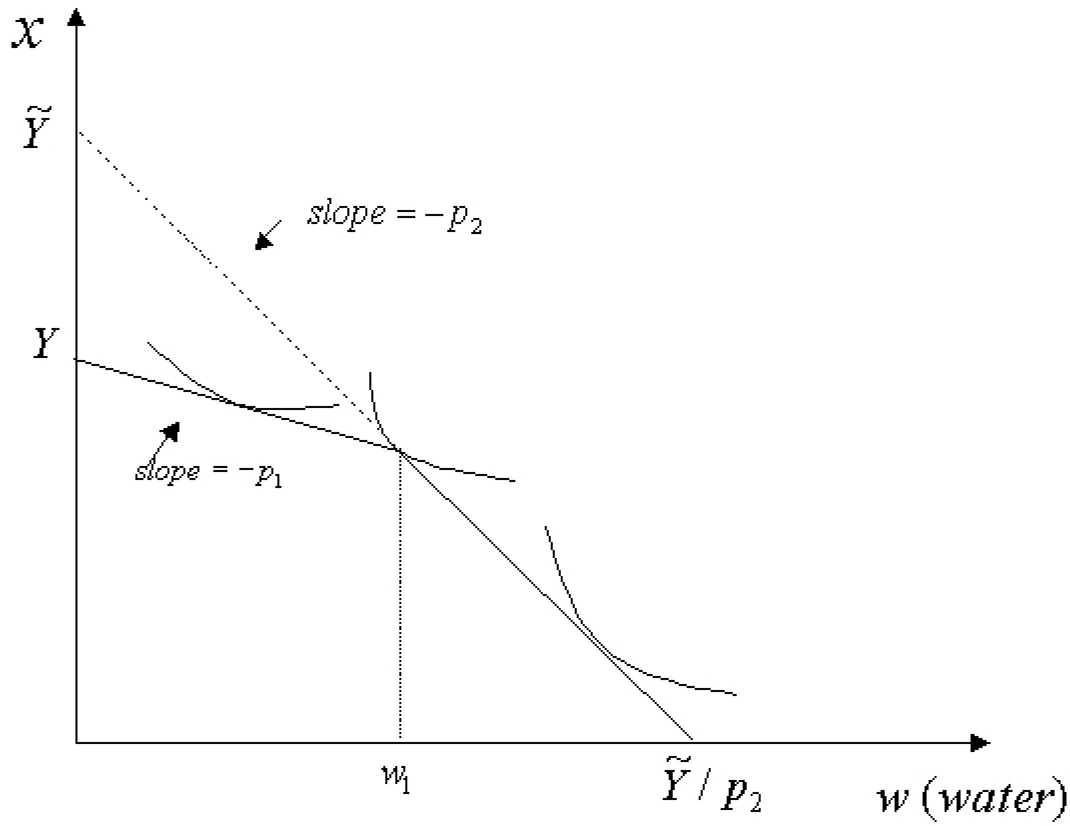


Where: w = water (units)
p_1 = marginal price of water in block 1
p_2 = marginal price of water in block 2
w_1=boundary quantity between block 1 and block 2

Figure 1. Two-tier increasing block price structure

---

utilities in our sample have decreasing block rates.



Where: Y = income
$\tilde{Y}$ = virtual income
p_1 = price of water in block 1
p_2 = price of water in block 2
w_1 = boundary quantity between block 1 and block 2

Figure 2. Utility Maximization under a Two-tier Increasing Block Price Structure

In theory, consumers equate the marginal price of water to the marginal benefit of water consumption in choosing the quantity to consume. But for households consuming anywhere on a piecewise linear budget constraint, with the exception of the first linear segment, the marginal price is not the price paid for every unit consumed. Marginal price is still the relevant price to include in a water demand function, but we must account for the implicit subsidy that derives from the block rate structure. We do so by adding to their income the difference between what they would pay if all units were priced at the price of the last unit consumed, and what they actually pay. This income supplement for a household consuming in the second block of a two-tier price structure is equal to the shaded region in Figure 1. Following standard practice, we use the term "virtual income" to refer to income so supplemented to account for infra-marginal prices.[8]

The kinks in the budget constraint create a more complicated relationship between price, income, and quantity than the relationship that results from a simple, linear budget constraint. The expected negative relationship between price and quantity demanded and positive relationship between income and quantity demanded will hold within blocks, but marginal income and price effects may be zero for households consuming at a kink (Moffitt 1990). The resulting consumer demand can be discontinuous, as consumption can stick at kinks in the budget constraint or, alternatively, change abruptly and non-marginally from one block to another. Thus, piecewise linear budget constraints require careful econometric treatment.

2.2 Piecewise linear budgets: water demand vs. labor supply

Labor supply under progressive income taxation and water demand under increasing block prices share the theoretical structure implied by piecewise linear budget constraints but differ in important practical ways. A controversial result of Hausman's labor supply models relevant to the discussion here is his estimation of a significant negative compensated wage elasticity of labor supply, indicating substantial deadweight losses from income taxation (Hausman 1981). Earlier studies had found uncompensated wage elasticities close to zero and thus concluded that income taxation had negligible effects on desired work hours. Through the use of his model accounting for the piecewise linear budget constraints created by progressive income taxation, Hausman recognized that the uncompensated effect was, in fact, close to zero because the income and substitution effects of a tax increase could effectively cancel each other out. Hausman's result is, in large part, attributable to the implicit lump-sum transfers to virtual income from the progressive tax schedule, which are quite large in labor supply analysis (Burtless 1981).

In the case of water demand, we measure a small, negative uncompensated price elasticity. In contrast to the labor supply case, the income and substitution effects of a water price increase

---

[8]Use of virtual income was developed in the labor supply literature (Hall 1973, Burtless and Hausman 1978), and in the electricity demand literature(Nordin 1976, Taylor 1975). It was first applied to the estimation of water demand functions by Billings and Agthe (1980). Without including virtual income in demand estimation, households facing the same marginal price but different infra-marginal rates would receive identical treatment, and shifts in prices below the marginal price would be modeled as if they had no effect. Virtual income treats the infra-marginal rate changes as lump-sum changes in household income.

operate in the same direction: water is relatively more expensive, and real income decreases, so consumers purchase less for both reasons. One would, therefore, expect the compensated price effect to be smaller than the uncompensated effect—not larger as in the case of labor supply. The increase in real income from the (very small) implicit lump-sum transfer due to the price schedule could, in theory, also diminish the compensated price effect relative to the uncompensated effect. But attempts to measure effects of these small infra-marginal rate income transfers in isolation have been unsuccessful, both in electricity and water demand (Hausman et al. 1979, Deller et al. 1986, Nieswiadomy and Molina 1989).[9]

3. Block Pricing and Price Elasticity

Most sources of variation in price elasticity of demand are related to the availability of substitutes. Apart from substitution possibilities that change the shape of the demand curve (and technological changes that may shift it inward or outward), elasticity may vary along the demand curve. With linear demand, for example, price elasticity of demand is declining in price. In water demand estimation, higher elasticity estimates are associated with long-run analyses, models using average rather than marginal price, data restricted to summer observations, and models of areas with increasing block rates (Hanemann 1997b, Espey et al. 1997).[10]

The greater elasticity results for summer uses (which presumably include a higher irrigation component) and long-run analyses can easily be justified by economic theory. Higher elasticity estimates for samples facing increasing-block rates are somewhat harder to understand. If consumers react to the price structure, as well as the marginal price itself, in determining their level of water consumption, there may be some behavioral explanation. For example, households that trigger a higher marginal price for use beyond some level of consumption may pay more attention to price than households that do not. If this were true, precise information about the level of prices would not be needed to cause an effect; even a vague notion that prices increase beyond some threshold would suffice to focus consumers' attention and increase price elasticity. In addition, it is possible that households facing block prices think that the cost of water consumption is higher than it really is— they may focus on the most expensive block price, for example. There is some evidence that consumers facing block pricing react to something other than marginal price (Shin 1985, Nieswiadomy and Molina 1991).

---

[9]Even in the fourth block in the sample cities analyzed here, where implicit lump-sum transfers are at their greatest, their average contribution to virtual income is less than $584 per year.

[10]In addition, one analysis estimated the effect of shifting from flat marginal pricing to an increasing rate structure. This is quite different from an increasing block structure, in that water consumption beyond some established cutoff increases the marginal price of *all* units consumed, not just those consumed in excess of the cutoff. Nonetheless, they estimate that the introduction of the increasing rate structure in the Maryland suburbs of Washington, D.C. in 1978 was responsible for reductions of 1.1 to 8.7 percent in quarterly household water demand (Young et al. 1983).

An alternative explanation is unobserved heterogeneity among consumers in cities implementing different price structures, if that heterogeneity is correlated with factors that affect elasticity. Households in cities with block prices may use a greater portion of total water consumption for purposes with substitution possibilities, leading to relatively more elastic demand in those cities. For example, households in block-price cities may use more water for irrigation (as a percentage of total use) due to arid conditions or longer growing seasons (Hewitt 2000). In addition, communities that implement block pricing may have experienced more significant price increases over time than communities with uniform prices, due to water scarcity or other factors. If this is true, then block-price households may have invested in technologies that make them more price-responsive. Higher current prices among block-price cities could also lead to higher elasticity estimates for those cities if we assume there is one demand curve for all consumers, and that elasticity is not constant.

While few studies have examined the potential influence of price structure on price elasticity, some have suggested a causal link (Hewitt 2000, Espey et al. 1997). Because differences in price elasticity across price structures may be due either to behavioral responses to price structure, or to heterogeneity, it is necessary to test for both possibilities.

4. Econometric Modeling of Demand Under Block Pricing

Non-linear budget constraints present challenges for the econometric estimation of demand functions. First, the discrete choice of block or kink and the continuous choice of quantity, made simultaneously, must both be modeled. In addition, the choice of block, hence quantity, and marginal price are endogenous. If a typical single-error stochastic specification were employed, the size of the error term, marginal price, and virtual income (a function of marginal price) would be systematically correlated.

4.1 Endogeneity of price and quantity

Under piecewise-linear budget constraints, ordinary least squares (OLS) estimates of the parameters of the demand function will be biased and inconsistent, due to the simultaneous determination of price and the block of consumption. A larger value of the error term will increase the likelihood that water consumption will be observed in a higher block (at a higher marginal price), and the opposite will be true for a small value of the error term. This problem was first treated by estimating simultaneous equations models, such as two-stage least squares (2SLS) and other instrumental variables models, for both labor supply and energy demand (Hausman et al. 1979, Agthe et al. 1986, Deller et al. 1986, Nieswiadomy and Molina 1988, 1989).

4.2 Insufficiency of traditional simultaneous equations models

While such models can address the problem of endogeneity and result in parameter estimates that reflect a downward-sloping demand curve, they do not model both portions of the consumption choice — the discrete choice of block or kink and the continuous choice of quantity. As a result, the effect of changes in the price structure, such as a shift from two blocks to three, or a change in the consumption quantity threshold that moves a household from one block to another, cannot be assessed. The true price elasticity under block rates includes both the conditional elasticity of demand given consumption within a block, and the elasticity of the probability of consuming within that block.

In addition, the usual IV methods disregard the fact that some households are observed consuming at or within the neighborhood of a kink point, where it is unclear which value of marginal price should be assigned to these observations (in an econometric model that includes one or more error terms). Arbitrary treatment of these observations, such as assigning them to one block or another or dropping them from the sample ignores the utility maximization process operating behind the demand curve.

4.3 The discrete-continuous choice model

The discrete-continuous choice (DCC) model, a maximum likelihood model, addresses these theoretical and econometric issues associated with block pricing. The model was first applied to water demand in the early 1990s (Hewitt 1993, Hewitt and Hanemann 1995), although the original development was for labor supply functions under graduated marginal income tax rates (Burtless and Hausman 1978), and subsequent generalizations noted the potential application of the model to water and electricity demand (Moffitt 1986, 1990).

The DCC model is a probability statement in which each observation is treated as if it could actually have occurred at any kink or linear portion of the household's budget constraint. The probability statement for an individual observation is a sum of joint probability statements, one for each kink and linear block in the budget constraint. Each joint probability includes the probability of the continuous choice of quantity consumed and the conditional probability that consumption occurred at that kink or block, given the choice of quantity. The form of the likelihood function differs depending on the price structure faced by each household; our data include households that face a uniform marginal price for water, as well as two- and four-tier increasing block price structures. Maximizing the probability statement, in the form of a likelihood function, generates the parameter estimates.

One major critique of this approach in labor supply analysis has been that the locations of consumers' budget constraint kink points are uncertain – they depend on information about income tax itemizations and deductions, typically unknown – creating measurement error presumed to bias estimates (Heckman 1983, Gan and Stahl 2002). In water demand analysis, the critique is irrelevant because budget constraint kink points correspond to administratively-set block cutoffs and, thus, are known with certainty.[11]

[11]A further critique of the DCC model in labor supply has been that results are sensitive to assumptions about the functional form of demand and distribution of error terms (Heckman 1983, Blomquist and Newey 2001). This may

5. DATA DESCRIPTION

The data comprise 1,082 households in 11 urban areas in the United States and Canada, served by 16 water utilities.[12] This group of areas is notable for its geographic and climatic diversity. Table 1 provides descriptive statistics for water demand model variables.

Daily household water demand and weather conditions are observed over two periods of two weeks each, one in the outdoor irrigation season and one in the non-irrigation season. For each household, less than 12 months passed between the two observation periods. Daily demand data were gathered by automated data loggers, attached to magnetic household water meters by utility staff. These devices were hidden from sight during most, if not all, uses of water. Daily weather observations were drawn from local data collection stations.

Each household faces one price structure throughout each season of observation, but many price structures changed between the two periods. As a result, there are 26 price structures in the data; eight two-tier increasing block structures, ten four-tier increasing block structures, and eight uniform structures. Marginal prices range from $0.00 per thousand gallons (kgal) for the first 4.5 kgal in Phoenix, to $4.96 per kgal in the most expensive block in the Las Virgenes Municipal Water District.

Data on characteristics of individual households, including gross annual household income, family size, and home age and size were collected by a one-time household survey. Households chosen for the study were randomly sampled from a subset of utilities' customer databases: residential single-family households. Surveys were anonymous and were field-tested prior to distribution. Sampling procedures, response rates, and statistical tests for differences between respondents and all single-family customers are described in Mayer et al. (1998).[13]

---

be true in water demand, as well. Determining the robustness of our results to changes in underlying assumptions is an important area for further research. Nauges and Blundell (2001) estimate a non-parametric model of water demand under block pricing, obtaining results that differ from their maximum likelihood results. They do not, however, estimate multiple maximum likelihood models to determine whether distributional and functional form assumptions are problematic.

[12]The household-level data used in this study were gathered by Mayer et al. (1998) for a study sponsored by the American Waterworks Association. See Table 4 for a list of the areas included.

[13]For an explanation of the treatment of top-censored categorical variables and missing observations, see Cavanagh (Olmstead) 2002.

Table 1. Variable Descriptions and Summary Statistics

Variable	Description	Units	Mean	Std. Dev.	Min.	Max.
w	Daily household water demand	kgal/day	0.40	0.58	0.00	9.78
p_1	Marginal price in block 1	\$/kgal	1.45	0.54	0.50	3.70
p_2	Marginal price in block 2	\$/kgal	1.84	0.40	0.84	4.06
p_3	Marginal price in block 3	\$/kgal	2.43	0.87	0.93	4.96
p_4	Marginal price in block 4	\$/kgal	3.28	1.30	0.99	5.98
y	Gross annual household income	\$000/yr	69.81	67.67	5.00	388.64
yd_2	Virtual income, block 2	\$000/yr	77.20	78.45	5.00	388.71
yd_3	Virtual income, block 3	\$000/yr	121.63	107.48	5.01	388.71
yd_4	Virtual income, block 4	\$000/yr	122.15	107.62	5.02	389.52
w_1	Water quantity at kink 1	kgal/day	0.21	0.08	0.06	0.37
w_2	Water quantity at kink 2	kgal/day	0.36	0.09	0.30	0.50
w_3	Water quantity at kink 3	kgal/day	1.82	0.80	0.75	2.49
seas	Season: 1 if irrigation season; 0 if not	0/1	0.51	0.50	0	1
weath	Evapotranspiration - effective rainfall	mm	5.06	8.42	-46.15	19.37
maxt	Maximum daily temperature	°C	24.12	8.78	0	42.78
famsz	Number of residents in household	integer	2.79	1.34	1	9
bthrm	Number of bathrooms in household	integer	2.58	1.30	1	7
sqft	Approximate area of home	000 ft^2	2.02	0.82	0.40	4.37
lotsz	Approximate area of lot	000 ft^2	10.87	9.22	1.00	45.77
age	Approximate age of home	yrs/10	2.88	1.62	0.07	5
evap	Evaporative cooling: 1 if yes; 0 if no	0/1	0.09	0.28	0	1
lasv	Indicator - Las Virgenes MWD	0/1	0.09	0.29	0	1
seat	Dummy - Seattle	0/1	0.10	0.28	0	1
sandg	Dummy - San Diego	0/1	0.09	0.28	0	1
tampa	Dummy - Tampa	0/1	0.09	0.29	0	1
phx	Dummy - Phoenix	0/1	0.09	0.29	0	1
tscot	Dummy - Tempe, Scottsdale	0/1	0.09	0.29	0	1
wcamb	Dummy - Waterloo, Cambridge	0/1	0.08	0.28	0	1
wvall	Dummy - Walnut Valley WD	0/1	0.10	0.29	0	1
lomp	Dummy - Lompoc	0/1	0.09	0.29	0	1

We anticipate that price and income will affect demand in the usual ways. In addition, most household water demand, with the exception of the small fraction used for drinking water, is actually derived demand in which the primary demand is for water-consuming goods and services, such as clean laundry, indoor bathroom use, and green lawns. As a result, household water demand also depends on characteristics that represent the household's tastes for water consumption in such services. We expect daily demand to be positively correlated with lot size, square footage of homes, number of bathrooms, and family size.

We include one idiosyncratic housing variable in the models—the presence or absence of an evaporative cooler (*evap*). Low-income households in arid areas frequently have evaporative coolers, rather than traditional air conditioning, which essentially substitutes water for electricity to produce cooling. In our sample, households with evaporative coolers use on average 40 percent more water per day. We include this variable to avoid biasing the income coefficient estimate.[14]

As we have specified the daily weather variables, each should be positively correlated with demand. Maximum daily temperature is represented by *maxt*, and *weath* represents the moisture requirements of green lawns not met by precipitation.[15] Given the very different distributions of water demand in the two seasons, we expect that the influence of season (*seas*), a dummy variable set equal to one during the outdoor watering season, will be positive.

We include home age in the water demand equation and expect the relationship to be non-monotonic. Very old homes are likely to have smaller connections to their city water system, and also fewer water-using appliances such as dishwashers and jacuzzis than do newer homes. The very newest homes were built after the passage of local ordinances in the 1980s and 1990s requiring various water-conserving fixtures.[16]

Finally, the models include a set of dummy variables that represent the 11 urban areas included in the study: Denver, Eugene, Seattle, San Diego, Tampa, Phoenix, Tempe/Scottsdale, Waterloo/Cambridge (Ontario), and Lompoc. These variables are included to account for fixed effects – variations in geography, conservation programs, regulations and culture not addressed by the other independent variables.

6. MODEL OF WATER DEMAND UNDER INCREASING-BLOCK AND UNIFORM MARGINAL PRICES

We use a multiplicative or log-log functional form for demand in this analysis, described

[14]While less than 10 percent of sample households have evaporative coolers, they are quite common in the arid cities. Forty-three percent of sample households in Phoenix have evaporative coolers, as do one-third of households in Tempe and Scottsdale, higher-income suburbs of Phoenix. Mean annual income among households with evaporative coolers is $56,000, compared with $71,000 among households without them. The t-statistic in a test of the significance of this difference in means is 9.94.

[15]Lawn moisture needs are captured through estimated evapotranspiration, less effective rainfall (equal to 0.6* measured precipitation). We use Hargreaves' approximation to the Penman-Monteith evapotranspiration equation, which requires mean, minimum, and maximum daily temperature, degrees latitude (to estimate a solar radiation parameter), and a readily-available constant associated with the crop of interest – green grass.

[16]The Energy Policy Act of 1992 established a national manufacturing standard of 1.6 gallons per flush for most toilets. Initial implementation occurred on January 1, 1994. All of the cities in our study have passed ordinances requiring low-flow plumbing fixtures in newly constructed and renovated residential structures, some of which were also required by state law.

by equations (1) and (2).[17]

$$w = \exp(Z\delta) p^{\alpha} \tilde{Y}^{\mu} \exp(\eta) \exp(\varepsilon) \qquad (1)$$

$$\ln w = Z\delta + \alpha \ln p + \mu \ln \tilde{Y} + \eta + \varepsilon \qquad (2)$$

The dependent variable, w, is daily household water demand, and Z comprises daily weather observations, as well as one-time observations of socio-demographic and housing characteristics. Marginal prices enter the demand equation as p, and virtual income as $\tilde{Y}$.

The model has two error terms. The first source of error is heterogeneity of preferences for water consumption among households, represented by the term η. The second source of error, thought of as optimization or perception error, ε, reflects the fact that actual use may not be coincident with intended use. We assume that $\eta \sim N(0, \sigma_{\eta}^{2})$ and $\varepsilon - N(0, \sigma_{\varepsilon}^{2})$. Given no reason to believe otherwise, we treat the two error terms as independent.

We estimate three models of daily household water demand – the two-error discrete continuous choice (DCC) model described above, and for comparison purposes, a generalized least squares (GLS) random-effects model, and an instrumental variables (IV) model.

6.1 Random effects and IV models

The GLS model, as explained earlier, generates estimates that will be biased and inconsistent because it does not account for the fact that marginal price and the choice of block are simultaneously determined. The IV model is similar to the two-stage least squares (2SLS) model estimated by Hewitt and Hanemann (1995) and originally due to Wilder and Willenborg (1975). The first stage equation is a regression of observed marginal price on the characteristics of the price structure (fixed charges and the full set of marginal prices), as well as all of the exogenous covariates. The second stage equation uses predicted values of price and the exogenous covariates to generate parameter estimates. While this model does not account for all of the theoretical and econometric issues associated with block pricing, it does account for the simultaneous determination of price and quantity, or more precisely of price and the block in which consumption occurs.

We employ instruments for marginal price that do not depend on specific block quantity cutoffs, which vary widely among price structures. We do this by creating a set of variables representing the marginal price of consuming certain quantities of water—the marginal price of

[17]For an explanation of how this demand function is derived from the underlying utility function, see Hewitt (1993).

1,000 gallons, 2,000 gallons, and so on.[18] The fixed charges and exogenous covariates complete the set of instruments.

We also account for the fact that observations are correlated across households. Our IV model is a two-stage GLS random-effects model for panel data. This is important not only because of the efficiency gain anticipated from recognizing the panel structure of the data, but because failing to do so may bias standard error estimates substantially downward if the error terms are positively correlated (Moulton 1986, 1990). This is particularly true when estimation of the coefficients of interest relies on between-group variation more than within-group variation, which is the case for the price coefficient in the sample (Moulton 1986).[19]

6.2 Discrete-continuous choice model

Parameters estimated by the IV model will ignore the possibility of block-switching in response to price and income changes. Thus, we estimate the DCC model to obtain elasticity estimates. In equation (3) below, we describe conditional demand under increasing block prices, with *K* blocks and *K-1* kinks, where w is observed consumption, $\underline{w}_i^*(Z, p_i, \tilde{Y}_i; \delta, \alpha, \mu)$ is optimal consumption in the interior of block segment *k*, and w_k is equal to consumption at kink point *k*.

[18]The method here is similar to that used by Gruber and Saez (2000) to estimate the elasticity of taxable income.

[19]Antweiler has termed our approach "mixed effects," in which the lower level of grouping (here households) is accounted for with random effects, and the higher level with fixed effects (Antweiler 2001).

$$
\ln w = \begin{cases}
\ln \underline{w}_1^*(Z, p_1, \tilde{Y}_1; \delta, \alpha, \mu) + \eta + \varepsilon \\
\qquad if \;\; -\infty < \eta < \ln w_1 - \ln \underline{w}_1^*(Z, p_1, \tilde{Y}_1; \delta, \alpha, \mu) \\
\ln w_1 + \varepsilon \\
\qquad if \;\; \ln w_1 - \ln \underline{w}_1^*(Z, p_1, \tilde{Y}_1; \delta, \alpha, \mu) < \eta < \ln w_1 - \ln \underline{w}_2^*(Z, p_2, \tilde{Y}_2; \delta, \alpha, \mu) \\
\ln \underline{w}_2^*(Z, p_2, \tilde{Y}_2; \delta, \alpha, \mu) + \eta + \varepsilon \\
\qquad if \;\; \ln w_1 - \ln \underline{w}_2^*(Z, p_2, \tilde{Y}_2; \delta, \alpha, \mu) < \eta < \ln w_2 - \ln \underline{w}_2^*(Z, p_2, \tilde{Y}_2; \delta, \alpha, \mu) \\
\ldots \\
\ln w_{\kappa-1} + \varepsilon \\
\qquad if \;\; \ln w_{\kappa-1} - \ln \underline{w}_{\kappa-1}^*(Z, p_{\kappa-1}, \tilde{Y}_{\kappa-1}; \delta, \alpha, \mu) < \eta < \ln w_{\kappa-1} - \ln \underline{w}_\kappa^*(Z, p_\kappa, \tilde{Y}_\kappa; \delta, \alpha, \mu) \\
\ln \underline{w}_\kappa^*(Z, p_\kappa, \tilde{Y}_\kappa; \delta, \alpha, \mu) + \eta + \varepsilon \\
\qquad if \;\; \ln w_{\kappa-1} - \ln \underline{w}_\kappa^*(Z, p_\kappa, \tilde{Y}_\kappa; \delta, \alpha, \mu) < \eta < \infty
\end{cases}
$$

The full derivation of the likelihood function, given below as equation (4), is described in Cavanagh (2002), Appendix 1.A.

$$\ln L = \sum_{uniform\ price\ hholds} \ln\left(\frac{1}{\sqrt{2\pi}} * \frac{\exp-(s_1)^2/2}{\sigma_v}\right) + \tag{4}$$

$$\sum_{block\ price\ hholds} \ln\left[\begin{array}{l}\sum_{k=1}^{K}\left(\frac{1}{\sqrt{2\pi}} * \frac{\exp-(s_k)^2/2}{\sigma_v}\right)*(\Phi(r_k)-\Phi(n_k)) \\ +\sum_{k=1}^{K-1}\left(\frac{1}{\sqrt{2\pi}} * \frac{\exp-(u_k)^2/2}{\sigma_\varepsilon}\right)*(\Phi(m_k)-\Phi(t_k))\end{array}\right]$$

$$Where: \begin{aligned} s_k &= (\ln w_i - \ln \underline{w}_k^*(.))\big/\sigma_v \\ u_k &= (\ln w_i - \ln w_k)\big/\sigma_\varepsilon \\ t_k &= (\ln w_k - \ln \underline{w}_k^*(.))\big/\sigma_\eta \\ r_k &= (t_k - \rho s_k)\big/\sqrt{1-\rho^2} \\ m_k &= (\ln w_k - \ln \underline{w}_{k+1}^*(.))\big/\sigma_\eta \\ n_k &= (m_{k-1} - \rho s_k)\big/\sqrt{1-\rho^2} \end{aligned}$$

The first summation refers to households served by utilities that charge a constant marginal price for water consumption – the parameters in this part of the equation if estimated in isolation would be equivalent to OLS estimates, because the budget constraint is linear. The second summation refers to households facing block prices and reflects both the discrete and the continuous choice. Within this portion of the equation, the first sub-summation represents the probability statement for consumption on *K* linear segments, and the second sub-summation represents the probability statement for consumption at the *K-1* kink points.[20]

7. EMPIRICAL RESULTS

Coefficient estimates, standard errors, and significance levels for the three models – GLS, IV and DCC – are reported in Table 2. Because of the known biases in the basic GLS model estimates, we do not directly interpret the parameter estimates, but note the large positive and statistically significant price coefficient, representing the slope of the increasing block price structures in the data, rather than a demand curve. Coefficients in the IV and DCC models cannot

[20]The DCC model does not account for the panel nature of the data. Rather, all observations are treated as if they are independent, and each makes an equally-weighted contribution to the value of the maximized likelihood function. Estimating a model that recognizes the panel nature of the data is a possible extension of this work.

be interpreted directly as elasticities, because they do not reflect the probability that a household switches blocks in response to a change in price or income. The coefficients are conditional elasticities, given consumption within a certain block. This is the best that can be done with the IV model, given that the discrete part of the consumption choice is not modeled directly. On the other hand, the DCC model reflects both the discrete and continuous parts of a household's choice, making it possible to simulate price and income elasticities with the parameter estimates. To simulate elasticities, we take expectations of the exponential form of the conditional demand function, simulate a percentage change in the explanatory variable, and then calculate the resulting change in expected demand.[21]

Table 2. Water Demand Model Estimates

	Parameter Estimates [a]		
Variable [b]	**GLS**	**IV**	**DCC**
lnprice	1.4019 (.0483)***	-.6336 (.1235)***	-.3408 (.0298)***
lnincome	.1257 (.0305)***	.1490 (.0323)***	.1305 (.0118)***
seas	.2579 (.0205)***	.3272 (.0215)***	.3070 (.0247)***
weath	.0075 (.0010)***	.0083 (.0011)***	.0079 (.0013)***
maxt	.0146 (.0015)***	.0207 (.0016)***	.0196 (.0018)***
famsz	.1694 (.0144)***	.1973 (.0153)***	.1961 (.0056)***
bthrm	.0325 (.0240)	.0533 (.0254)**	.0585 (.0093)***
sqft	.0713 (.0362)**	.1390 (.0385)***	.1257 (.0140)***
lotsz	.0041 (.0023)*	.0067 (.0025)***	.0065 (.0009)***
age	.0778 (.0565)	.0908 (.0598)	.0867 (.0219)***
age2	-.0150 (.0094)	-.0154 (.0099)	-.0137 (.0036)***
evap	.1922 (.0787)**	.2477 (.0833)***	.2277 (.0300)***

[21]This method is described in detail in Cavanagh (2002), Appendix 1.B.

Table 2. (Continued)

	Parameter Estimates [a]		
Variable [b]	**GLS**	**IV**	**DCC**
lasv	-.7523 (.0971)***	.3925 (.1206)***	.2592 (.0409)***
seat	-1.0216 (.0914)***	.0510 (.1133)	-.1231 (.0380)***
sandg	-.6722 (.0894)***	.0786 (.1032)	.0136 (.0366)
tampa	-.6098 (.0893)***	-.3406 (.0956)***	-.3881 (.0373)***
phx	-.0734 (.0927)	.0482 (.0983)	-.0043 (.0385)
tscot	.0377 (.0915)	-.1508 (.0974)	-.1032 (.0360)***
eug	.9175 (.0931)***	-.1598 (.1150)	.0050 (.0388)
wcamb	-.6669 (.0909)***	-.1416 (.1004)	-.1938 (.0362)***
wvall	-.3842 (.0921)***	.2862 (.1042)***	.1785 (.0388)***
lomp	-.9395 (.0915)***	.0627 (.1115)	-.0646 (.0373)*
constant	-3.7236 (.1539)***	-3.7198 (.1627)***	-3.6993 (.0653)***
σ_η	----	----	1.0768 (.0103)***
σ_ε	----	----	.3554 (.0277)***

(a) Standard errors are in parentheses. Asterisks denote significance levels: *** at $\alpha \leq .01$; ** at $.01 < \alpha \leq .05$; and * at $.05 < \alpha \leq .10$.
(b) All city effects are relative to households in Denver, Colorado.

7.1 Effects of price and income on water demand

Simulated price and income elasticities for the DCC model are provided in Table 3, along with the respective IV coefficient estimates for purposes of comparison .

Table 3. Effects of Changes in Price and Income on Water Demand

Change in Variable	DCC Model Simulated Elasticity	IV Coefficient Estimate
1% price increase	-0.3319	-0.6336
1% income increase	0.1273	0.1490

7.1.1 Income elasticity

The estimated income coefficient in the IV model is 0.15, and the estimated elasticity for the full sample in the DCC model is 0.13, where the simulated income elasticities are derived from changes to base income, meaning that virtual income in all blocks rises by the same absolute amount. These income elasticity estimates and those of others who have applied DCC models, using indirect measures of income, are low compared with most previous estimates. The range of income elasticity estimates from 1951-1991 was 0.18 to 2.14, with most estimates falling in the range of 0.2 to 0.6 (Hanemann 1997b). Most previous studies, however, did not include household-level information on housing characteristics, which are strongly positively correlated with income. The omission of these variables would have overestimated the influence of income on water demand.

7.1.2 Price elasticity

The simulated price elasticities in Table 3 are derived from equivalent percentage increases in prices of all blocks at once, including resulting changes in virtual income. As expected, in the IV model, the use of instruments addressed the endogeneity of price and quantity, resulting in the estimation of a downward-sloping demand curve. The price coefficient of -0.63, an elasticity conditional on households remaining within their currently observed segment of the budget constraint, is estimated at a very high level of significance and is within the range of previous estimates.[22] But unconditional elasticities cannot be obtained from the IV parameter estimates.[23]

---

[22]The IV model in Hewitt and Hanemann (1995) did not identify a significant price effect. There is a good deal more price variation in this sample, as well as a significantly greater number of households (1,082 compared to 121 in that study), so it is not surprising to find an effect where they did not.

[23]In addition, the IV model requires the assumption that the block in which a household is observed (and resulting marginal price and virtual income) is the block in which a household actually consumes. The IV model cannot account for the fact that a large draw from the distribution of the error term may be incompatible with the observed consumption block and marginal price equaling actual block and marginal price. This amounts to arbitrary

The price elasticity measured by the DCC model is -0.33, about half the magnitude of the IV estimate.

7.1.3 Differences from previous estimates

Our simulated price elasticity is quite different from that previously estimated by Hewitt and Hanemann (1995), who use a DCC model and estimate an elasticity almost six times the magnitude of ours. Their estimate might have been considered no more than an outlier, since among hundreds of studies in the literature since 1965, only three studies prior to Hewitt and Hanemann (1995) reported estimates in the elastic range (Howe and Lineweaver 1967, Danielson 1979, Deller et al. 1986). But since it was also the first study to model the piecewise-linear budget constraint fostered by block pricing, the result caused some to ask whether previous studies had underestimated elasticity through incorrect modeling.

On the surface, this may seem similar to the situation with labor supply, in which the DCC model (often called the Hausman model) has led to estimates of greater labor supply distortions from taxation than other models (Hausman 1981, Blundell and MaCurdy 1999, Heim and Meyer 2001). But we do not expect the same differential to arise in the water demand literature, for reasons stated earlier. The Hausman result is driven, in large part, by the substantial implicit lump-sum transfers that result from progressive income taxation and increase the compensated wage elasticity of labor supply. The equivalent transfers are minuscule in water pricing. Thus, the DCC model would be unlikely to estimate high price elasticities in water demand, as it has wage elasticities in labor supply. Our DCC elasticity estimates, produced by exploiting the widest variation yet available in price, price structure, and individual household characteristics, do not depart drastically from other studies.

Reasons for the difference between our estimate and that of Hewitt and Hanemann (1995) are likely related to differences in model specification and respective samples. We exclude marginal wastewater charges from the present analysis, and we include a variety of additional household-level variables. More importantly, much of our sample lies outside of the range of the price variation in Hewitt and Hanemann (1995).

7.2 Relationship of Price Structure to Demand

Another important factor that may have caused price elasticity of demand to differ in our analysis from previous studies is linked with the fact that 40 percent of households in this sample face uniform marginal prices, with the other households facing two-tier or four-tier increasing block structures. A causal link between price elasticity and price structure has been suggested (Espey et al. 1997, Hewitt 2000). We test whether or not price elasticity varies with price structure, and we examine some possible explanations for observed differences in elasticity among price structures.

---

assignment of households to locations on the budget constraint, particularly in the neighborhoods of kink points.

7.2.1 Observed differences in price elasticity

When we estimate the DCC model for households facing increasing block prices, we obtain a simulated price elasticity of approximately -0.60, whereas for households facing uniform marginal prices, we obtain an elasticity estimate of -0.19, and the effect is statistically insignificant (Table 4).

Table 4. Price and Income Elasticity Estimates by Type of Price Structure

Category of Price Structure	**Price Elasticity Estimate**	**Income Elasticity Estimate**
All sample households	-0.3319	0.1273
Uniform price households	-0.1937	0.0175
Two- & four-block households	-0.6007	0.1962

We also estimate an additional full-sample model, allowing price elasticity to vary with price structure. The demand function in this model includes both a dummy variable denoting whether a household faces block prices (1) or a uniform marginal price (0), and an interaction term between marginal price and the block price dummy.[24] The demand function with the block-pricing variables is included here in log form as equation (5), in which *D* is the block pricing dummy variable.

$$\ln w = Z\delta + \lambda D + \alpha \ln p + \gamma D \ln p + \mu \ln \tilde{Y} + \eta + \varepsilon \qquad (5)$$

[24]In the DCC model, the likelihood function for households facing block prices includes many different possible marginal prices, each of which are interacted with the block pricing dummy.

Table 5. Water Demand Model Estimates with Block Pricing Test

Variable [a]	Coefficient Estimate [b]	Standard Error
lnprice	0.1291*	0.0118
Dblocks	-0.6434*	0.1847
Dblocks*lnprice	-0.9941*	0.2290
lnincome	0.6606*	0.2300
seas	0.3077*	0.0246
weath	0.0073*	0.0013
maxt	0.0206*	0.0018
famsz	0.1952*	0.0056
bthrm	0.0586*	0.0093
sqft	0.1279*	0.0140
lotsz	0.0066*	0.0009
age	0.0927*	0.0219
age2	-0.0150*	0.0036
evap	0.2276*	0.0300
lasv	0.2436*	0.0410
seat	-0.1815*	0.0429
sandg	0.0048	0.0368
tampa	-0.3989*	0.0374
phx	-0.0157	0.0383
tscot	-0.1107*	0.0359
eug	-0.8156*	0.2476
wcamb	-0.4809*	0.0818
wvall	-0.0455	0.0717
lomp	-0.2004*	0.0630
constant	-3.0755*	0.1926
σ_η	1.0766*	0.0103
σ_ε	0.3530*	0.0280

(a) All city effects are relative to households in Denver, Colorado.
(b) Asterisks (*) mark estimates significant at $\alpha \leq .01$.

The results of this test for an effect of block pricing on water demand and price elasticity are consistent with the results we obtained when estimating separate demand functions for households facing uniform vs. block pricing. The presence of block pricing has a strong negative effect on price elasticity and on water demand, as well.[25] The effects of household characteristics and weather variables are essentially unchanged from the previous model. The city fixed effects are somewhat different, especially for cities with uniform prices. This is to be expected, however, given that we are now controlling for a source of variation (price structure) which was previously absorbed in the city fixed effects.

However, the results of this model are difficult to interpret. The estimated price coefficient for households facing uniform prices is positive. Although the model we estimate above for uniform-price households alone fails to identify a statistically significant effect of price on water demand (not an altogether uncommon result in the water demand literature), the estimated price coefficient is, at least, negative. For unknown reasons, the income coefficient we estimate in this model is much higher than that in the DCC model without the block price variables.

Modeling the precise nature of the relationship between block pricing and price elasticity is part of our ongoing research. Results thus far do seem to indicate that households facing block prices are more sensitive to price and income changes than households facing uniform prices. We are exploring explanations for this observed difference.

7.2.2 Utility-level heterogeneity and observed price elasticity variation

Although correlation does not demonstrate causality, it is at least conceivable that part of the observed difference in price elasticity of demand across price structures is indeed driven by price structures themselves, which may have the effect – in the case of the multi-tiered structures – of focusing consumer attention. Such phenomena have been observed in other contexts, such as electricity demand (Newell, Jaffe and Stavins 1999).

There are four other possible explanations that merit discussion. First, it is possible that current prices are simply higher among block-price utilities than among uniform-price utilities, but it is doubtful that prices explain the observed elasticity variation. The average price for 1,000 gallons of water faced by uniform-price households ($1.72) is higher than the average price faced by block-price households ($1.53).[26] Thus, if the magnitude of prices were responsible for different elasticities, we would expect the opposite result from that described here.

---

[25]The effect on water demand is particularly interesting, given that mean daily water consumption (without controlling for other factors) is approximately 30 percent higher for households facing block prices than for households facing uniform prices.

[26]Using a t-test for difference in means, the t-statistic=25.8.

Second, if expenditures for water comprise a larger budget share for households served by utilities with block pricing, this could help explain the larger observed price elasticity. But estimated annual costs for water consumption (including fixed charges) in our sample are approximately 0.5 percent of annual income, and there is no significant difference in budget share between households facing different price structures. Thus, observed differences in price elasticity cannot be explained by this factor.

The third possible source of heterogeneity is the extent to which households may have adopted water-efficient technologies, allowing greater responses to price changes. This may be related to greater long-term aridity, or to greater variation in prices over time. The elasticities estimated here are a mixture of short-run and long-run estimates. Most of the variation in price and income occurs in the cross section, and it is commonly assumed that cross-sectional data reflect steady-state variation (Baltagi and Griffin 1984, Taylor 1975). However, the time series is very short, and in many cities there has been little variation in prices over time. We can test for the possibility of differential technology adoption among households facing different price structures.[27]

Fourth, households facing block prices may devote a greater share of total water consumption than their counterparts in uniform-price cities to uses for which substitutes exist. In our sample, households facing block prices devote about 26 percent of total water consumption to outdoor uses, including irrigation and swimming pools. Households facing uniform prices devote only 18 percent of water consumption to outdoor uses. The fact that households in cities with block pricing use more water outdoors may be due to aridity; the simple correlation between the presence of block pricing and utility-level long-term average monthly precipitation is -0.315. If utilities in drier cities are more likely to implement block pricing, it is no surprise that a greater fraction of water consumption in these areas is devoted to outdoor uses, and we would anticipate that demand for such purposes would be particularly sensitive to price.

Economists have long suggested that outdoor water use is more price-elastic than indoor water use. Indeed, this is one reason that many utilities have implemented increasing-block rates – they hope to reduce water consumption (especially during the peak season) by increasing prices on the most price-elastic units. Further research is needed to determine whether this characteristic of increasing block prices partly explains the higher observed price elasticities for block price cities.

7.3 Effect of weather and household characteristics on water demand

The weather, household, and housing characteristics consistently have the expected signs in both the IV and DCC models, and the results of the two models with respect to these

[27]Our data do allow us to examine whether households facing block prices in the sample have, on average, adopted water-conserving technologies for indoor uses in greater proportion than households facing uniform prices.

covariates differ little (Table 6).[28] Results regarding the age of homes were weaker than anticipated, but the non-monotonic effect is as expected. The highest water demand occurs among homes in the range of 20 to 40 years old; both newer homes and older homes use less water.

Table 6. Effects of Selected Independent Variables on Water Demand

	Percent Change in Demand from Simulated Variable Change	
Independent Variable (simulated change)	**IV**	**DCC**
Season (change from non-irrigation to irrigation season)	38.74	35.10
Evapotranspiration less rainfall (add 1 mm)	0.83	0.78
Maximum daily temperature (add 1°C)	2.09	1.94
Number of household members (add one member)	21.83	21.23
Number of bathrooms (add one bathroom)	5.51	5.91
Area of home (add 1,000 ft^2)	15.00	13.13
Area of lot (add 1,000 ft^2)	0.67	0.64
Evaporative cooler (add evaporative cooler)	28.11	25.04

7.4 Between-city variation in water demand

Results for the city dummy variables represent fixed city-level variation not otherwise accounted for in the model, including attitudes toward water use and water conservation, utility conservation education programs, and water requirements of residential landscaping. Table 7 exhibits the percentage increase or decrease in daily water demand associated with residence in a given city, *ceteris paribus*, in comparison with residence in Denver, Colorado.

[28]Figures reported in Table 6 are $(\exp(\delta + \sigma_{\varepsilon}^{2} / 2) - 1) * 100)$ for the IV estimates, and simulated percent changes in demand for the DCC estimates.

Table 7. Estimated Percentage Increase or Decrease in Average Daily Water Demand Associated with Residence in Sample Cities, DCC Model

City or Utility District	Estimated Percentage Increase or Decrease in Average Daily Demand from Residence in City
Las Virgenes MWD, California	29.03
Walnut Valley WD, California	19.18
San Diego, California	1.34
Denver, Colorado	0.00
Eugene, Oregon	0.49
Phoenix, Arizona	-0.42
Lompoc, California	-6.14
Tempe/Scottsdale, Arizona	-9.63
Seattle, Washington	-11.37
Waterloo/Cambridge, Ontario	-17.30
Tampa, Florida	-31.61

8. CONCLUSIONS

In this study, we have estimated price and income elasticities of water demand among urban households facing a variety of price structures, including increasing block pricing, using a structural model that accommodates piecewise-linear budget constraints. We exploit the most price-diverse, detailed, household-level water demand data yet available to obtain elasticity estimates more representative of U.S. urban areas than those of previous studies.

We estimate substantial differences in the price elasticity of residential water demand among households facing increasing block prices versus households facing uniform marginal prices. We have offered a set of potential explanations for these differences, related to substitution possibilities of outdoor water uses, as well as potential behavioral reactions to the price structure.

Economists in the policy arena have long recommended price-based instruments for environmental and natural resource management, because of their cost-effectiveness relative to command-and-control instruments. Yet, in part due to the very low price elasticities typically estimated, water utilities rarely implement price increases in response to scarcity, instead relying primarily on technology standards and specific water use restrictions. Our research suggests that, given consumers' low elasticity at current prices, price structure may be an important and cost-effective alternative to command-and-control approaches, and possibly a more effective alternative, in terms of its ability to reduce water consumption, than increases in the magnitude of marginal price, itself.

In conclusion, we note some areas for further research. First, we continue to explore the observed difference in price elasticity and water demand across price structures to determine whether the effect is, indeed, attributable to price structure, or whether it is due to unobserved heterogeneity in utility service areas. Second, the use of non-parametric techniques to estimate price and income elasticities of water demand may be fruitful. Results using non-parametric models should be contrasted with those from the maximum likelihood model described here, employing different demand functions and distributional assumptions.

References

Agthe, Donald E., R. Bruce Billings, John L. Dobra, and Kambiz Raffiee (1986), "A Simultaneous Equation Demand Model for Block Rates," *Water Resources Research* 22(1): 1-4.

Antweiler, Werner (2001), "Nested random effects estimation in unbalanced panel data," *Journal of Econometrics* 101: 295-313.

Baltagi, Badi and James M. Griffin (1984), "Short and Long Run Effects in Pooled Models," *International Economic Review* 25(3): 631-645.

Bauman, D. D., J. J. Boland and W. M. Hanemann (1997), *Urban Water Demand Management and Planning* (McGraw-Hill, New York).

Baumol, W. J. and W. E. Oates (1988), *The Theory of Environmental Policy*, Second Edition (Cambridge University Press, New York).

Billings, R. B. (1982), "Specification of Block Rate Price Variables in Demand Models," *Land Economics* 58(3): 386-394.

Billings, B. R. and D. E. Agthe (1980), "Price Elasticities for Water: A Case of Increasing Block Rates," *Land Economics* 56(1): 73-84.

Blomquist, Sören, Matias Eklöf, and Whitney Newey (2001), "Tax reform evaluation using non-parametric methods: Sweden 1980-1991," *Journal of Public Economics* 79:543-568.

Blundell, Richard and Thomas MaCurdy (1999), "Labor Supply: A Review of Alternative Approaches," in Ashenfelter, O. and D. Card, eds., *The Handbook of Labor Economics* (North-Holland, New York): 1559-1695.

Burtless, Gary (1981), Comments on Jerry A. Hausman, "Labor Supply," in Henry J. Aaron and Joseph A. Pechman, eds., *How Taxes Affect Economic Behavior* (Brookings Institution, Washington, D.C.): 76-83.

Burtless, Gary and Jerry A. Hausman (1978), "The Effect of Taxation on Labor Supply: Evaluating the Gary Income Maintenance Experiment," *Journal of Political Economy* 86(December): 1101-1130.

Cavanagh (Olmstead), Sheila M. (2002), "Essays in Environmental Economics and Policy," Ph.D. Dissertation, Harvard University, Cambridge, Massachusetts.

Chicoine, D. L. and G. Ramamurthy (1986), "Evidence on the Specification of Price in the Study of Domestic Water Demand," *Land Economics* 62(1): 26-32.

Corral, L. R. (1997), Price and Non-Price Influence in Urban Water Conservation, Unpublished Ph.D. Dissertation, University of California at Berkeley, Berkeley, CA.

Crocker, T. D. (1966), "The Structuring of Atmospheric Pollution Control Systems," in Wolozin, H., ed., *The Economics of Air Pollution* (Norton, New York): 61-86.

Dales, J. H. (1968), *Pollution, Property and Prices* (University of Toronto Press, Toronto).

Danielson, Leon E. (1979), "An Analysis of Residential Demand for Water Using Micro Time-Series Data," *Water Resources Research* 22(4): 763-767.

Deller, Steven C., David L. Chicoine and Ganapathi Ramamurthy (1986), "Instrumental Variables Approach to Rural Water Service Demand," *Southern Economic Journal* 53(2): 333-346.

Espey, M., J. Espey and W. D. Shaw (1997), "Price elasticity of residential demand for water: A meta-analysis," *Water Resources Research* 33(6): 1369-1374.

Foster, H. S., Jr. and B. R. Beattie (1979), "Urban Residential Demand for Water in the United States," *Land Economics* 55(1): 43-58.

Gan, Li and Dale Stahl (2002), "Estimating Labor Supply under Piecewise Linear Budget Constraints with Measurement Error," Working paper, Department of Economics, University of Texas, Austin (October).

Gibbs, K. C. (1978), "Price Variable in Residential Water Demand Models," *Water Resources Research* 14(1): 15-18.

Greene, D. L. (1990), "CAFÉ or Price? An Analysis of the Effects of Federal Fuel Economy Regulations and Gasoline Price on New Car MPG, 1978-89," *The Energy Journal* 11(3): 37-57.

Gruber, Jonathan and Emmanuel Saez (2000), "The Elasticity of Taxable Income: Evidence and Implications," NBER Working Paper 7512 (National Bureau of Economic Research, Cambridge, MA).

Hall, Robert E. (1973), "Wages, Income and Hours of Work in the U.S. Labor Force," in Glen G. Cain and Harold W. Watts, eds., *Income Maintenance and Labor Supply* (Rand McNally College Publishing Company, Chicago): 102-159.

Hanemann, W.M. (1997a), "Determinants of Urban Water Use", in Baumann, D.D., J.J. Boland, and W.M. Hanemann, eds., *Urban Water Demand Management and Planning* (McGraw-Hill, New York): 31-76.

Hanemann, W.M. (1997b), "Price Rate and Structures," in Baumann, D.D., J.J. Boland and W.M.

Hanemann, eds., *Urban Water Demand Management and Planning* (McGraw-Hill, New York): 137-179.

Hausman, Jerry A. (1981), "Labor Supply," in Henry J. Aaron and Joseph A. Pechman, eds., *How Taxes Affect Economic Behavior* (Brookings Institution, Washington, D.C.): 27-83.

Hausman, Jerry A. (1985), "Taxes and Labor Supply," in Auerbach, A. and M. Feldstein, eds., *The Handbook of Public Economics* (North-Holland, Amsterdam): 213-263.

Hausman, J. A., M. Kinnucan, and D. McFadden (1979), "A Two-level Electricity Demand Model: Evaluation of the Connecticut Time-of-Day Pricing Test," *Journal of Econometrics* 10:263-289.

Heckman, James J. (1983), "Comment on Jerry A. Hausman's 'Stochastic Problems in the Simulation of Labor Supply'" in Martin Feldstein, ed., *Behavioral Simulation Methods in Tax Policy Analysis* (University of Chicago Press, Chicago, IL): 70-82.

Heim, Bradley T. and Bruce D. Meyer (2001), "Structural Labor Supply Models when Budget Constraints are Nonlinear," Working paper presented at National Bureau of Economic Research, November 2, 2001.

Hewitt, Julie A. (2000), "An Investigation into the Reasons Why Water Utilities Choose Particular Residential Rate Structures," in Ariel Dinar, ed., *The Political Economy of Water Pricing Reforms* (New York: Oxford University Press): 259-277.

Hewitt, Julie A. (1993), "Watering households: The two-error discrete-continuous choice model of residential water demand," Unpublished Ph.D. Dissertation, University of California, Berkeley, CA.

Hewitt, Julie A. and W. Michael Hanemann (1995), "A Discrete/Continuous Choice Approach to Residential Water Demand under Block Rate Pricing," *Land Economics* 71(2): 173-192.

Hirst, E. and R. Goeltz (1984), "The Economics of Utility Residential Energy Conservation Programs: A Pacific Northwest Example," *The Energy Journal* 5(3): 159-169.

Howe, C. W. and F. P. Lineweaver (1967), "The Impact of Price on Residential Water Demand and its Relation to System Demand and Price Structure," *Water Resources Research* 3(1): 13-32.

Jaffe, A. B. and R. N. Stavins (1994a), "The Energy Paradox and the Diffusion of Conservation Technology," *Resource and Energy Economics* 16(2): 91-122.

Jaffe, A. B. and R. N. Stavins (1994b), "Energy-Efficient Investments and Public Policy," *The Energy Journal* 15(2): 43-65.

Kiefer, J. C., B. Dziegielewski and E. M. Opitz (1993), Analysis of Water Savings from the LITEBILL Program: An Evaluation of Alternative Research Methods (Planning and Management Consultants, Ltd., Carbondale, IL).

Mayer, P. W., W. B. DeOreo, E. M. Opitz, J. C. Kiefer, W. Y. Davis, B. Dziegielewski and J. O. Nelson (1998), *Residential End Uses of Water* (American Water Works Association Research Foundation, Denver).

Michelsen, A. M., J. T. McGuckin and D. M. Stumpf (1998), *Effectiveness of Residential Water Conservation Price and Nonprice Programs* (American Water Works Association Research Foundation, Denver, CO).

Moffitt, Robert (1990), "The Econometrics of Kinked Budget Constraints," *Journal of Economic Perspectives* 4(2): 119-139.

Moffitt, Robert (1986), "The Econometrics of Piecewise-Linear Budget Constraints: A Survey and Exposition of the Maximum Likelihood Method," *Journal of Business and Economic Statistics* 4(3): 317-328.

Montgomery, W. D. (1972), "Markets in Licenses and Efficient Pollution Control Programs," *Journal of Economic Theory* 5(3): 395-418.

Moulton, B. R. (1990), "An Illustration of a Pitfall in Estimating the Effects of Aggregate Variables on Micro Units," *Review of Economics and Statistics* 72(2):334-338.

Moulton, B. R. (1986), "Random Group Effects and the Precision of Regression Estimates," *Journal of Econometrics* 32(3): 385-397.

Nauges, Céline and Richard Blundell (2001), "Estimating Residential Water Demand Under Block Rate Pricing: A Nonparametric Approach," Working paper (November).

Newell, Richard G., Adam B. Jaffe, and Robert N. Stavins (1999), "The Induced Innovation Hypothesis and Energy-Saving Technological Change," *Quarterly Journal of Economics* 114(3): 941-975.

Nieswiadomy, Michael L. and David J. Molina (1991), "A Note on Price Perception in Water Demand Models," *Land Economics* 67(3): 352-359.

Nieswiadomy, Michael L. and David J. Molina (1989), "Comparing Residential Water Demand Estimates under Decreasing and Increasing Block Rates Using Household Data," *Land Economics* 65(3): 280-289.

Nieswiadomy, Michael L. and David J. Molina (1988), "Urban Water Demand Estimates under

Increasing Block Rates," *Growth and Change* 19(1): 1-12.

Nordin, J. A. (1976), "A Proposed Modification of Taylor's Demand Analysis: Comment," *The Bell Journal of Economics* 7(2): 719-21.

Pigou, A.C. (1920), *The Economics of Welfare* (Macmillan and Co., London).

Pint, E. M. (1999), "Household Responses to Increased Water Rates During the California Drought," *Land Economics* 75(2): 246-266.

Renwick, Mary E. and Richard D. Green (2000), "Do Residential Water Demand Side Management Policies Measure Up? An Analysis of Eight California Water Agencies," *Journal of Environmental Economics and Management* 40(1): 37-55.

Schipper, L. (1979), "Another Look at Energy Conservation," *American Economic Review* 69(2): 362-368.

Schultz, M. T., S. M. Cavanagh, B. Gu and D. J. Eaton (1997), The Consequences of Water Consumption Restrictions During the Corpus Christi Drought of 1996, Draft Report (LBJ School of Public Affairs, University of Texas at Austin).

Shin, Jeong-Shik (1985), "Perception of Price When Price Information is Costly: Evidence from Residential Electricity Demand," *Review of Economics and Statistics* 67(4): 591-598.

Stelzer, I. M. (1982), "What Role Can Utilities Play in Energy Conservation Programs?" *The Energy Journal* 3(1): 90-94.

Taylor, L. D. (1975), "The Demand for Electricity: A Survey," *The Bell Journal of Economics* 6(1): 74-110.

Wilder, Ronald P. and John F. Willenborg (1975), "Residential Demand for Electricity: A Consumer Panel Approach," *Southern Economic Journal* 42(October): 212-217.

Wirl, F. (1997), *The economics of conservation programs* (Kluwer Academic, London).

Wirl, F. and W. Orasch (1998), "Analysis of United States' Utility Conservation Programs," *Review of Industrial Organization* 13(4): 467-486.

Young, C. E., K. R. Kinsley and W. E. Sharpe (1983), "Impact on Residential Water Consumption of an Increasing Rate Structure," *Water Resources Bulletin* 19(1): 81-86.