The Value of Scarce Water: Measuring the Inefficiency of Municipal Regulations

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ABSTRACT

Rather than allowing prices to capture scarcity rents during periods of drought-induced excess demand, policy makers have mandated the adoption of specific technologies and the curtailment of certain uses, primarily outdoor watering. Using unique panel data on residential end-uses of water, we examine the welfare implications of typical drought policies. Using price variation across and within markets, we identify end-use specific price elasticities. Our results suggest that the current policies target those water uses households, themselves, are most willing to forgo. Nevertheless, we find that use restrictions have costly welfare implications, primarily due to household heterogeneity in willingness-to-pay for water under conditions of scarcity.

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1. Introduction

Competitive markets are the most efficient avenues through which to allocate scarce resources to their highest-valued uses, under the standard assumptions. Yet, many important resources traditionally have not been managed through markets. Prominent examples include space on public roadways and airport runways, water supply, some telecommunications networks, electricity transmission and distribution, and, until recently, broadcast space on the electromagnetic spectrum and electricity generation. These goods and services have been heavily regulated, for a variety of reasons, with allocation and pricing (if any) controlled by public authorities. Where markets are not employed to allocate scarce resources, the potential welfare gains from a market-based approach can be estimated.

We assess the potential welfare gains, and possible distributional outcomes, from switching from non-market to market-based regulation of municipal water supply during periods of drought. During droughts, municipal water restrictions focus almost exclusively on the residential sector, rather than on commercial and industrial water users.[1] Rather than allowing prices to capture scarcity rents during periods of excess demand, policy makers have mandated the adoption of specific technologies and the curtailment of certain uses, primarily outdoor watering, requiring the same limitations on consumption of all households. If households are heterogeneous in willingness-to-pay for water under conditions of scarcity, a price-based approach to drought policy has a theoretical welfare advantage over the current approach.

Using unique panel data on residential end-uses of water for 1,082 households in 11 urban areas in the United States and Canada, we examine the implications of the current approach to urban drought. Using price variation across and within markets, we identify price elasticities specific to indoor and outdoor water demand. Our econometric approach presents a substantial challenge, in that: (1) sixty percent of sample households

[1] Of municipal water consumption, one-half to two-thirds is residential (Dixon et al. 1996, Gleick et al. 2003).

face increasing-block prices, creating piecewise linear budget constraints under which demand and marginal price are systematically correlated; and (2) outdoor water use is censored at zero – households do not use water outdoors every day.

We use these estimates to determine whether current policies target the water uses that households, themselves, would reduce in response to price increases. We find that outdoor watering restrictions do mimic household reactions to price increases, on average, as outdoor use is much more price-elastic than indoor use. However, the real advantage of market-based approaches lies in their accommodation of heterogeneous marginal benefits. During periods of scarcity, in which regulators impose identical frequency of outdoor watering, shadow prices for the marginal unit of water may vary greatly among households.

We divide households into four groups, based on income and lot size, and estimate separate end-use elasticities by group to assess the degree of household heterogeneity in these markets. We estimate shadow prices for water among all customers under four drought policy scenarios of increasing stringency. We then estimate utility-level market-clearing prices for these drought policy scenarios. Under a drought policy limiting outdoor watering to two days per week, estimated shadow prices are, on average, 185 percent higher than the current average marginal water price in these markets, while market-clearing prices are 124 percent higher than current prices.

Our results have important implications for the efficiency of current policies. By ignoring consumer heterogeneity, current drought policies misallocate resources; for an average summer drought, we estimate welfare losses from the typical command-and-control regulation of approximately $73 per household, about 29 percent of average annual household expenditures on water in our sample.

However, switching to a market-based policy also would have distributional consequences. A market would enable those customers who are the least price sensitive, the wealthy consumers with large lots, to consume at levels approximately equal to those without a drought. In contrast, the poor may stop using outdoor water entirely, even under modest drought conditions. The financial consequences of these changes depend on the allocation of water rights.

The paper proceeds as follows. In Section 2, we discuss the related literature. Section 3 examines the econometric models, and Section 4 presents the data. In Section 5, we show our price elasticity estimates for various end uses and consumer groups. In Section 6, we discuss the economic consequences of current regulatory policies and distributional impacts of switching to a price-based allocation mechanism. In Section 7, we conclude.

2. Related Literature

The questions addressed in this paper arise from the general theoretical literature on the conditions under which gains in social welfare are possible through the introduction of markets for managing scarcity (Weitzman 1977, Suen 1990, Newell and Stavins 2003). Studies have produced theoretical and empirical estimates of the gains from increasing the influence of markets on: the distribution of space in the broadcast spectrum (Melody 1980, McMillan 1994); traffic congestion on roadways (Winston 1985, Kraus 1989, Small and Yan 2001, Parry and Bento 2002) and at airports (Daniel 2001, Pels and Verhof 2004); electricity generation (Gilbert *et al.* 1996, De Vany and Walls 1999, Kleit and Terrell 2001); and water supply across sectors (Howe et al. 1986, Hearne and Easter 1997). The gains from market-based approaches to resource allocation in these cases derive largely from heterogeneity in marginal benefits across potential consumers.

A related, prolific literature has focused on welfare comparisons of market-based and command-and-control approaches to environmental policy (Pigou 1920, Crocker 1966, Dales 1968, Montgomery 1972, Baumol and Oates 1988, Tietenberg 1995). The gains from market-based approaches in the case of pollution control policies derive from cost heterogeneity among regulated firms. The advantage of market-based policies over command-and-control policies for pollution control has been demonstrated empirically for the regulation of lead in gasoline (Kerr and Maré 1997), sulfur dioxide in power plant emissions (Burtraw *et al.* 1998), and many other applications.

In the literature on the economics of pollution control, analysts have focused on distributional outcomes where pollutants are non-uniformly mixed, creating variation in the marginal social benefits to different parties of market-based policies (i.e., some get more "clean air" than others). In the smaller, general resource allocation literature, distributional concerns arise from a related issue – the fact that high-value consumers of the good or service will purchase more, and low-value consumers less, if a market is introduced. This is the advantage of a market; it is the source of welfare gains. But under certain conditions, like rationing during wartime, or in the aftermath of a natural disaster, willingness-to-pay, and its close associate, ability to pay, seem unjust allocation criteria.

Weitzman (1977) describes goods and services for which this is the case a "class of commodities whose just distribution to those having the greatest *need*," (emphasis added) as distinguished from want or preference, "is viewed by society as a desirable end in itself." Within this formulation of the problem, the price system turns out to be the more effective way to allocate "needs" when preferences are heterogeneous or income distribution egalitarian; rationing is more effective under the opposite conditions.

Water for residential consumption is an example of a good, some portion of which may fall within Weitzman's need-based commodity regime (drinking, bathing, cooking), and some portion of which would not (swimming pools, lawns, hot tubs). The standard approach to allocating water under conditions of scarcity, wisely, restricts these less necessary uses, not basic needs. Yet, the standard approach does not recognize heterogeneity in willingness-to-pay for these uses, and thus is likely to result in welfare losses, when compared with management through prices.

To date, few studies have addressed municipal drought policies in this framework.[2] Collinge (1994) proposes a theoretical water entitlement transfer system similar to the municipal-level market-based approach we analyze empirically. One experimental economics study simulates water consumption from a common pool, and predicts that customer heterogeneity will generate welfare losses from command-and-control water conservation policies (Krause *et al.* 2003). Neither of these analyses estimates the magnitude of potential welfare gains, nor do they explore distributional implications in any depth. Renwick and Archibald (1998) estimate water demand elasticities by income quartile in Santa Barbara, California, and use these estimates to compare the distributional implications of price and non-price water conservation policies, but do not consider welfare impacts.

A few studies have empirically estimated the impacts of non-price conservation programs on aggregate residential water demand, sometimes jointly with price elasticities (Michelsen *et al.* 1998, Corral 1997). Such studies analyze city-level demand across cities, constructing indices of the relative stringency of residential demand management programs. This approach impedes the direct comparison of the costs of specific non-price programs with hypothetical price increases to achieve the required level of reduction in aggregate water use.[3] We know of only one case in which such a comparison has been made. Timmins (2003) compares a mandatory low-flow appliance regulation with a modest water tax, using aggregate consumption data from 13 groundwater-dependent California cities. Under all but the least realistic of assumptions, he finds the tax to be more cost-effective than the technology standard in reducing groundwater aquifer lift-height in the long run.

In part, the dearth of studies analyzing the welfare impacts of type-of-use restrictions in the residential sector is attributable to a variety of methodological

[2] This is surprising, given that energy policies have frequently been examined in this context (Jaffe and Stavins 1994a, 1994b, Schipper 1979, Hirst and Goeltz 1984, Wirl and Orasch 1998, Wirl 1997).

[3] Not all studies of non-price demand management are able to detect an impact of such programs on residential demand. One study of draconian outdoor watering restrictions in Corpus Christi, Texas, during a drought of record in summer 1996 found no impact of the restrictions on residential consumption (Schultz *et al.* 1997).

challenges. Few data reliably disaggregate residential water consumption into its component uses. In addition, the presence of non-linear prices and the censoring of outdoor demand at zero complicate econometric analyses of this type. In our choice of econometric models, we draw on a well-developed literature on the price elasticity of residential water demand to meet these challenges.

Of the hundreds of water price elasticity studies published since 1960, a few are related closely to our current work.[4] One end-use water demand study has been performed, including estimates of price elasticity for specific uses of water (Mayer *et al.* 1998). This study was the first to use data that reliably separate household water consumption into its component end uses. We use the same data, but ask different questions, and account for price endogeneity.[5]

Non-linear prices typically require a structural approach to obtain unbiased estimates of price elasticity and other model parameters. In the case of increasing-block prices, marginal price and the quantity consumed are positively correlated (Figure 1). Econometric techniques that treat piecewise-linear budget constraints in a manner consistent with utility theory derive from studies of the wage elasticity of labor supply under progressive income taxation (Burtless and Hausman 1978), and have benefited greatly from the generalizations of Moffitt (1986, 1990). Four structural models of water demand under non-linear prices have been estimated: Hewitt and Hanemann (1995), Rietveld *et al*. (1997), Pint (1999), and Olmstead *et al.* (2005). We use the parameter estimates from Olmstead *et al.* (2005) to construct our price instruments, building on their approach to ask and answer policy-relevant questions regarding droughts, non-price demand management, and hypothetical markets.

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[4] Meta-analyses suggest that the central tendency of short-run elasticity estimates over the past four decades is about -0.3, and of long-run estimates about -0.6 (Espey et al. 1997, Dalhuisen et al. 2004).

[5] In estimating demand elasticities, Mayer *et al.* (1998) use average prices, not marginal prices. Ours is the first application of end-use demand estimation to treat non-linear prices in a simultaneous equations framework.

3. Econometric Models

For the estimation of end-use demand, in our case a pair of partial demand equations for indoor and outdoor consumption, the likelihood function in a structural approach would be complicated. The likelihood function for such models is constructed in part by using the block "cutoffs" (the quantities at which marginal price increases), based on total billing period water consumption, to determine the probabilities of consumption at all possible locations along the household's budget constraint (linear segments and kink points). The likelihood function, in our case, would include a total demand equation, which would determine marginal price, and separate end-use demand equations for indoor and outdoor consumption.

We develop an alternative approach here, thanks to the availability of water demand parameter estimates from an existing study using the same data (Olmstead *et al.* 2005). From the structural model, we derive the probability for each household of consuming at each possible marginal price on each day. Probabilities are calculated as functions of the structural parameter estimates, the data, and characteristics of each household's water price structure (number and magnitude of marginal and infra-marginal prices, as well as block cutoffs). We use these probabilities to estimate an expected marginal price, the sum of the products of marginal prices, times the probabilities of facing those prices.[6] Price instruments are then calculated as the seasonal average, by household, of these daily probability-weighted prices. We describe the estimation of these price instruments in greater detail in Appendix A.

We begin with a test of the validity of our identification strategy by examining a model of total daily water demand, using probability-weighted marginal prices as

[6] Kink point probabilities are, on average, 5 percent for two-block price structures, and they range from 1 to 3 percent, on average, for four-block price structures. We divide the kink probabilities evenly (for each household day) between the marginal prices on either side of the kink.

instruments. Using the same data as Olmstead *et al.* (2005), we replicate their results for total water demand, since we know those estimates to be unbiased.[7]

The equation for this total demand function is (1), in which w is total daily water demand for household *i* on day *t*, p is the marginal water price for which we instrument, $\tilde{Y}$ is virtual income, Z is a matrix of daily and seasonal weather variables, and X is a matrix of household characteristics.[8] The error structure comprises θ, a household heterogeneity parameter, and ν, the residual.

$$\ln w_{total_{it}} = \alpha \ln p_{it} + \mu \ln \tilde{Y}_i + \delta Z_{it} + \beta X_i + \theta_i + \nu_{it}. \tag{1}$$

Having tested the usefulness of the price instruments, we proceed with the estimation of end-use models. We adopt different models for indoor and outdoor demand, due to the fact that the fraction of outdoor demand observations equal to zero is approximately 0.58 (and we observe no such censoring of indoor demand). The indoor demand model is identical to the model described in (1) for total demand, but with the log of daily indoor demand, rather than total demand, on the left-hand side. For outdoor demand, we estimate a Tobit model, described in (2).

$$\begin{aligned} w^*_{outdoor_{it}} &= \alpha \ln p_{it} + \mu \ln \tilde{Y}_i + \delta Z_{it} + \beta X_i + \theta_i + u_{it} \\ w_{outdoor_{it}} &= w^*_{outdoor_{it}} \text{ if } w^*_{outdoor_{it}} > 0 \\ w_{outdoor_{it}} &= 0 \text{ if } w^*_{outdoor_{it}} \leq 0 \end{aligned} \tag{2}$$

Dealing with endogenous prices in the Tobit framework involves one extra step to obtain unbiased estimates (Newey 1987). In the first stage, we estimate fitted prices as functions of the price instruments and all of the exogenous covariates. In the second

[7] Identification is not a problem because the estimated probabilities used to create the price instruments incorporate variation in price structures (number and magnitude of prices and block cutoffs), characteristics not incorporated in the second stage of our demand estimation.

[8] Virtual income is annual household income, plus the difference between total water expenditures if the household had purchased all units at the marginal price, and actual total expenditures. This standard technique treats the implicit "subsidy" of the infra-marginal prices as lump-sum income transfers, and is originally due to Hall (1973).

stage, we include both the fitted prices and the residuals from the fitted price equation as independent variables.

4. Data

Daily demand data are drawn from 1,082 households, randomly selected from billing databases in 11 urban areas across the United States and Canada, served by 16 water utilities. Observed households are detached, single-family homes, with no apartments or other multi-family housing in the sample.[9] Table 1 provides descriptive statistics.

Households were observed for two periods of two weeks each, once in an arid season, and once in a wet season. Total demand was disaggregated into its indoor and outdoor components using magnetic sensors attached to water meters. These sensors recorded water pulses through the meter, converting flow data into a flow trace, which detects the “flow signatures” of individual residential appliances and fixtures (Mayer *et al.* 1998). We add together consumption from all indoor fixtures to obtain indoor demand, and consumption from all outdoor uses (irrigation and pools) to obtain outdoor demand. Leaks and unknown uses are included in total demand, but are not modeled explicitly as either indoor or outdoor consumption.[10]

Water demand varies by season, but only for outdoor use. Outdoor water demand in an arid season is, on average, five times outdoor demand during a wet season. In addition, the fraction of observations using any water outdoors, at all, is 0.42 – the reason for choosing a censored regression model.

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[9] Households were randomly sampled from single-family homes within utility customer databases. Sampling procedures, response rates, and statistical tests for selection are described in Mayer *et al.* (1998), Appendix A.

[10] Leaks comprise approximately 6 percent of total consumption, and unknown uses approximately 1 percent. In outdoor use, we can distinguish between water consumption for swimming pools from that for all other outdoor uses. We cannot distinguish among irrigation, car-washing, and washing of sidewalks and driveways, but these uses are all typically regulated or prohibited by drought policies.

In our tests of consumer heterogeneity, we divide the sample into four subgroups, based on income and lot size. Income is our best available proxy for ability to pay, and lot size is our best available proxy for preferences for the services that households derive from outdoor water consumption (such as lawns, gardens, pools, and looking better than the neighbors). Those with both incomes and lot sizes above the medians ($55,000 per year, and 9,000 ft^2, respectively) are categorized as "rich, big lot" households; those with both incomes and lot sizes below the medians are categorized as "poor, small lot" households; and so on for the two groups in between. In the absence of any drought policy these groups divide total sample water consumption as follows: rich, big lot (43 percent); rich, small lot (23 percent); poor, big lot (15 percent); poor, small lot (19 percent).

The households in the sample face either uniform marginal prices (39 percent); or two-tier (44 percent) or four-tier (17 percent) increasing block prices.[11] Given cross-sectional variation and changes over time, the data contain 26 price structures and 47 different marginal prices, ranging from zero to just under $5 per thousand gallons. Average total expenditures on water in the sample, including fixed charges, are $256 per year, or about 0.51 percent of average annual household income.

Price variation in the sample is primarily in the city cross-section. If we regress observed marginal prices on our set of regional dummies, we obtain an R-squared of 0.30. Regressing prices on city fixed effects results in an R-squared of 0.69, on household fixed effects, 0.92, and on our price instruments, 0.81. Table 1 demonstrates that the mean and standard deviation of our price instrument, $\hat{p}$, are similar to the mean and standard deviation of observed marginal water prices.[12] Data on household

[11] Less than one-third of households in the United States face increasing-block prices. Thus, these price structures are over-sampled in the data. This matters for elasticity estimates only if elasticity varies across price structures – an unresolved empirical question (Olmstead *et al.* 2005).

[12] For some sample utilities, marginal wastewater charges are assessed on current water consumption. In addition, some sample utilities benchmark water use during the wet season as a basis for volumetric wastewater charges assessed the following year. For households observed during these periods (and there are some in the data), effective marginal water prices include some function of the present value of expected future wastewater charges associated with current use. Marginal wastewater charges are excluded from the present analysis.

characteristics, including information on annual income, family size, square footage and age of homes, lot size, number of bathrooms, and the presence of evaporative cooling were gathered by survey.[13] We control for season (arid vs. wet), as well as daily weather variables, including maximum daily temperature, and evapotranspiration less effective rainfall (0.6 times total rainfall). Finally, we construct regional fixed effects to control for long-run climate variation not absorbed by the daily and seasonal weather variables.[14]

5. Results

5.1. End-Use Price Elasticity

Our first task is to estimate a total water demand model as in (1), using our 2SLS approach, hoping to obtain parameter estimates close to those we know to be unbiased. Table 2 reports coefficient estimates and standard errors from two such models. The first column contains estimates from Olmstead *et al.* (2005), for the purpose of comparison. The second column reports estimates from a 2SLS random-effects model for panel data, in which the independent variables in the demand function are identical to those in Olmstead *et al.* (2005). This model generates parameter estimates that are similar to those from the structural model, with an important exception – the price elasticity estimate is not significantly different from zero.

In the third column of Table 2, we group the city-level fixed effects from "test model 1" into regional fixed effects. This model captures exogenous sources of geographic and climatic variation, leaving enough price variation to identify a price elasticity. With regional fixed effects, we obtain estimates that are very close to those

[13] Evaporative cooling, common in arid climates, substitutes water for electricity in the provision of air conditioning. Less than 10 percent of sample households have evaporative coolers, but 43 percent of sample households in Phoenix have them, and about one-third of households in Tempe and Scottsdale. Households with evaporative cooling use, on average, 35 percent more water than households without.

[14] Regions are as follows: (1) Southern California (Las Virgenes Municipal Water District, City of San Diego, Walnut Valley Water District, and City of Lompoc); (2) Arizona and Colorado (Phoenix, Tempe, Scottsdale, and Denver); (3) Northern (City of Seattle Public Utilities, Highline Water District, City of Bellevue Utilities, Northshore Utility District, Eugene Water and Electric Board, Regional Municipality of Waterloo, Ontario).

from the structural model with, unsurprisingly, somewhat less precision. The price elasticity estimate is -0.35, and strongly significant. We use test model 2, a 2SLS random-effects model with regional fixed effects, as our point of departure for the rest of the analysis.

We then separate total demand into indoor and outdoor consumption and estimate the models given in (1), for indoor use only, and (2). In Table 3, we report the full set of parameter estimates and standard errors (with the exception of constants and regional fixed effects) for indoor and outdoor demand models, first annually, and then by season.

Indoor use appears to be influenced by income and family size, and little else. Outdoor demand parameters are all significant, with the exception of home age in some models. Many significant outdoor demand parameters would seem to be drivers of indoor, rather than outdoor consumption (for example, the number of bathrooms). It may be that these parameters are correlated with omitted variables that represent preferences for outdoor water consumption.

Table 4 summarizes the results of these models with respect to price elasticity, the parameter of interest, and reports elasticities, rather than price coefficient estimates, for the outdoor models.[15] Our discussion proceeds on the basis of the summary of elasticity estimates in Table 4.

The partial demand models reveal striking variation in elasticity across uses and, for outdoor use, across seasons. None of the indoor elasticity estimates are significantly different from zero (the estimates are very small in magnitude, as well). Outdoor demand in the wet season is the most price-elastic (-1.28), and is still quite responsive to price in

[15] The Tobit coefficients are not price elasticities – see the notes to Table 4 for the calculation of Tobit elasticities. For these estimates, we calculate the Tobit model allowing different coefficients on price for in season and off-season.

the arid season (-0.75). In fact, essentially all of the strong seasonal variation we observe in total water consumption is attributable to outdoor use.[16]

Results reported in Table 4 might suggest that regulators' focus on outdoor consumption as a target of command-and-control water conservation policies provides a good first approximation to a market-based approach. Indeed, outdoor uses are the uses that households, themselves, would choose to cut back the most in response to a price increase. But an important cost of the command-and-control approach derives from the heterogeneity of regulated entities – in this case, households. If households are heterogeneous in their preferences for outdoor consumption, across-the-board outdoor water use restrictions will ignore that heterogeneity, generating welfare losses relative to a market-based approach.

5.2. Robustness

We test the robustness of our elasticity estimates by exploring a number of other model specifications. These include demand functions with household fixed effects and functions with city fixed effects (rather than regional fixed effects). We also examine a model that collapses the daily variation in household demand, obtaining parameter estimates from regressions of aggregate seasonal household demand on the covariates of interest. We find our results to be qualitatively robust to these models. See Appendix B for details.

5.3. Consumer Heterogeneity

To test whether households are, in fact, heterogeneous in their preferences for water consumption, we divide the sample into four sub-groups, based on income and lot size. We estimate separate elasticities for the four groups. Results, reported in Table 5, indicate a high degree of heterogeneity. Households presumed to have the strongest

[16] We have no information on available substitutes for municipal tap water in outdoor uses, which might include groundwater wells or public surface water sources. To the extent that these substitutes are available in the sample, our outdoor elasticity estimates are lower than they would be in the absence of substitutes.

preferences for outdoor water consumption, the “rich, big lot” group, exhibit the least elastic outdoor demand (-0.45). Those presumed to have the weakest preferences for outdoor consumption, the “poor, small lot” group, exhibit the most elastic outdoor demand (-0.86). Lot size appears to make a larger difference than income in this regard, as the “rich, small lot” group (-0.76) is somewhat more elastic than the “poor, big lot” group (-0.69).[17]

6. Discussion

In this section, we discuss the implications of consumer heterogeneity for water policy. In particular, we discuss the variation in households’ willingness-to-pay for water under the current drought policies (i.e., their shadow prices). Command and control regulations typically limit the number of days in a week that households may use water outdoors, whether for watering lawns, washing automobiles, or filling swimming pools. A common policy is to limit outdoor watering to two days a week. We examine the implications of this policy, as well as limits of three, one, and zero days per week. Then we discuss the allocation in a market for water. Finally, we examine the potential welfare gains and distributional implications of such a market.

6.1. Shadow Prices

This evidence of heterogeneity among households suggests that, when outdoor water consumption is restricted by drought policy, shadow prices for the marginal unit of water will vary significantly. Variation in shadow prices would indicate potential gains from trade. For example, a market would lead to smaller reductions in outdoor consumption by the least elastic groups, larger reductions by the most elastic groups, and, perhaps, small reductions in indoor use by all groups (Figure 2). Based on the separate

[17] F-tests find “rich, big lot” to differ significantly from “rich small,” “poor big,” and “poor small” (P-values are 0.01, 0.06, and 0.01, respectively). The other categories did not differ significantly from each other. In the arid season, the season in which drought regulations are implemented most frequently, the differences between the “rich, big lot” group and all others are even more pronounced (and the differences among the three other groups less pronounced). During the arid season, the “rich, big lot” outdoor elasticity is only weakly significant, and is less than half the magnitude of the next least-elastic group.

elasticity estimates for our household sub-groups, we calculate shadow prices and market-clearing prices under drought policies of varying stringency.

To calculate shadow prices, we estimate the constrained level of consumption for each household under each drought policy, and then "back up" along that household's demand curve to obtain their willingness-to-pay for the marginal unit of water.[18] Some households are unconstrained by the policies—their probability of watering on a given day is less than or equal to the probability imposed by the watering restrictions. For constrained households, we calculate the difference in their expected quantity demanded in the unrestricted and restricted scenarios. For example, for a once-per-week watering policy, the restricted probability of watering is 1/7 (assuming full compliance). So a household with a probability of watering greater than 1/7 is constrained, and will have a resulting decrease in expected quantity demanded.

For the irrigation season, Table 6 reports shadow prices. In the most extreme policy, when no watering is allowed, our nonlinear functional form implies an infinite shadow price for all customers. We assume that the willingness-to-pay is at most $50 per thousand gallons. The most common policy (of allowing outdoor watering two days per week) has an average shadow price of $5.25 per thousand gallons. Note that this is almost three times the average price consumers actually pay ($1.84).

As we would expect, shadow prices increase with the stringency of the drought policy. Furthermore, the standard deviation of shadow prices across all customers is increasing in drought policy stringency. These standard deviations reflect the potential benefits from establishing a market-based policy. For example, Figure 3 provides histograms of shadow prices in two cities given a policy limiting outdoor watering to two days a week.[19] Even in cities with relatively small standard deviations in shadow prices (like Eugene, Oregon), we see there is a lot of variation and that shadow prices tend to be right-skewed. There are some households with much higher shadow prices than the

[18] Indoor and outdoor elasticities reported in Table 5 are used for estimation of shadow and market prices.
[19] We chose two cities (Eugene, Oregon, and San Diego, California) that provide examples with low and average amounts of variation in shadow prices, respectively.

average. Under a market, all households will consume such that shadow prices are equal; all gains from trade will be realized.

6.2. Market-clearing Prices

We then construct market-clearing prices under drought policies of varying stringency. Here we assume that each utility's goal is simply to save the aggregate quantity of water it would save by implementing each type of drought policy, no matter how that aggregate water consumption reduction is achieved. From the households' perspective, these savings could be achieved indoors, outdoors, or by purchasing "credits" from another household.[20] We then identify the market-clearing price for that aggregate reduction, constraining households to non-negative consumption.[21]

In Figure 3, the solid vertical lines denote the market-clearing prices for two cities. In Eugene, where there was relatively little shadow price variation, the market-clearing price is close to the average of the shadow prices. In San Diego, we see a larger difference between the shadow prices and the market price.

The last column of Table 6 reports market-clearing prices across all cities. Like the shadow prices, market-clearing prices increase monotonically with the stringency of watering restrictions. Within a utility, there is a common price. Thus, the only variation in these prices is across utilities. While not as large as the variation within shadow prices, we still find that market-clearing prices vary substantially across utilities. For the most common drought policy, the average market-clearing price is $4.13 per thousand gallons, or slightly more than twice the current mean marginal price.

Market-clearing prices assume that the necessary aggregate (utility-wide) demand reduction is that which would be achieved under full compliance with the various drought

[20] An actual tradable credit system would likely be infeasible due to large transactions costs. However, with no uncertainty, a regulator could equivalently set a higher price so as to clear the market.
[21] Market-clearing prices are estimated for the irrigation season, assuming that drought regulations are implemented primarily during arid months.

policies. We could easily add a probability of compliance that is less than one to our analysis, but doing so would be equivalent to simulating different drought policies than those we selected. In any case, were utilities to estimate market-clearing prices, they might want to charge slightly higher prices, anticipating less than full compliance, depending on the importance of meeting the demand reduction.[22] It is likely that compliance will be higher under a market-based policy than under the current command-and-control approach, since "cheating" in the market context would require that households figure out how to consume piped water outdoors "off-meter", and in the current context can easily be accomplished by watering at night, or in some other way that avoids observation by utility staff or vengeful neighbors.

6.3. Welfare implications

The management of water scarcity through residential outdoor watering restrictions results in substantial welfare losses, given the observed heterogeneity in willingness-to-pay. Welfare losses calculated in this context (with a reduced-form model, and Marshallian demand curves) should be considered very rough estimates. Nonetheless, we do calculate them.[23]

For each utility, we simulate the welfare effects of a two day per week watering policy over a 180-day drought-struck irrigation season. Table 7 summarizes our findings on a per household basis. Deadweight losses (DWL) under the current regime represent the estimated benefit to the average household of introducing a market. Estimated DWL ranges from $1 per household, in the service area of Seattle Public Utilities, to $463 per household, in the service area of Northshore Utility District. The variation in DWL from the command-and-control approach is attributable, in part, to the standard deviation of

[22] This is always the case with a market approach based on prices, rather than quantities (Weitzman 1973). While a quantity instrument (like tradable permits) would be preferable in cases where utilities had very strict quantity constraints, such as the threat of violation of a treaty over shared water resources, the transactions costs involved in establishing a household-level trading regime would likely be prohibitive.
[23] We estimate deadweight loss by integrating demand curves, as discussed in Table 7.

shadow prices, a strong indicator of potential gains from trade.[24] In our sample, society would be better off by $73 per household through the introduction of a market.

The discussion of market-based approaches is largely hypothetical under the current regulatory structure. Estimated market-clearing prices are greater than current average marginal prices in all of these markets, in some cases by very large magnitudes. If they are also higher than average costs, prices this high would be impossible to implement without significant rebates of some form, as utilities in the United States are usually restricted to zero (or very small) profits.

In addition, we face the standard worry about the welfare effects of a theoretical first-best policy in a second-best setting (Lipsey and Lancaster 1956, Harberger 1974). The fact that ours is a partial equilibrium analysis may be of less concern than in the well-known theoretical and empirical studies of environmental taxation in a second-best setting (Sandmo 1975, Goulder *et al.* 1999). The closest case to our own, analytically, is that of proposed congestion pricing regimes. In that case, introducing market-based approaches changes commuting costs, spilling over into labor markets, which are already distorted at the margin by income taxation (Parry and Bento 2002, Small and Yan 2001). No such spillover (no pun intended) is engendered by changes in water expenditures which, in any case, comprise one-half of one percent of household income in our sample.

In our case, the distortions of greatest concern are within water markets, themselves. In most cases, marginal water prices are well below the marginal social cost of water supply (Hanemann 1997, Timmins 2003). Applying a market-based approach to drought policy will result in higher marginal prices for all households (even if total expenditures fall for some households through lump-sum transfers). To the extent that price-based drought management results in more households paying something closer to marginal social cost, the pre-existing distortion does not change the basic nature of our

[24] Presumably, if we included multi-family homes in a market, gains from trade would be somewhat larger. Their inclusion would add their (currently unregulated) indoor use to the "common pool" which, if sensitive to price increases, would be an additional source of reductions to support increases in higher-valued uses.

results.[25] However, if marginal prices in the sample are well below marginal social cost, the welfare impacts of moving to price-based drought management policy may pale in comparison to the impacts of moving to marginal social cost pricing, period. This is an important area for further research, but it is beyond the scope of this analysis.

6.4. Distributional implications

While the shift from outdoor watering restrictions to a price-based municipal drought policy would be welfare-improving in all markets, the distributional implications of shifting regulatory approaches depend on the allocation of property rights. A market would re-distribute scarce water such that those with high willingness-to-pay for water consumption outdoors would consume more than they do under outdoor use restrictions, and those with low willingness-to-pay would consume less.

Figure 4 describes the movement of water consumption across income/lot size quartiles that would be brought about by the shift from command-and-control to market-based allocation of water during periods of scarcity. For the purpose of exposition, we choose a two day per week watering policy, and the market-clearing price that would generate the equivalent level of aggregate demand reduction. The consumption share of rich, big lot households would grow substantially, primarily as a result of consumption reduction by poor, small lot households, and to a smaller extent by the other groups. Hence, a market would result in a water allocation that would "soak the rich."

Under a market relative to the command and control approach, the consumption share of the least elastic group (at least for outdoor uses), the rich, big lot households, would rise from 35 to 48 percent; the consumption share for the most elastic group, the poor, small lot households, would fall from 23 to 16 percent, with smaller reductions in consumption shares by the remaining two quartiles. Absolute consumption falls quite drastically among all quartiles under both types of drought policies.

[25] An increase in the marginal price of municipal water supply will generate an increase in the consumption of substitutes. Where groundwater is a viable substitute for municipal tap water, spatial and intertemporal externalities may result (or increase, where they are already present).

Under the current approach, the largest DWL, as a fraction of average income, is experienced by the rich, big lot households (see Table 8). However, the poor, small lot households experience the second-largest "effective" DWL. Interestingly, above the median lot size, the command-and-control approach hurts rich more than poor households, and below the median lot size, the opposite is true. (We can think of this as differences along the dimension of ability to pay, controlling for tastes.) Policymakers are often willing to compromise on efficiency if a policy is perceived to be progressive; the current CAC approach is neither efficient, nor consistently progressive.

A progressive market-based approach can always be designed through the use of transfers. In the present case, this could occur through the utility billing process. Any exogenous household characteristic (historical consumption, for example) might be used to determine the size of a household's water "budget" over a billing period. All units would be charged at the market-clearing volumetric marginal price, but households below their budget constraint could receive credits toward the next bill; households above the constraint would pay additional charges.[26]

Before concluding our discussion of distributional implications, it is important to ask whether residential water consumption, as a whole, fits into the category of "needs" defined earlier. Weitzman (1977) concludes that the comparative advantage of prices over rationing (δ, in his terms) is equal to twice the variance of demand, conditional on income, less the mean square deviation in demand at market prices, the difference between a "taste distribution effect" and an "income distribution effect", described in (3).

$$\delta = 2V[\varepsilon] - \sigma^2 \tag{3}$$

If tastes predominate in the allocation of the good, it is best left to markets. If income predominates, rationing may be a better alternative. Using our predictions of the quantity of water demanded under a market, we measure a conditional variance, $V[\varepsilon]$, of

[26] Collinge (1994) outlines one plausible tax/rebate system.

0.061 and an unconditional variance, σ^2, of 0.033.[27] We conclude that there is substantial taste variation in water demand. A market will allocate water "needs" better than rationing.

7. Conclusions

Using unique panel data on residential end-uses of water, we examine the welfare implications of command-and-control municipal drought policies. Using price variation across and within markets, we identify price elasticities for indoor and outdoor consumption. Outdoor uses are more elastic than indoor uses, suggesting that current policies target those water uses households, themselves, are most willing to forgo. Nevertheless, we find that use restrictions have substantial welfare implications, primarily due to household heterogeneity in willingness-to-pay for water under conditions of scarcity.

Heterogeneity is often ignored in economic analyses, which proceed from the viewpoint of the "representative consumer." For heavily regulated goods, estimating the welfare gains from introducing markets requires the opposite starting point—it is precisely the variation in marginal benefits that opens up the potential gains from trade within non-market allocations.

Of all the currently regulated markets in which alternative price-based policies have been proposed, municipal water markets may be the easiest in which to imagine actually introducing a market-based approach, even one that involves lump-sum transfers to achieve equity goals. Household water use is metered, and monitored by utility staff for the purpose of billing and collection.[28] Were such a system to be implemented, a

[27] We calculate the conditional variance by regressing total water consumption on a set of indicator variables for each of the 21 income categories in our sample. We further control for other proxies of wealth, including quadratic functions of size of house and size of lot.

[28] This is quite unlike the case of market-based pollution regulation, which requires the installation of continuous emissions monitoring infrastructure (for tradable permits), or the case of congestion pricing, which requires a new system with which regulators can track consumers' use of priced roadways.

municipality would have the rare opportunity to affect an actual Pareto improvement, in which gains not only exceed losses, but no household is made worse off.

If concern about "everyone doing their part" during a drought is the reason for the current predominance of command-and-control, rather than market approaches to the management of scarce water resources, economists' discussion of potential lump sum transfers and actual Pareto improvements may fall on deaf ears. There is irony in this. In the long run, command-and-control regulations provide no incentive for the invention, innovation, and diffusion of water conserving technologies (outdoors or indoors). Water priced below marginal social cost, drought or no drought, also results in inefficient land-use patterns, like the establishment of large, lawn-covered lots and thirsty non-native plant species where water is scarce. Further investigation of the welfare gains from water marketing, both within and across sectors, is both an important area for further research, and an important subject for further dialogue between economists, policymakers, and environmental advocates.

Table 1. Descriptive Statistics

Variable	Description	Units	Mean	Std. Dev.	Min.	Max.
w	Daily household water demand	kgal/day	.40	.58	0	9.78
	in season	kgal/day	.54	.71	0	9.78
	off season	kgal/day	.25	.34	0	7.16
outdoor	Daily water demand outdoors	kgal/day	.22	.55	0	9.50
	in season	kgal/day	.36	.69	0	9.50
	off season	kgal/day	.07	.30	0	6.79
indoor	Daily water demand indoors	kgal/day	.17	.13	0	1.91
	in season	kgal/day	.17	.13	0	1.04
	off season	kgal/day	.17	.13	0	1.91
P(outdoor>0)	Fraction obs. for which outdoor>0		.42	.49	0	1
P(indoor>0)	Fraction obs. for which indoor>0		>.99	.02	0	1
price	Observed marginal water price	$/kgal/mo	1.76	.60	0.5	4.96
phat	Marginal water price instrument	$/kgal/mo	1.70	.53	0.76	4.75
income	Gross annual household income	$000/yr	69.81	67.67	5.00	388.64
seas	Irrigation season=1 / not=0		0.51	0.50	0	1
weath	Evapotransp. less effective rainfall	mm/day	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		2.79	1.34	1	9
bthrm	Number of bathrooms in household		2.58	1.30	1	7
sqft	Area of home	000 ft^2	2.02	0.82	0.40	4.37
lotsz	Area of lot	000 ft^2	10.87	9.22	1.00	45.77
age	Age of home	yrs/10	2.88	1.62	0.07	5
evap	Evaporative cooling=1 / not=0		0.09	0.28	0	1
region1	Southern California		.37	.48	0	1
region2	Arizona/Colorado		.28	.45	0	1
region3	Northern		.26	.44	0	1
region4	Florida		.09	.29	0	1

Table 2. Model Testing Probability-Weighted Prices as Instruments

	DCC Estimates (OHS Paper)		Test Model 1		Test Model 2	
Variable	**Estimate**	**SE**	**Estimate**	**SE**	**Estimate**	**SE**
lnprice	-0.3408*	0.0298	-0.1759	0.1090	-0.3508*	0.0683
lnincome	0.1305*	0.0118	0.1417*	0.0319	0.1422*	0.0328
season	0.3070*	0.0247	0.3117*	0.0212	0.3163*	0.0209
weath	0.0079*	0.0013	0.0081*	0.0011	0.0078*	0.0010
maxtemp	0.0196*	0.0018	0.0193*	0.0015	0.0202*	0.0015
famsize	0.1961*	0.0056	0.1911*	0.0151	0.1940*	0.0156
bathrooms	0.0585*	0.0093	0.0489*	0.0251	0.0571*	0.0260
sqft	0.1257*	0.0140	0.1243*	0.0380	0.1372*	0.0395
lotsize	0.0065*	0.0009	0.0061*	0.0025	0.0078*	0.0025
home age	0.0867*	0.0219	0.0878	0.0591	0.1086#	0.0615
home age^2	-0.0137*	0.0036	-0.0153	0.0098	-0.0187#	0.0102
evap cooling	0.2277*	0.0300	0.2352*	0.0823	0.2448*	0.0831
constant	-3.6994	0.0652	-3.7148*	0.1607	-4.1537*	0.1552
Fixed Effects	City-level		City-level		Region-level	
R^2 overall within between			0.20 0.10 0.35		0.19 0.10 0.32	

Notes: * significant at 5% (# at 10%). Dependent variable is natural log of daily household water demand (kgal). Model in column 1 is discrete-continuous choice model from Olmstead *et al.* (2005). Test models 1 and 2 are two-stage least squares random effects model for panel data, in which we instrument for marginal water prices. Estimates for city-level and region-level fixed effects are not reported. In all cases, N=25,668, with 1,082 households.

Table 3. Models of Indoor and Outdoor Water Demand

Variable	Indoor (annual)	Indoor (by season)	Outdoor (annual)	Outdoor (by season)
lnphat	-0.0713 (0.0567)	-0.0640 (0.0564)	-0.4181* (0.0447)	-0.3858* (0.0453)
lnphat offseas		-0.0353 0.0289		-0.1947* (0.0384)
perr			1.2885* (0.0537)	1.4524* (0.0721)
perr offseas				-0.3672* (0.0914)
lnincome	0.0667* (0.0279)	0.0668* (0.0278)	0.0888* (0.0202)	0.0862* (0.0198)
season	-0.0167 (0.0164)	-0.0346 (0.0219)	0.4274* (0.0212)	0.3197* (0.0290)
weath	0.0006 (0.0008)	0.0006 (0.0008)	0.0074* (0.0012)	0.0076* (0.0012)
maxtemp	-0.0012 (0.0012)	-0.0012 (0.0012)	0.0316* (0.0016)	0.0320* (0.0016)
famsize	0.2389* (0.0132)	0.2392* (0.0132)	0.0354* (0.0093)	0.0380* (0.0095)
bathrooms	0.0028 (0.0220)	0.0031 (0.0220)	0.0535* (0.0159)	0.0553* (0.0157)
sqft	0.0149 (0.0335)	0.0156 (0.0335)	0.1466* (0.0253)	0.1480* (0.0254)
lotsize	0.0031 (0.0022)	0.0031 (0.0022)	0.0110* (0.0017)	0.0105* (0.0017)
home age	0.0660 (0.0522)	0.0660 (0.0521)	0.0610# (0.0373)	0.0587 (0.0369)
home age^2	-0.0135 (0.0086)	-0.0134 (0.0086)	-0.0079 (0.0062)	-0.0072 (0.0062)
evap cooling	0.1477* (0.0705)	0.1485* (0.0704)	0.0834# (0.0506)	0.0928# (0.0520)

Notes: * significant at 5% (# at 10%). Indoor model is a 2SLS Random Effects model. Outdoor is a 2SLS Tobit Random Effects model. Results for constant and regional fixed effects not reported. The number of observations is 25,136 for indoor consumption, and 25,707 for outdoor. The variable *perr* is the residual from the first stage (fitted price) equation.

Table 4. Summary of Elasticity Estimates

Elasticity of	Overall	In Season	Off Season
total consumption	-0.3508* (0.0683)	-0.2769* (0.0679)	-0.5908* (0.0757)
indoor consumption	-0.0713 (0.0567)	-0.0640 (0.0564)	-0.0993 (0.0629)
outdoor consumption	-0.6430* (0.0687)	-0.7491* (0.0890)	-1.2819* (0.1090)

Notes: * significant at 5% (# at 10%). The number of observations is 25,668 for total consumption; 25,136 for indoor; and 25,707 for outdoor (13,181 in season and 12,525 off).

Elasticities for Tobit model are estimated as follows:

$$\alpha^{Tobit} = \frac{\alpha^{Tobit} * \Pr(w_{outdoor} > 0)}{\mu_{w_{outdoor}}}$$

Table 5. Price Elasticities of Demand, by Income/Lot size Sub-group

Household Sub-group	Total Demand Elasticities	Indoor Demand Elasticities	Outdoor Demand Elasticities
Rich, big lot	0.0508 (0.0953)	-0.1172 (0.0789)	-0.4535* (0.1054)
Poor, big lot	-0.3689* (0.1264)	-0.0914 (0.1065)	-0.6949* (0.1212)
Rich, small lot	-0.3961* (0.0853)	-0.0639 (0.0700)	-0.7618* (0.0957)
Poor, small lot	-0.5604* (0.0910)	-0.0419 (0.0756)	-0.8551* (0.0932)

Notes: * significant at 5% ($^{\#}$ at 10%). The number of observations is 7,188 for rich, big lot; 4,016 for poor, big lot; 7,386 for rich, small lot; and 7,117 for poor, small lot.

Table 6. Shadow Prices, Market-clearing Prices under Various Drought Policies

Drought Policy	Current Price Mean ($/kgal) [Std. Dev.]	Shadow Price Mean ($/kgal) [Std. Dev.]	Market-clearing Price Mean ($/kgal) [Std. Dev.]
Status quo (no drought policy)	1.84 [0.64]		
No outdoor watering		50.00 [0.00]	18.00 [14.69]
Outdoor watering once/week		7.62 [10.95]	7.12 [7.53]
Outdoor watering twice/week		5.25 [7.17]	4.13 [3.33]
Outdoor watering 3 times/week		3.67 [4.82]	2.91 [1.72]

Notes: All prices are for irrigation season, only. We assume willingness-to-pay is at most $50 per thousand gallons.

Table 7: Estimated Welfare Impacts by Utility

Utility	Shadow Price ($/kgal) Mean	Std. Dev.	Market Price ($/kgal)	DWL ($/household/summer)
Seattle, WA	3.0	0.3	3.0	1.0
Waterloo, Ontario	2.2	0.7	2.1	3.1
Eugene, OR	1.2	0.5	1.1	3.3
Cambridge, Ontario	1.9	0.8	1.8	3.4
Lompoc, CA	3.1	1.1	2.9	5.9
Highline, WA	3.6	1.6	3.2	10.4
Tampa, FL	2.2	1.5	2.2	12.8
Tempe, AZ	3.2	2.6	2.3	28.5
Bellevue, WA	3.3	2.4	3.0	28.5
San Diego, CA	4.0	3.2	3.1	28.8
Denver, CO	3.7	3.3	2.7	34.4
Phoenix, AZ	5.4	6.1	4.0	103.5
Walnut Valley, CA	7.7	8.3	4.9	107.6
Scottsdale, AZ	7.9	6.8	5.3	109.4
Las Virgenes, CA	16.5	12.6	13.3	323.7
Northshore, WA	10.8	13.3	9.7	463.1
Average	**5.3**	**7.2**	**4.1**	**73.3**

Notes: Shadow prices and deadweight losses (DWL) are calculated for a two-day per week outdoor watering policy. DWL is dollars per household per summer, estimated as the area under constant-elasticity demand curves. For indoor demand, $\widehat{w}_{in} = e^{Z\hat{\delta}} e^{X\hat{\beta}} p^{\hat{\alpha}_{in}} \widetilde{Y}^{\hat{\mu}}$. Let $C_{in} = e^{Z\hat{\delta}} e^{X\hat{\beta}} \widetilde{Y}^{\hat{\mu}}$. Then, $\widehat{w}_{in} = C_{in} p^{\hat{\alpha}_{in}}$ and the integral is: $C_{in} p(\frac{p^{\hat{\alpha}_{in}}}{(\widehat{\alpha}_{in}+1)}) = \frac{\widehat{w}_{in} p}{\widehat{\alpha}_{in}+1}$. For outdoor demand, $\widehat{w}_{out} = \widehat{\alpha}_{out} \ln p + \widehat{\mu} \ln \widetilde{Y} + \widehat{\delta} Z + \widehat{\beta} X$. Let $C_{out} = \widehat{\mu} \ln \widetilde{Y} + \widehat{\delta} Z + \widehat{\beta} X$. Then, $\widehat{w}_{out} = C_{out} + \widehat{\alpha}_{out} \ln p$, and the integral is: $C_{out} p - \widehat{\alpha}_{out} p + \widehat{\alpha}_{out} p \ln p$.

Table 8. Average Deadweight Loss by Quartile

Quartile	Average Deadweight Loss ($/arid season)	Average Deadweight Loss/ Average Annual Income
Rich, big lot	$197.3	.0017
Rich, small lot	23.8	.0003
Poor, big lot	11.7	.0004
Poor, small lot	31.7	.0011

Figure 1. Two-tier increasing block price structure

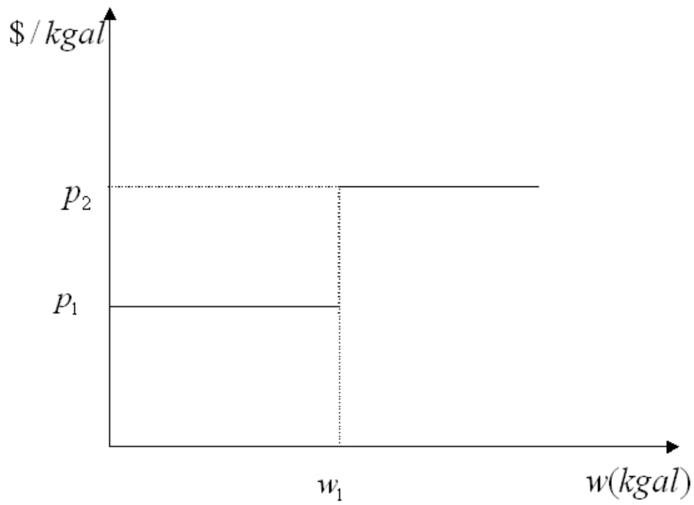
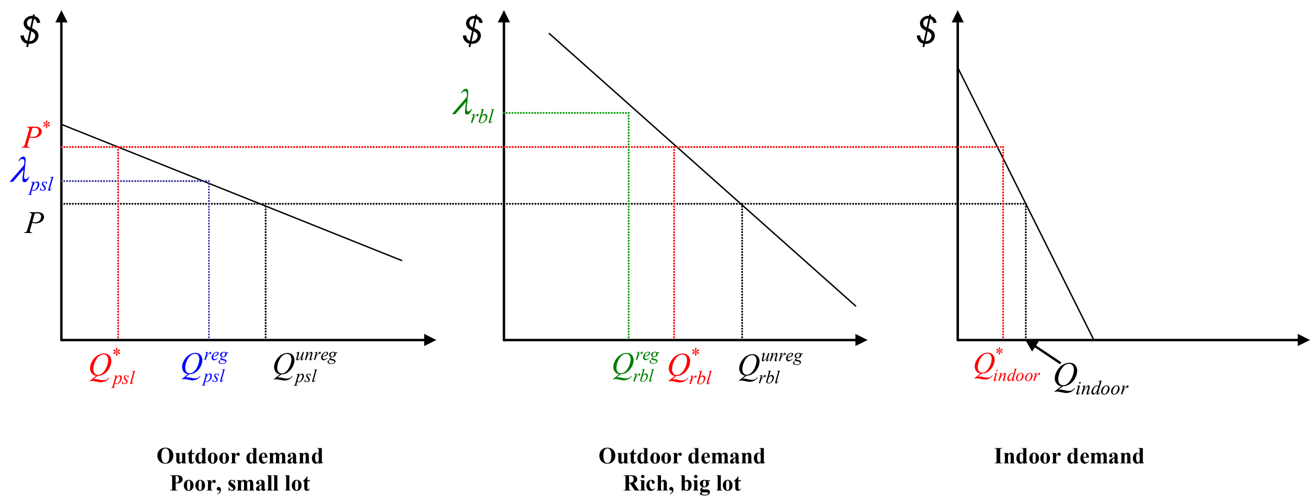
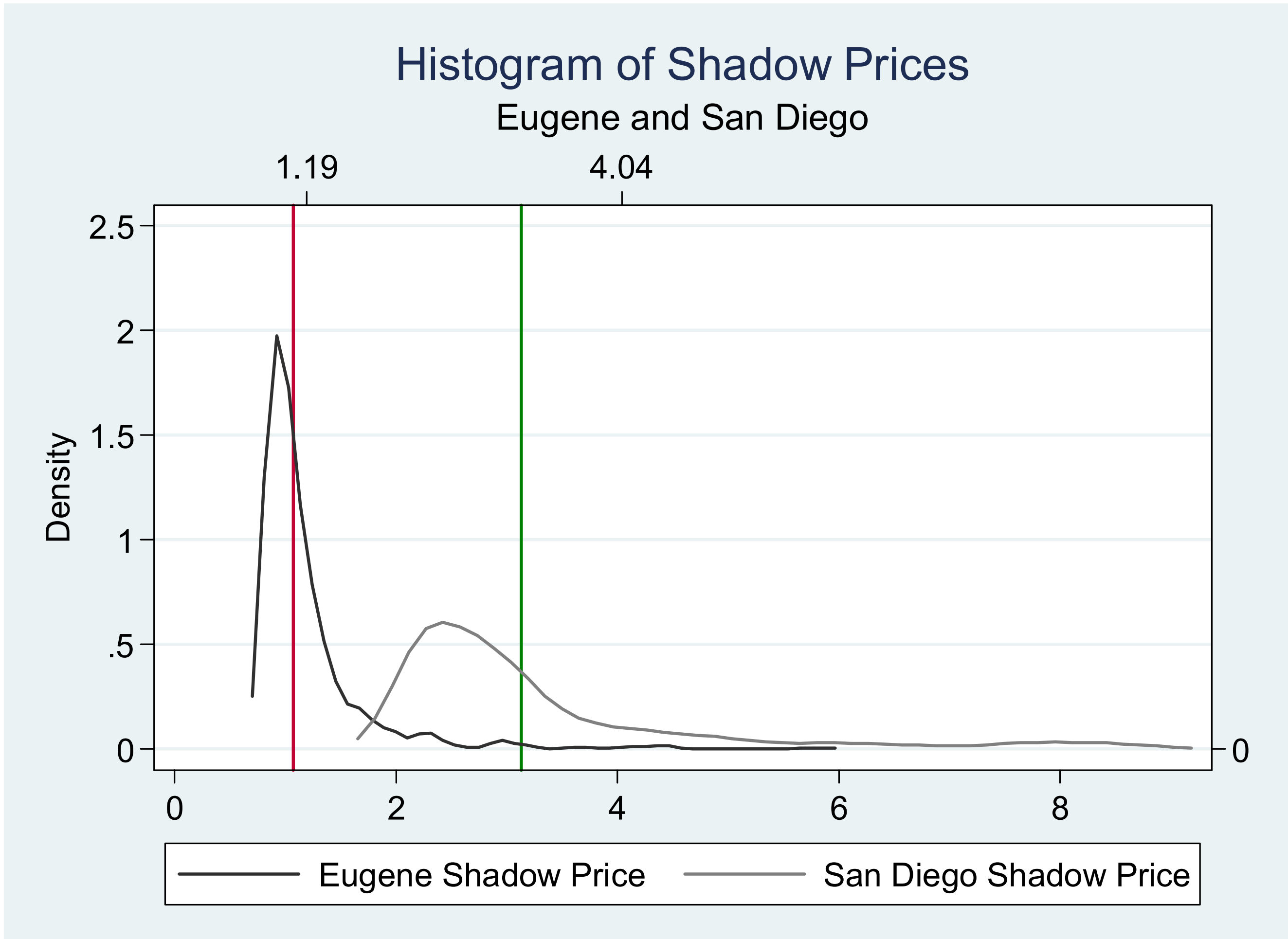


Figure 2. Stylistic Model of a Market-Clearing Price, with Shadow Prices



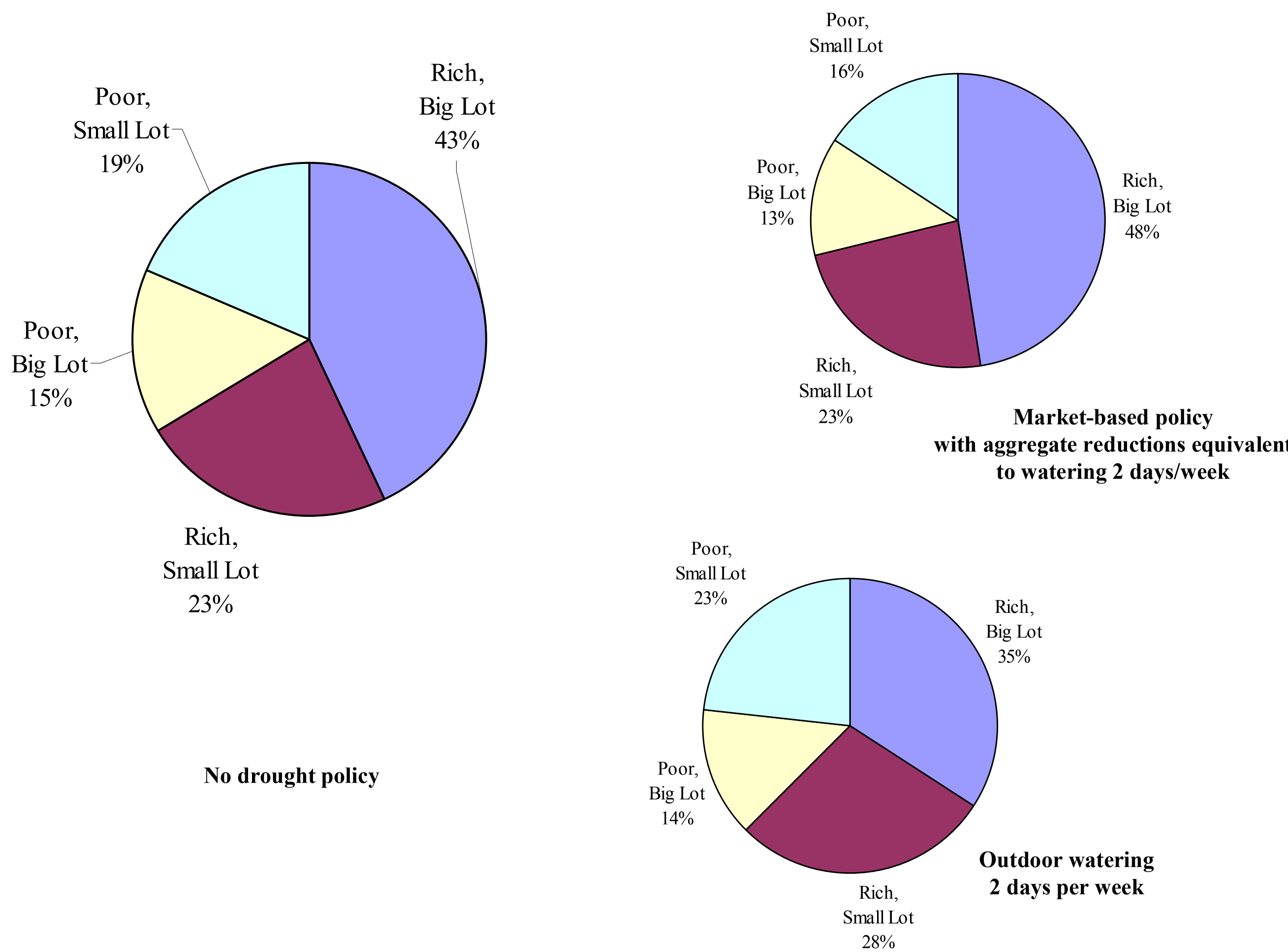
(Where P^* is the market-clearing price for $Q^{reg}_{psl} + Q^{reg}_{rbl} + Q_{indoor} = Q^*_{psl} + Q^*_{rbl} + Q^*_{indoor}$).

Figure 3. Distribution of Shadow Prices



Notes: Distributions are of estimated shadow prices for Eugene, Oregon and San Diego, California, given a two days per week water policy. Numbers at the top of the figure identify average shadow prices for each city, and vertical lines represent the alternative policy of a market-clearing price, in dollars per thousand gallons.

Figure 4: Allocation of Water Consumption by Drought Policy



Notes: Total irrigation season consumption (the “size of the pie”) is 6,815 kgal/day with no drought policy, and 4,632 kgal/day with either drought policy.

Appendix A. Estimation of Price Instruments

The water demand function (A.1) is in exponential form, where w is total daily water demand, Z is a matrix of seasonal and daily weather conditions, X is a matrix of household characteristics, p is the marginal water price, $\widetilde{Y}$ is virtual income, η is a measure of household heterogeneity, ε is optimization or perception error; and δ, β, α, and μ are parameters.[29]

$$w = e^{Z\delta} e^{X\beta} p^{\alpha} \widetilde{Y}^{\mu} e^{\eta} e^{\varepsilon} \tag{A.1}$$

Let $\underline{w}_k^*(.) = e^{Z\delta} e^{X\beta} p_k^{\alpha} \widetilde{Y}_k^{\mu}$, or optimal consumption in block k. Then, conditional demand under a two-tier increasing-block price structure, in which w_1 is the kink point, can be represented as in (A.2), and conditional price as in (A.3).

$$w = \begin{cases} \underline{w}_1^*(.)e^{\eta}e^{\varepsilon} & if \quad 0 < e^{\eta} \le \dfrac{w_1}{\underline{w}_1^*(.)} \\ w_1 e^{\varepsilon} & if \quad \dfrac{w_1}{\underline{w}_1^*(.)} < e^{\eta} \le \dfrac{w_1}{\underline{w}_2^*(.)} \\ \underline{w}_2^*(.)e^{\eta}e^{\varepsilon} & if \quad \dfrac{w_1}{\underline{w}_2^*(.)} < e^{\eta} \end{cases} \tag{A.2}$$

$$p = \begin{cases} p_1 & if \quad 0 < e^{\eta} \le \dfrac{w_1}{\underline{w}_1^*(.)} \\ \text{indet.} & if \quad \dfrac{w_1}{\underline{w}_1^*(.)} < e^{\eta} \le \dfrac{w_1}{\underline{w}_2^*(.)} \\ p_2 & if \quad \dfrac{w_1}{\underline{w}_2^*(.)} < e^{\eta} \end{cases} \tag{A.3}$$

Consumption only occurs at the kink point if the consumer maximizes utility for choices that are unavailable at all (p_k, Y_k), so for kink observations, $\underline{w}_1^*(.) > w_1$ and $\underline{w}_2^*(.) < w_1$ (see Figure A.1).

---

[29] The structural model includes two additional parameters, σ_η and σ_ε. Our 2SLS approach does not allow separate identification of the two error variances. We use the structural estimate of σ_η to calculate block and kink probabilities in (A.5).

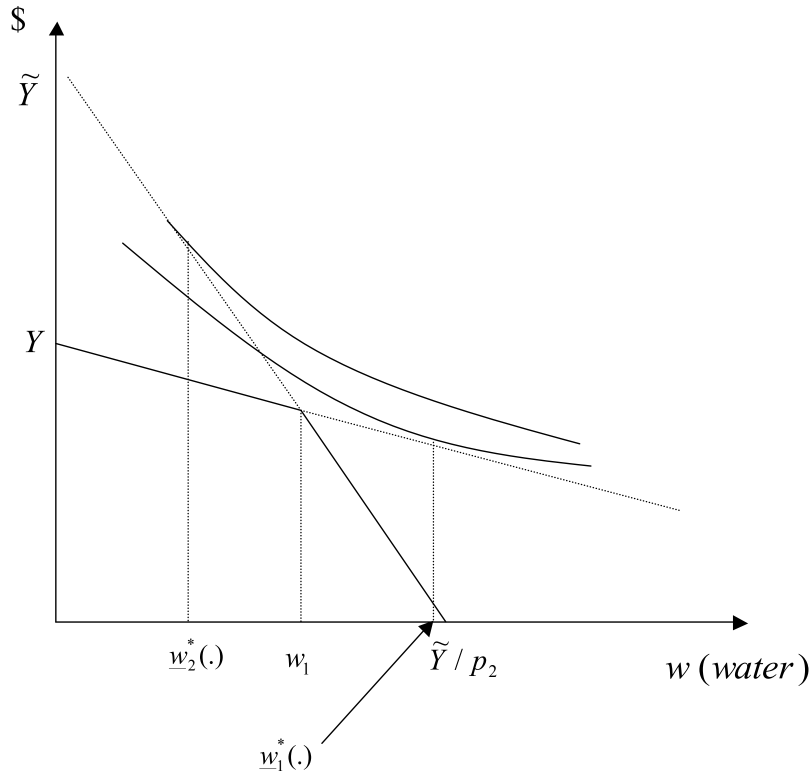


Figure A.1. Preferences resulting in consumption at a kink point under a two-tier increasing block price structure

From the conditional price equation, we derive a daily probability-weighted price (A.4). Our price instrument is the seasonal average, by household, of $\widehat{p}$. Errors are assumed to be independent and normally distributed. Thus, $e^{\eta} \sim LN(\mu_{e^{\eta}}, \sigma_{e^{\eta}})$, and integrations in (A.5) are over the probability density function of this lognormal distribution.

$$\widehat{p} = \Pr A * p_1 + \Pr B * (.5 p_1 + .5 p_2) + \Pr C * p_2 \tag{A.4}$$

Where:

$$\Pr A = \int_0^{\frac{w_1}{\underline{w}_1^*(.)}} f(e^{\eta}) de^{\eta}$$

$$\Pr C = \int_{\frac{w_1}{\underline{w}_2^*(.)}}^{\infty} f(e^{\eta}) de^{\eta} \tag{A.5}$$

$$and \quad \Pr B = 1 - \Pr A - \Pr C$$

Appendix B. Robustness of Estimation

The ideal data for this analysis would include a longer time-series component. Indeed, with only two price observations per household, these are hardly panel data at all (at least along the dimension of prices). For this reason, we cannot estimate a model with household fixed effects (FEs), or even city FEs, without reducing price variation so substantially as to prevent reasonable interpretation of parameter estimates, provided we are able to estimate effects, at all, that are significantly different from zero (recall that about 92 percent of variation in sample prices can be explained by household FEs, alone).

The best way we can control for household heterogeneity in this context without losing too many degrees of freedom is to include a household random effect in the models, as we have done. A Hausman test rejects the null hypothesis that the random effects model is consistent and efficient for some, but not all of our models. Where we reject random effects, we do so largely because the estimates are more precise than they are in the models in which the Hausman test does not reject the null.

As a test of robustness, we do estimate models with household and city FEs. Table B.1 reports results from these robustness checks, as well as others to be described in the paragraphs that follow. The first column of Table B.1 reports results from our random-effects demand models (originally reported in Tables 2 and 3), for the purpose of comparison.

The household fixed effects (FE) model generates a total demand elasticity similar to that of our basic model. The indoor demand function is upward-sloping, though the price coefficient is insignificant; both are likely artifacts of the small amount of price variation available to estimate the indoor elasticity. The incidental parameters problem prevents us from estimating a maximum likelihood outdoor model with household FEs. In addition to losing degrees of freedom, we lose descriptive power through our inability to estimate the effects of individual household characteristics on water demand.

When we estimate daily water demand models with city FEs, the elasticity of total consumption is statistically insignificant, indoor demand is upward-sloping (though the elasticity is also insignificant), and the magnitude of the outdoor elasticity is about one-half the magnitude of the outdoor elasticity with regional FEs (see Table B.1). For this sample, the inclusion of city FEs substantially reduces price variation and results in noisy estimates of price elasticities. Fixed effects of an even finer level, that of utilities, might control for residential water conservation programs—and other utility policies—unlike either the regional or city FEs. The inclusion of regional fixed effects does control for long-run climate variation but cannot capture utility-specific heterogeneity.

We test the robustness of our estimates by constructing one additional alternative model. We collapse daily observations to seasonal observations, creating two demand observations per household, and obtain estimates from regressions of aggregate seasonal household demand on the independent variables (reported in Table B.1). The aggregate demand models of total, indoor, and outdoor consumption provide estimates that are very similar to their daily counterparts. That our models are robust to this seasonal specification is encouraging. However, the daily water demand models are preferable in that they provide more detailed information regarding the impact of daily weather conditions on water consumption.

Table B.1. Robustness of Elasticity Estimates

Elasticity of	Overall	Household FEs	City FEs	Aggregate consumption
total consumption	-0.3508* (0.0683)	-0.2936* (0.1440)	-0.1759 (0.1090)	-0.3997* (0.0691)
indoor consumption	-0.0713 (0.0567)	0.0562 (0.1117)	0.0884 (0.0886)	-0.0673 (0.0612)
outdoor consumption	-0.6430* (0.0687)	.	-0.3270 * (0.1284)	-0.5794* (0.1010)

Notes: * significant at 5% (# at 10%).

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