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Sheila Olmstead, Michael Hanemann, and Robert Stavins

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by

Sheila M. Olmstead
Yale University - School of Forestry and Environmental Studies
W. Michael Hanemann
University of California, Berkeley - The Richard & Rhoda Goldman School of Public Policy
Robert N. Stavins
Harvard University - John F. Kennedy School of Government; Resources for the Future

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

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

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

Please Contact:
Prof. Sheila M. Olmstead
School of Forestry and Environmental Studies
Yale University
230 Prospect Street
New Haven, Connecticut 06511
Phone: 203-432-6247
Fax: 203-432-0026
E-Mail: sheila.olmstead@yale.edu

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ABSTRACT

We perform an empirical analysis of the influence of price and price structure on residential water demand, using the most price-diverse, detailed, household-level water demand data yet available for this purpose. Ours is the first analysis to address both the simultaneous determination of marginal price and water demand under block pricing, and the possibility of endogenous price structures in the cross section. We find that households facing increasing block prices may be more sensitive to price increases than households facing uniform marginal prices. Tests for endogenous price structures cannot rule out a behavioral response to the shape of supply, but suggest that observed differences in price elasticity under supply curves of varying shapes may result, in part, from underlying heterogeneity among utility service areas.

Keywords: water demand, price elasticity, non-linear pricing

JEL Codes: Q21, Q25, Q28, L95, D12

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1. INTRODUCTION

Water utility managers traditionally have maintained that consumers do not respond to price signals, and so demand management has occurred most frequently through restrictions on specific water uses and requirements for the adoption of specific technologies. In theory, raising prices to bring about a given level of water conservation is less costly than implementing a command-and-control approach, even if the prices in question are inefficient. Water utilities increasingly hear this message and respond, in many cases by implementing increasing block prices (IBPs), under which marginal prices increase with the quantity consumed. The price elasticity of water demand is a key policy variable of interest in this context, because: (1) regulators may use price to manage demand during periods of scarcity; and (2) utilities often face zero-profit constraints, so the impact of price changes on total revenues is a concern.

In 2000, approximately one-third of U.S. urban residential water customers faced increasing block prices, up from four percent of customers in 1982 (OECD 1999; Raftelis 1999). Consumer responses to IBPs in the market for water are unclear, at best. Estimating water demand functions under non-linear price regimes requires econometric techniques that can separate demand from supply. Under IBPs, this problem is twofold. First, marginal prices rise as consumption increases. Thus, even though price schedules are set by regulators, the upward-sloping nature of IBPs means that the price and choice of "block" in which to consume are simultaneously determined. Second, when estimating water demand functions using data in which most of the price variation occurs in the cross-section (water price schedules vary little over time), a utility's choice of price schedule, itself, may be endogenous, potentially biasing cross-sectional elasticity estimates and casting doubt on cross-city comparisons.

Ours is the first analysis to address both the simultaneous determination of marginal price and water demand under block pricing, and the possibility of endogenous price structures. Some previous studies tackled the first problem but not the second, generating results that suggest that price elasticity is higher under IBPs than under a linear marginal price. Thus, a one percent price increase under IBPs would result in a greater reduction in water consumption than a one percent price increase under a uniform marginal price, all else equal. If this is true, it implies that consumers exhibit a demand response to the shape of the supply curve, and not simply its height (the magnitude of marginal price).

Using the most price-diverse, detailed, household-level water demand data yet available, we exploit substantial cross-sectional and time-series variation in the price incentives faced by households to estimate price elasticities under heterogeneous price structures. We estimate a full-sample price elasticity of water demand of approximately -0.33, and we then examine how the price elasticity of demand varies between linear and non-linear pricing regimes.[1]

Models that split our sample by price structure suggest that demand is more price-elastic under IBPs, implying either a demand response to price structure, or underlying heterogeneity among the two sub-samples. We explore the possibility of underlying heterogeneity, with mixed results, but we cannot rule out a behavioral response to the shape of supply curves.

[1]Our elasticity estimates are a mixture of short-run and long-run. Much of the variation in price comes from the cross-section. But there is also substantial variation within individual households over time, due to IBPs. There is also a small amount of variation in price over time due to utility price changes.

The questions we examine bring to mind an earlier debate in the labor economics literature with respect to the Hausman model of labor supply under progressive income taxation. We apply the Hausman model and subsequent modifications to the estimation of water demand functions. However, we explain why the use of this model should not result in dramatic changes in how economists think about the price elasticity of water demand, as it did in how economists think about the wage elasticity of labor supply. The contrast with applications from tax policy is relevant to IBPs for commodities other than water, and must be considered carefully as analysts import the Hausman model in other markets.

The paper is organized as follows. In Section 2, we review the literature on block pricing, the price elasticity of residential water demand, and the economic theory and econometrics of piecewise linear budget constraints. In Section 3 we describe the data, and in Section 4 we describe our models. In Section 5, we describe the results, and in Section 6, we conclude.

2. THE LITERATURE AND THE ECONOMIC CONTEXT

2.1 Block Pricing and Efficiency

Urban residential water service pricing typically takes one of three forms: (1) a uniform marginal price; (2) increasing block prices; or (3) decreasing block prices. Each of these price structures is typically accompanied by a fixed water service fee. Under constant or uniform rates, households pay 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.

Water prices in North America generally lie below the long-run marginal cost (LRMC) of water supply, the efficient price (Hanemann 1997, Timmins 2003). Utilities' adoption of increasing block prices may expand the fraction of consumption that is priced at the LRMC. The LRMC can be greater than short-run average cost, because LRMC reflects the cost of new supply acquisition, and new supplies are typically more costly to develop than current supplies (Hanemann 1997). Thus, pricing all units at LRMC may cause utility revenues to exceed expenses, sometimes substantially (Moncur and Pollack 1988; Hall 2000). The efficient way to deal with this problem would be to rebate net revenues in some lump-sum fashion. IBPs, instead, charge something approaching LRMC for the "marginal" uses (lawn-watering and the like), while meeting zero-profit constraints through the manipulation of block cutoffs and infra-marginal prices. Increasing the fraction of consumption that faces efficient prices would result in welfare gains. But if the highest-tier price in an IBP schedule does reflect LRMC, some welfare losses must result from infra-marginal subsidies. Thus, the relative efficiency of uniform and IBP structures depends on a variety of factors.

Water utilities in North America are regulated monopolies, some public and some private. IBPs have been described as a possible reflection of monopoly price discrimination (Hewitt 2000). However, since water prices are usually set by elected or appointed regulatory entities, such as city councils and public utility commissions, consumers influence the adoption of particular price structures, and even the magnitude of marginal price, through the political process.[2] We do not directly model the avenues through which water prices are established. Instead, we focus on the

[2]For example, in the City of Corpus Christi, Texas, a local taxpayers' organization successfully lobbied for a regulation limiting water price increases by the public utility to the rate of inflation during the early 1990s, despite a series of severe regional droughts.

demand side – whether consumers react differently to different price structures, controlling to the extent possible for potential endogenous price structure choice (either by utilities or by collective action on the part of consumers).

2.2 Consumer responses to water prices and price structure

Previous analyses suggest that water demand is inelastic. 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, with 90 percent of all estimates between 0 and -0.75.[3]

Many of the largest price elasticity estimates — some in the range of -1 to -2 — have been obtained with structural analyses that model directly the piecewise-linear budget constraint created by block pricing (Hewitt and Hanemann 1995; Rietveld et al. 1997; Pint 1999), drawing upon an approach developed to estimate labor supply effects of progressive income taxation.[4] In addition, two meta-analyses have found that consumers facing IBPs are more price-elastic than consumers facing linear marginal prices (Espey et al. 1997; Dalhuisen et al. 2003). In comparing elasticities across different studies, the meta-analyses face many confounding factors that differentiate the studies in their samples. They control for some, but not all, of these factors. These results have caused some conjecture that previous studies had underestimated price elasticity through incorrect modeling, or that IBPs themselves are responsible for greater sensitivity of demand to price (Dalhuisen et al. 2003).

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[3]Dalhuisen et al. (2003) obtain a mean price elasticity of -0.41 in a meta-analysis of almost 300 price elasticity studies, 1963-1998.

[4] Prior to Hewitt and Hanemann (1995), only three published studies had estimated price elasticities in the elastic range (Howe and Lineweaver 1967; Danielson 1979; Deller et al. 1986).

In water demand estimation, higher elasticity estimates are associated with long-run analyses, data restricted to summer observations, and areas with increasing block rates (Hanemann 1997; Espey et al. 1997). Greater elasticities for summer use (which includes a higher irrigation component) and long-run analyses are reasonable on the basis of economic theory; but higher elasticities for areas facing IBPs are more difficult to understand. If consumers react to price structure, as well as the magnitude of marginal price, there may be a behavioral explanation (Liebman and Zeckhauser 2004).

An alternative explanation is unobserved heterogeneity among cities implementing different price structures. Any utility service area heterogeneity that is systematically correlated with price structure choice might explain the higher observed demand elasticities in IBP cities.

2.3 Block Pricing and Piecewise Linear Budget Constraints

2.3.1 Theory of piecewise linear budget constraints

Under block pricing, consumers face a piecewise-linear budget constraint. Figure 2 depicts the budget constraint in a two-tier increasing block price system with a set of hypothetical indifference curves. 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.[5]

2.3.2 Econometric Modeling of Demand Under Block Pricing

[5]Under decreasing-block structures, budget sets are non-convex, enabling multiple tangencies between consumer indifference curves and the budget constraint. 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); in the literature on labor supply and taxation (Burtless and Hausman 1978; Hausman 1985); and for the case of water supply (Hewitt 1993; Hewitt and Hanemann 1995).

The use of IBPs creates a wedge between average and marginal price and makes these prices endogenous to the consumer's water use decision. The water demand literature has focused mainly on the former; the labor supply literature handled both aspects within the framework of the discrete/continuous choice model of Hausman. The econometric complications caused by these features have been discussed in the literature (Burtless and Hausman 1978; Hanemann 1984; Moffitt 1986, 1990).

Regarding the wedge between average and marginal price, for households consuming anywhere above the first linear segment on a piecewise linear budget constraint, the marginal price is not the price paid for every unit consumed. The implicit subsidy from the infra-marginal rates can be accounted for by adding to income the difference between what a household would pay if all units were charged at the marginal price, and what they actually pay. This income supplement for a household consuming in the second block of a two-tier price structure would be equal to the shaded region in Figure 1, commonly referred to as "virtual income".[6]

Regarding the simultaneous determination of block and consumption, 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. Thus, ordinary least squares (OLS) estimates of the parameters of the demand function will be biased and inconsistent.

This problem has been addressed by estimating instrumental variables (IV) models, such as two-stage least squares, for both labor supply and water and energy demand (Hausman et al. 1979; Agthe et al. 1986; Deller et al. 1986; Nieswiadomy and Molina 1988, 1989). While such models

[6]The treatment of infra-marginal price changes as lump-sum changes in household income was developed in the labor supply literature (Hall 1973; Burtless and Hausman 1978), and in the electricity demand literature (Nordin 1976; Taylor 1975). The first application to water demand was by Billings and Agthe (1980).

can estimate downward-sloping demand curves, they exhibit three key limitations. First, simultaneous equations models can be used only to estimate elasticities conditional on remaining within the observed block of consumption. Second, the IV methods disregard the fact that some households consume within the neighborhood of a kink point, where it is unclear which value of marginal price should be assigned (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, cannot be reconciled with utility theory. Third, the effect of changes in the price structure, such as a change in block quantity thresholds, cannot be assessed, as these elements of the price structure do not enter IV models.

The discrete/continuous choice (DCC) model, a maximum likelihood model, addresses these issues. The DCC model is a stochastic expression in which each observation is treated as if it could have occurred at any kink or along any linear portion of the household's budget constraint. The probability statement for an individual observation is the 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 nature of the supply curve faced by each household.

2.4 Piecewise Linear Budgets: Water Demand vs. Labor Supply

The structural DCC model was developed in the study of labor supply by Hausman (Burtless and Hausman 1978; Hausman 1985) and generalized by Moffitt (1986, 1990). The theoretically similar context implied by discrete/continuous consumption problems was developed by Hanemann (1984), and first applied to water demand by Hewitt (1993).

Labor supply under progressive income taxation and consumer demand under IBPs share the theoretical structure implied by piecewise linear budget constraints but differ in important ways. First, compensated wage effects on labor supply are positive, but compensated price effects on demand are negative. This is an important distinction, because income effects are negative in both cases. When water prices increase, water is relatively more expensive (causing substitution toward other goods) and real income decreases, so consumers purchase less for both reasons. When wages increase, leisure is relatively more expensive (causing substitution toward labor), but real income increases (decreasing labor and increasing leisure). Applications of Hausman's model to labor supply suggested that the uncompensated effect of a tax increase was, in fact, close to zero because the income and substitution effects of a tax increase could effectively cancel each other out in the data (Hausman 1981). In the case of water demand, one would expect the compensated price effect to be *smaller* than the uncompensated effect – not larger.

In addition, Hausman's result is attributable in part to virtual income – the implicit subsidies from infra-marginal tax rates are relatively large in labor supply analysis (Burtless 1981). These implicit subsidies are very small in water demand, however, given low water prices, small differences between price "tiers" in IBP structures, and the fact that annual water expenditures, on the whole, are a very small fraction of annual income to begin with.[7] Without the substantial "compensation" through virtual income from the progressive tax rate analysis, and without competing income and substitution effects, one would not expect price elasticities of water demand

[7]In fact, attempts to measure effects of these small infra-marginal rate subsidies have been unsuccessful in water and electricity demand (unlike in labor supply), most likely for these reasons (Hausman et al. 1979, Deller et al. 1986, Nieswiadomy and Molina 1989).

to be larger, on average, where DCC models are employed to estimate them. Yet this is, in fact, what has previously been found.

Unobserved household budget constraints are a key concern in the labor context; in water demand, the price structure is administratively determined and there is no such uncertainty. Another critique of the Hausman model is that it artificially constrains wage effects to be positive and income effects to be negative (MaCurdy et al. 1990, Blundell and MaCurdy 1999), although some disagree with this contention (Blomquist 1995). The possibility of non-convex preferences due to work costs and other factors make this a serious concern in labor supply analysis (Heim and Meyer 2004), but non-convex preferences in the consumer demand context would be highly unlikely. While the application of Hausman-type methods to labor supply analysis continues to be somewhat controversial, they are very well-suited to discrete/continuous demand modeling, and we use them here.[8]

3. DEMAND MODELS

Building on Olmstead (2006), we use a log-log demand function in all of the models in this analysis, described by equation (1).[9]

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

[8]This has not prevented experimentation with nonparametric approaches (Nauges and Blundell 2005).

[9]This functional form is extremely common in water demand analysis. We use it here in order to be able to compare our estimates with Olmstead (2006), as well as Hewitt and Hanemann (1995) and Rietveld (1997). One reason for the prevalence of this functional form is the typically extreme right-skewness of water demand.

The dependent variable, w, is daily household water demand. Marginal prices enter the demand equation as p, and virtual income as $\tilde{Y}$. Water demand, with the exception of the small fraction used for drinking, is derived demand in which the primary demand is for water-consuming goods and services, such as clean laundry, indoor bathroom use, and green lawns. Thus, in addition to price and income, the demand models also include household characteristics. These variables, as well as daily weather observations and city fixed effects, are included in the matrix Z. The model has two error terms. The first source of error (η) represents heterogeneous water consumption preferences among households not explained by the household characteristics in Z. The second source of error (ε) reflects random error unobservable to both the household and the analyst. From the household's perspective, actual use may not coincide with optimal use due to leaks, running toilets, or other causes. From the analyst's perspective, ε captures these optimization errors, as well as the usual measurement error. We assume that $\eta \sim N(0,\sigma_\eta^2)$ and $\varepsilon \sim N(0,\sigma_\varepsilon^2)$, and that the two errors are independent.

Equation (2) describes conditional demand under increasing block prices, with K blocks and K-1 kinks, where w is observed consumption, $\underline{w}_k^*(Z, p_k, \tilde{Y}_k; \delta, \alpha, \mu)$ is optimal consumption on block segment k, and w_k is equal to consumption at kink point k. Equation (3) is the DCC likelihood function implied by (2) and our distributional assumptions regarding the errors, in which Φ is the standard normal cumulative distribution function.[10] We apply the DCC model developed in

[10]The derivation of the IBP portion of the likelihood function is described in Olmstead (2006).

Olmstead (2006), but add an additional term to (3), the first summation, to account for households facing uniform marginal prices.

$$\ln w = \begin{cases} \ln \underline{w}_1^*(Z, p_1, \tilde{Y}_1; \delta, \alpha, \mu) + \eta + \varepsilon \\ \qquad \text{if } -\infty < \eta < \ln w_1 - \ln \underline{w}_1^*(Z, p_1, \tilde{Y}_1; \delta, \alpha, \mu) \\ \ln w_1 + \varepsilon \\ \qquad \text{if } \ln w_1 - \ln \underline{w}_1^*(Z, p_1, \tilde{Y}_1; \delta, \alpha, \mu) < \eta < \ln w_1 - \ln \underline{w}_2^*(Z, p_2, \tilde{Y}_2; \delta, \alpha, \mu) \\ \ln \underline{w}_2^*(Z, p_2, \tilde{Y}_2; \delta, \alpha, \mu) + \eta + \varepsilon \\ \qquad \text{if } \ln w_1 - \ln \underline{w}_2^*(Z, p_2, \tilde{Y}_2; \delta, \alpha, \mu) < \eta < \ln w_2 - \ln \underline{w}_2^*(Z, p_2, \tilde{Y}_2; \delta, \alpha, \mu) \\ \ldots \\ \ln w_{K-1} + \varepsilon \\ \qquad \text{if } \ln w_{K-1} - \ln \underline{w}_{K-1}^*(Z, p_{K-1}, \tilde{Y}_{K-1}; \delta, \alpha, \mu) < \eta < \ln w_{K-1} - \ln \underline{w}_K^*(Z, p_K, \tilde{Y}_K; \delta, \alpha, \mu) \\ \ln \underline{w}_K^*(Z, p_K, \tilde{Y}_K; \delta, \alpha, \mu) + \eta + \varepsilon \\ \qquad \text{if } \ln w_{K-1} - \ln \underline{w}_K^*(Z, p_K, \tilde{Y}_K; \delta, \alpha, \mu) < \eta < \infty \end{cases} \quad (2)$$

$$
\ln L = \sum_{uniform\ price\ households} \ln\left(\frac{1}{\sqrt{2\pi}} * \frac{\exp-(s_1)^2/2}{\sigma_\nu} \right) + \tag{3}
$$

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

$$
Where: \quad \begin{array}{l} \nu = \eta + \varepsilon \\ \rho = corr(\nu, \eta) \\ s_k = (\ln w_i - \ln \underline{w}_k^*(.)) \Big/ \sigma_\nu \\ u_k = (\ln w_i - \ln w_k) \Big/ \sigma_\varepsilon \\ t_k = (\ln w_k - \ln \underline{w}_k^*(.)) \Big/ \sigma_\eta \\ r_k = (t_k - \rho s_k) \Big/ \sqrt{1-\rho^2} \\ m_k = (\ln w_k - \ln \underline{w}_{k+1}^*(.)) \Big/ \sigma_\eta \\ n_k = (m_{k-1} - \rho s_k) \Big/ \sqrt{1-\rho^2} \end{array}
$$

4. THE DATA

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

Daily household water demand and weather conditions are observed over two periods of two weeks each, one in an arid season and one in a wet season. Daily demand data were gathered by automated data loggers, attached to magnetic household water meters by utility staff.[12]

Each household faces one price structure throughout each season of observation, but some price structures changed between the two periods. As a result, the data include 26 price structures: eight two-tier IBP structures, ten four-tier IBP structures, and eight uniform structures. Marginal prices exhibit substantial variation, from \$0 to \$4.96 per thousand gallons (kgal), with a mean observed marginal price of \$1.71/kgal.

Household demographic data were collected by a one-time survey.[13] These include lot size, square footage of homes, number of bathrooms, and family size. Home age is also included in the water demand equation, and we expect that both very old and new homes may use less water than

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[11]The household-level data were gathered by Mayer et al. (1998) for a study sponsored by the American Waterworks Association Research Foundation. Sample utilities include Denver Water (CO); Eugene Water and Electric Board (OR); Seattle Public Utilities (WA); City of Bellevue Utilities Department (WA); Highline Water District (WA); Northshore Utility District (WA); City of San Diego Water Department (CA); Tampa Water Department (FL); Phoenix Water Department (AZ); City of Tempe Water Utilities (AZ); City of Scottsdale Water Department (AZ); Regional Water Services, Municipality of Waterloo (ONT); City of Cambridge (ONT); Walnut Valley Water District (CA); City of Lompoc Utilities Department (CA); and Las Virgenes Municipal Water District (CA).

[12]These devices were hidden from sight during most, if not all, uses of water.

[13]Households chosen for the study were randomly sampled from a subset of utilities' customer databases: residential single-family households. Sampling procedures, response rates, and statistical tests for selection bias are described in Mayer et al. (1998) and Cavanagh (Olmstead) (2002).

“middle aged” homes.[14] We include one idiosyncratic housing variable in the models—the presence or absence of an evaporative cooler, which essentially substitutes water for electricity in air conditioning.[15]

We include a set of daily weather variables. Maximum daily temperature is represented by *maxt*, *weath* represents the moisture requirements of green lawns not met by precipitation, and *seas* is a dummy variable set equal to one during the arid (peak outdoor watering) season.[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), Walnut Valley, Las Virgenes, and Lompoc. These city fixed effects account for variations in geography, regulations and other factors not addressed by the other independent variables.

5. EMPIRICAL RESULTS

The full-sample DCC model, with the likelihood function in (3), is our baseline model. We explore the possibility of variation in demand and price elasticity with the shape of supply by estimating two further demand models: the baseline DCC model, but for IBP households only; and

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[14]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 newest homes were built after the passage of local ordinances in the 1980s and 1990s requiring various water-conserving fixtures.

[15]In our sample, households with evaporative coolers use on average 40 percent more water per day. We include this variable to avoid biasing the other parameter estimates, particularly the income coefficient estimate, since evaporative cooling is relatively more common in low-income households.

[16]Lawn moisture needs are captured through estimated evapotranspiration (ET), 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. For information on our ET calculation, see Cavanagh (Olmstead) (2002), Appendix C.

a panel random-effects model for UP households only. We then perform some additional analysis regarding the potential for endogenous price structures.

5.1 Full-Sample Discrete/continuous Choice Model

Results of the full-sample DCC model are reported in Table 2. Due to the non-linearity of the DCC model, parameter estimates are not marginal effects (though they are very close). To simulate elasticities, we take expectations of the exponential form of the conditional demand function, simulate a percentage change in the explanatory variable, and calculate the resulting change in expected demand. We obtain standard errors for simulated elasticities through Monte Carlo simulation: we draw 100 random samples of size N, with replacement, from our data; estimate the DCC model parameters for each Monte Carlo sample; use the parameter estimates and Monte Carlo data to simulate elasticities; and, finally, take the sample standard deviations of these 100 replications. Price and income elasticities and these bootstrapped standard errors for the full-sample model are reported in the first row of Table 3.[17]

Weather, income and the other household-level characteristics affect demand in expected ways. Results regarding the age of homes are weak, but the effect is as expected: the highest water demand occurs among homes aged 20 to 40 years; both newer homes and older homes use less water. Our income elasticity estimate (0.13) is low compared with previous studies. Most income elasticity estimates from 1951 to 1991 were in the range of 0.2 to 0.6 (Hanemann 1997). Few previous studies, however, have included household-level information on housing characteristics, which are strongly correlated with income. The omission of these variables may have overestimated the influence of income on water demand.

[17]For a more detailed treatment of these simulations, see Olmstead (2006).

The variable of greatest interest, price elasticity, indicates that a one percent price increase will reduce demand by 0.33 percent, all else equal.[18] Our simulated price elasticity, produced by exploiting the widest variation yet available in price, price structure, and individual household characteristics, is comparable to the central tendency of estimates over the past 40 years, but much smaller than those previously estimated using the DCC approach. Hewitt and Hanemann (1995), for example, obtain an elasticity almost six times the magnitude of ours.

5.2 Models Exploring Elasticity Magnitude and Price Structure

When we estimate the DCC model for households facing IBPs only, we obtain a simulated price elasticity of approximately -0.64. When we estimate a panel random-effects model for households facing uniform marginal prices only, we obtain an elasticity estimate of -0.33, and the effect is statistically insignificant, even with about 400 households in this group. This pattern reflects what we see in the existing literature, even though the 95 percent confidence intervals for these two estimates overlap, as the UP estimate is very noisy. In comparisons of studies using different samples, we observe greater price elasticities among block-price samples than among uniform-price samples. Does price elasticity vary with price structure, all else equal, or is something else going on?

5.2.1 Tests of simple theoretical explanations

To examine the question of a causal link between price elasticity and the shape of supply, we return to basics and investigate whether the observed difference can be explained by theory. First, we would expect price elasticities to be larger in the long run than in the short run. While it is possible that this explains the differences seen in the previous literature, it does not

[18]Price elasticity in our DCC models is a percent change in quantity demanded for a one percent increase in all prices in a household's price structure.

explain the difference we see in our data, since all households (IBP and UP) were observed over approximately the same time period.

Second, if the average magnitude of prices were higher in IBP cities than in uniform-price cities, this could explain the observed difference in the magnitude of price elasticity. In our sample, the average magnitude of marginal price in uniform-price cities is $1.72 per kgal, while in block-price cities it is $1.78 (assuming that marginal price for IBP households is price in the block of observed consumption). With a difference this small, it is unlikely that heterogeneity in the average magnitude of prices is responsible for the difference in elasticities we observe.

Third, if the percentage of total income devoted to water consumption were lower among households in uniform price cities, on average, than among households facing IBPs, this might explain the observed difference in elasticities. Estimated annual expenditures on water consumption (including fixed charges) in our sample are approximately 0.5 percent of annual income; we observe no statistically significant difference in the budget share for water across price structures.

We have eliminated some likely candidates from the list of potential explanations for variation in price elasticity with the shape of supply. There are two remaining possibilities. First, it is conceivable that part of the observed difference in price elasticity of demand across price structures is, indeed, driven by the shape of the supply curves themselves. But correlation does not demonstrate causality. The second possibility is that price structures, themselves, are endogenous.

5.2.2 Informal tests for endogenous price structures

Long-term aridity and long growing seasons may lead utilities to adopt IBPs, seen as a tool to reduce peak demand (chiefly summer lawn watering), and may also cause consumers to be more responsive to water price changes. IBP cities may also have implemented more extensive residential non-price conservation programs, which may increase consumers' attention to water consumption and price.

Ideally, to test for the possibility of endogenous price structures, we would estimate a two-stage model, first predicting price structure at the utility level, then estimating elasticity, given the predicted price structure. With the appropriate choice of instruments, this would presumably give us unbiased estimates of price elasticity, free of the effects of underlying heterogeneity in "supply curve design" among utilities. However, the data include only 16 water utilities, a sample too small to support the first stage of such an analysis. Even with a sufficient number of utilities, given a set of predicted types of price structure (uniform vs. IBP), how would we construct the actual schedules of prices? For example, if the first-stage model were to predict block pricing in Eugene, Oregon, when Eugene actually had implemented a uniform marginal price, the choice of number of blocks, consumption cutoffs marking block divisions, and price levels at each block would involve too many parameters for which to instrument convincingly and would, in the end, be arbitrary.

A second-best approach would be to test our full-sample models, allowing the price elasticity parameter to vary with the shape of supply, controlling perfectly for the utility-level characteristics we believe to be correlated with the choice of price structure and with price elasticity. It is unlikely, however, that we could specify a model that controls for all of these omitted variables and leaves no endogenous variation in price structure choice remaining. In addition, while our data are the best known to us for this purpose, there is not enough price variation within the uniform price cities, alone, to tolerate the necessary loss of degrees of freedom that would accompany this approach and jointly estimate a pair of price elasticities.[19]

Thus, we take a third-best approach. We estimate a new full-sample DCC demand model, including a larger set of fixed effects for the uniform price cities. This larger set of fixed effects (13 of them) absorbs all of the price variation from the uniform price utilities over time and space. The parameter estimates for these fixed effects can be interpreted as average daily water consumption among households within each uniform-price utility service area, within each season, controlling for all of the remaining covariates in the full-sample models. We place these fixed effects, devoid of any household-level variation in home characteristics and socio-demographics, and devoid of any sub-city-level variation in daily weather, on the left-hand side of a simple second-stage model. This left-hand-side variable contains only utility-level variation in price and characteristics omitted from the main model, some of which we believe to be correlated with price elasticity. We run ordinary least squares (OLS) regressions of these fixed effects on marginal price and utility-level characteristics that may be driving the observed difference in price elasticity among price structures: long-run temperature, long-run precipitation, and the approximate budget for water conservation programs (per residential connection) of the household's water utility in the year the data were collected. Results are reported in Table 4.

This model comprises only 13 observations, so we include only two regressors at a time, and our conclusions are necessarily limited. In general, there are no surprises regarding the impact of these city-level characteristics on average water consumption. To the extent that these variables

[19]Taking into account the variation in uniform prices, alone, over time and space, we have 13 degrees of freedom. Variation in block prices (both across utilities and within the individual price structures, themselves) is much more extensive.

have statistically significant effects, water consumption is increasing in long-run average July and January temperatures, and decreasing in both long-run average precipitation and utility expenditures per account on water conservation programming.[20]

The price coefficient estimates in these models can be considered price elasticities for the uniform-price households. Price elasticities vary greatly in magnitude and precision with model specification. Elasticity estimates for these uniform-price cities as a group range from -.01 (indistinguishable from zero) in a model controlling for long-run average July temperature to -.96 in a model controlling for per-connection expenditures on water conservation programming. Even with no variation in the shape of supply, price elasticity appears to be very sensitive to omitted cross-sectional variables.

These results do not rule out the possibility of a behavioral reaction to the shape of the supply curve. But they lend support to the hypothesis that underlying city-level heterogeneity in characteristics such as aridity and conservation programs, which may be correlated with both consumers' price elasticity and utilities' price schedule choice, explains part of the observed pattern of higher price elasticity under IBPs.

6. CONCLUSION

This is the first study to have addressed both aspects of the separation of demand and supply under non-linear pricing – simultaneous choice of price block and quantity consumed at the household level, and endogenous price structure at the utility level. We estimated price elasticity

[20]The conservation budget variable is likely endogenous (utilities with higher overall water demand may be more likely to implement conservation policies and programs). We expect that the true relationship of *cbudget* to water demand is negative. Endogeneity would tend to drive the parameter estimate toward zero, shrinking the measured effect in an OLS model, implying that the true coefficient may be larger in magnitude than the estimate we report here.

of water demand among urban households facing a variety of price structures, including increasing block pricing and uniform marginal prices, using a structural model that accommodates piecewise-linear budget constraints. We exploited 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 using such models.

In response to a literature that suggests that price elasticity varies with the shape of administratively-determined supply curves, we estimated separate price elasticities for two sub-samples of households — those facing IBPs and those facing uniform marginal prices. We observed a difference in elasticity (indeed, we failed to identify a price elasticity significantly different from zero for the uniform-price households, alone), and posited a set of potential explanations for these differences. We eliminated a number of possibilities that would be consistent with economic theory. We then tested for the remaining suspect among neoclassical explanations — the possibility of endogenous price structures. Results from our tests for endogenous price structures did not eliminate the possibility of some kind of behavioral response to the shape of supply, in addition to the magnitude of marginal price, but they cast doubt on this possibility as the sole explanation.

Even if our results indicate that water utilities should not adopt IBPs with the sole intention of increasing the potential "bang for the buck" (water demand reductions for a given price increase) from price-based conservation policies, an important efficiency justification for the implementation of IBPs remains. To the extent that IBPs increase the portion of consumers facing efficient prices on the margin, they may well be welfare-improving. Exploration of the efficiency advantages of IBPs, and of consumer responses to non-linear prices, is an important area for further research.

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Table 1. Descriptive 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
income	Gross annual household income	\$000/yr	69.81	67.67	5.00	388.64
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
blockp	Price structure: 1 if IBP; 0 if UP	0/1	0.61	0.49	0	1
seas	Season: 1 if irrigation season; 0 if not	0/1	0.51	0.50	0	1
weath	Evapotranspiration less 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	number	2.79	1.34	1	9
bthrm	Number of bathrooms in household	number	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

Notes: City fixed effects not summarized here. Approximately 10 percent of observations drawn from each city (Las Virgenes MWD, Seattle, San Diego, Tampa, Phoenix, Tempe/Scottsdale, Waterloo/Cambridge, Walnut Valley WD, Lompoc).

Table 2. Water Demand Model Estimates

Variable	Discrete/Continuous Choice model, Full Sample	Split Sample Models	
		DCC – IBP Households Only	Random Effects – UP Households Only
lnprice	-.3407*	-.6411*	-.3258
	(.0298)	(.0424)	(.5644)
lnincome	.1306*	.1959*	.0432
	(.0118)	(.0163)	(.0499)
seas	.3072*	.3370*	.2656*
	(.0247)	(.0330)	(.0331)
weath	.0079*	.0094*	.0054*
	(.0013)	(.0017)	(.0016)
maxt	.0196*	.0238*	.0193*
	(.0018)	(.0024)	(.0025)
famsz	.1960*	.1959*	.2095*
	(.0056)	(.0080)	(.0225)
bthrm	.0585*	.0553*	.0498
	(.0093)	(.0127)	(.0393)
sqft	.1257*	.2001*	-.0066
	(.0140)	(.0191)	(.0600)
lotsz	.0065*	.0097*	-.0010
	(.0009)	(.0011)	(.0046)
age	.0869*	.1814*	-.0870
	(.0219)	(.0304)	(.0940)
age2	-.0137*	-.0252*	.0109
	(.0036)	(.0050)	(.0160)
evap	.2278*	.2416*	-.8139
	(.0300)	(.0334)	(.6129)
lasv	.2590*	.3391*	
	(.0408)	(.0486)	
seat	-.1232*	-.0482	
	(.0379)	(.0491)	
sandg	.0129	.1361*	
	(.0365)	(.0416)	

VARIABLE	DISCRETE/CONTINUOUS CHOICE MODEL, FULL SAMPLE	SPLIT SAMPLE MODELS	
		DCC – IBP Households Only	Random Effects – UP Households Only
tampa	-.3879*	-.3225*	
	(.0373)	(.0427)	
phx	-.0044	-.0271	
	(.0385)	(.0444)	
tscot	-.1033*	-.1436*	
	(.0359)	(.0406)	
eug	.0050		-.0018
	(.0387)		(.5846)
wcamb	-.1939*		-.1646
	(.0362)		(.1568)
wvall	.1784*		.3272*
	(.0387)		(.1320)
lomp	-.0647		-.0440
	(.0373)		(.1132)
constant	-3.6994*	-4.3288*	-3.1353*
	(.0652)	(.0888)	(.6854)
σ_η	1.0769*	1.1493*	
	(.0103)	(.0115)	
σ_ϵ	.3552*	.4220*	
	(.0277)	(.0227)	

Notes: *Significant at α=.05. Standard errors in parentheses below estimates. Dependent variable is natural log of daily water demand. Models in columns 1and 2 are discrete/continuous choice, estimated by maximum likelihood. Model in column 3 is random effects GLS regression for panel data. City fixed effects are relative to Denver, CO, except in column 3, where they are relative to Seattle, WA. For full-sample model (column 1), N=25,666. For IBP split-sample model (column 2), N=15,621; and for UP split-sample model (column 3), N=10,045.

Table 3. Price and Income Elasticities of Demand

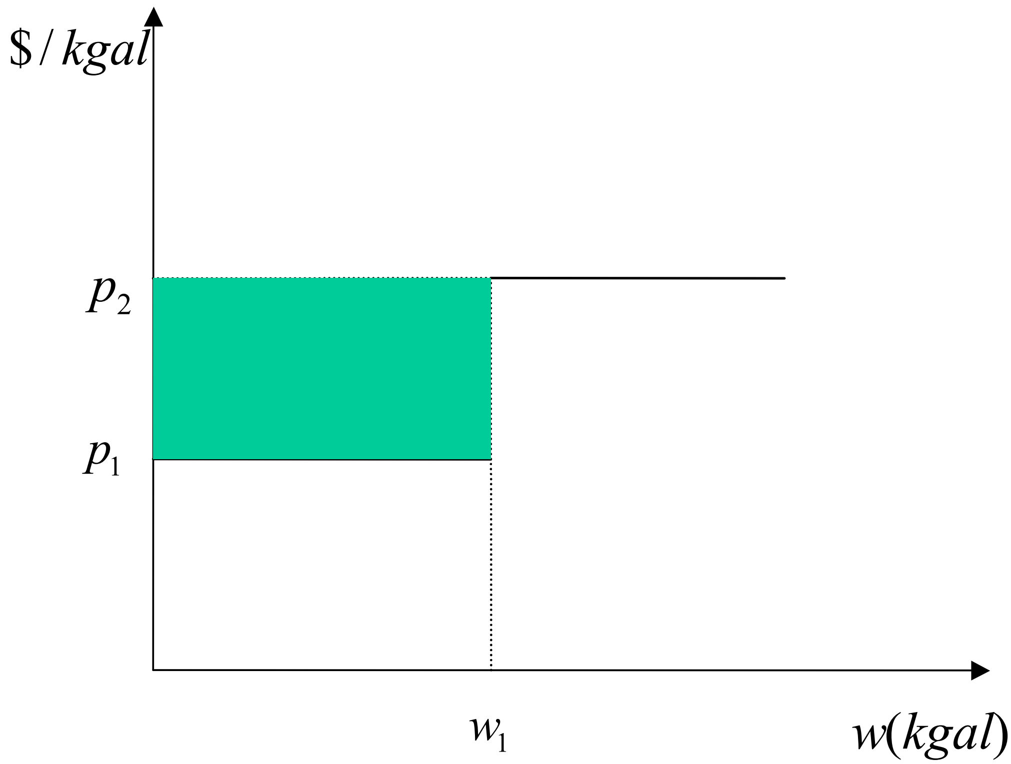
	Elasticity Estimates	
Model	**Price**	**Income**
Discrete / Continuous Choice – full sample	-.3319* (.0366)	.1273* (.0133)
Discrete / Continuous Choice – block-price households only	-.6090* (.0568)	.1865* (.0159)
Random Effects – uniform price households only	-.3258 (.5644)	.0432 (.0499)

Notes: *Significant at α=.05. Elasticity estimates for discrete/continuous choice models (rows 1 and 2) calculated by taking expectation of conditional demand in exponential form, simulating a one percent change in the explanatory variable, and calculating resulting percent change in expected demand. Standard errors obtained by Monte Carlo simulation. See Appendix A for a detailed description of simulations and standard error calculations. Simulated DCC elasticity estimates for other independent variables do not vary appreciably from coefficient estimates, and are not reported here. Elasticity estimates for random effects model (row 3) are repeated from Table 2, for comparison purposes (no simulations necessary for this model).

Table 4. Informal Tests for Endogenous Price Structures

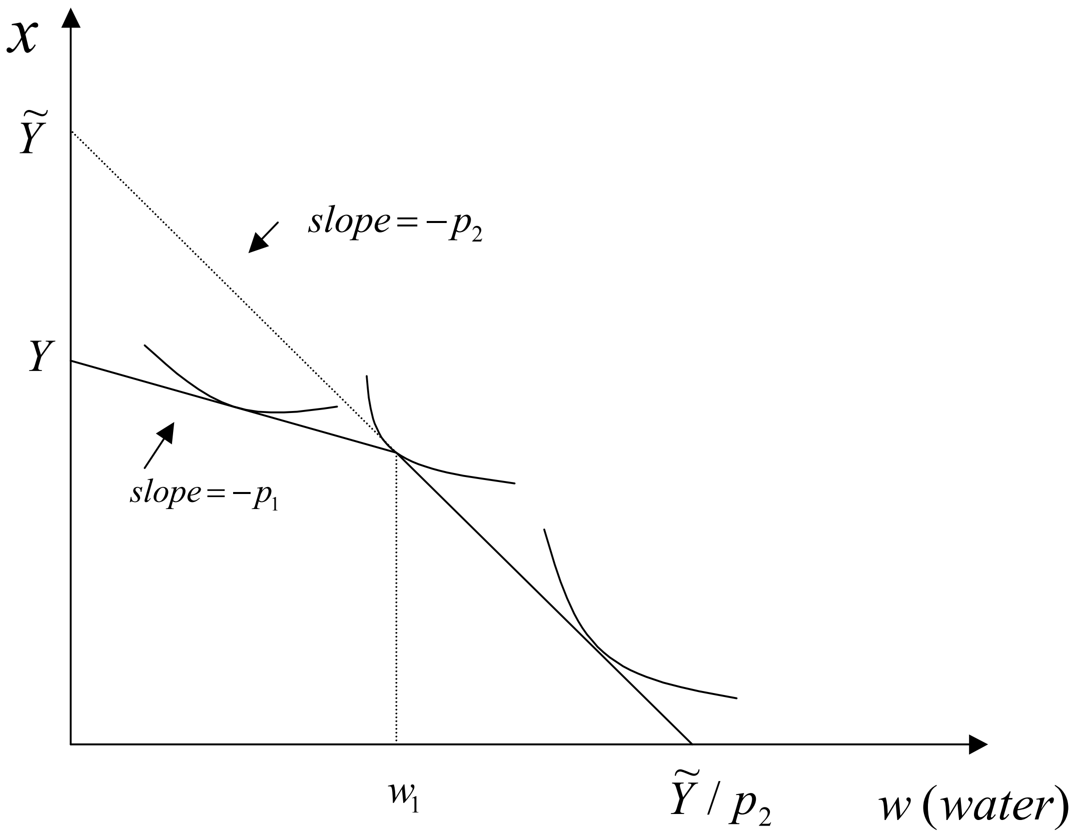
	Parameter Estimates				
Variable	***Model A***	***Model B***	***Model C***	***Model D***	***Model E***
lnprice	-.0695 (.0391)	-.0914* (.0466)	-.0120 (.0475)	-.2037* (.0867)	-.9632* (.3243)
long-run average January temperature		.0113* (.0043)			
long-run average July temperature			.0245* (.0103)		
long-run average precipitation				-.0032 (.0022)	
conservation budget per residential connection					-.0487* (.0111)
constant	.0199 (.0328)	-.0049 (.0252)	-.6513* (.2801)	.3110 (.2095)	.8906* (.2710)

Notes: *Significant at α=.05. Standard errors, robust to clustering by utility, in parentheses. Dependent variable is set of city/season fixed effects for uniform price cities, estimated in first-stage equation (full-sample DCC model in which all price variation is absorbed by fixed effects). First-stage results are not reported here. The price coefficient in the first-stage model is -0.33 (t-statistic=-10.76). Models in Table 4 are all ordinary least squares regressions, with N=13, except for Model E (N=12), as conservation budget information was not available in Eugene, OR.



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



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