**Estimating the Price Elasticity of Demand for Water with Quasi Experimental Methods**

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## I. Introduction

There is a growing recognition in both the professional and popular literatures that water scarcity is a key policy issue which is essential to address in evaluating the effects of climate change and long term sustainability of economic growth.[1] Glennon's [2009] observations in his new book, Unquenchable: America's Water Crisis and What to Do About It, describe the problem well:

> "Water presents a surprising riddle. We can neither make nor destroy it, so our supply is fixed yet it's exhaustible because, as a shared resource used repeatedly, some uses preclude future reuse. Water policy suffers from a profound discontinuity between science and law…the result epitomizes the tragedy of the commons: limitless access to a finite resource" (p. 324)

Those evaluating the water problem usually conclude prices must be reformed so that incentives facing water users change to reflect this scarcity. Demand functions provide the basic economic relationships required to understand how water use will respond to such changes. Here is where the problems arise. Access to detailed information on household water use including household attributes (i.e. lot size,

[1] See Covich [2009] for a detailed assessment of the climate change and water challenges.

landscaping composition, and presence of swimming pools), the composition of use (i.e. indoor versus outdoor usage), and pricing is limited. Where there is metering and pricing of water there has also been growing use of increasing block pricing structures. These pricing policies transform the methods required to use the records on households' water consumption patterns for estimating how the quantity of water used responds to price changes.[2] They also make the requirements for detailed information on household attributes even more important. Finally, due to these data limitations it has been difficult to characterize the differences in water demand responses by small versus large residential users. This shortcoming has direct policy relevance because there is broad consensus in the U.S. that affordable access to water for "ordinary" household use must be maintained for lower income groups.

This paper proposes a simple method for estimating the price elasticity of demand that meets policy needs and can accommodate current data limitations as well as the presence of increasing block pricing structures. In short, we propose a method that can estimate how water usage responds to price changes under the typical conditions confronting applied researchers of what we label as "under data duress." Our strategy uses the seasonal and temporal changes in the Phoenix municipal water system's residential water rates to reconstruct a record of price changes that are exogenously imposed on customers. By matching months that experienced changes in the marginal price for the same consumption block it is possible to isolate the quantity adjustment associated with these specific price changes and be assured there is no shifting of

[2] See Olmstead et al. [2007] for a derivation of the relationship between conditional and unconditional price elasticities.

households between price blocks[3]. Our sample was constructed using percentiles of water consumption at the level of the census block group. This approach allows the analysis to control for housing attributes, external water usage, and socio-economic characteristics of consumers. The findings from our simple strategy are striking. Despite limited price variation we estimate statistically significant demand elasticities that are more inelastic for large users. This approach was applied to two pairs of years where the marginal prices changed, one involved a price change in a pair of "normal" years (in terms of precipitation) and a second had a price change between the same normal base year and a different "treatment" year that had exceptionally dry conditions. Thus, a comparison of the results for approximately comparable price changes across these distinctive set of natural precipitation conditions provides a simple plausibility check for the logic of the model. That is, we should expect natural increases in the need for water (due to dry conditions) would make households with significant outside uses of water less responsive to price increases. This is exactly what we find. The estimates of price elasticities for residential customers of all size classes are much less responsive to the price changes when the price comparison involves responses during dry conditions. In both the normal / normal and normal / dry case we estimate separate models for winter and summer demand.

Section two outlines our model and discusses its relationship to the conditional demand approach widely advocated for estimating price elasticities with increasing block

[3] This strategy is consistent with Chetty's [2009] call for a middle ground between structural and reduced form (or treatment effect) approaches to policy evaluation. He describes how research in public economics has focused on what he labels a "sufficient statistic" or estimation strategies that focus on measuring parameters in transparent ways with credible identifying information and yet focus on the key *economic* parameter for policy evaluation and often welfare statements.

prices.[4] The third section describes our data and empirical estimates. We conclude with comments on the applicability of our new approach in other situations where response of water usage to price changes must be evaluated with limited information.

## II. Modeling Water Demand

### A. Background

Most of the recent literature measuring the price elasticity of demand has been based on the discrete / continuous choice (DCC) model developed by Hausman [1979] for applications to labor supply.[5] For the case of demand models derived based on choices with convex budget sets (increasing block price structures), one need only specify a conditional demand function to estimate the model. This relationship describes how the quantity demanded responds to the marginal price within each budget segment (corresponding to the segment associated with each step in the increasing block pricing structure). Conventional practice assumes this demand function includes two errors. One is usually hypothesized to be associated with household preference heterogeneity not captured in observable variables. This feature is known to the household but not the analyst. The second is an error hypothesized to be unknown to both the household and the analyst. It could arise from leaks in the water system, measurement errors, or a composite

[4] An alternative structure based on using the first order conditions from the constrained utility maximization problem is developed and illustrated in Strong and Smith [forthcoming].

[5] See Burtless and Hausman [1978] for discussion of the model with non-convex budget constraint, Heckman [1983] for a critique and Reiss and White [2006] for discussion of how welfare analysis can be undertaken with non-linear budget constraints.

of both effects. The two error terms allow the model to describe choice outcomes, including some households' consumption choices that would appear to have them falling at the kinks of the budget constraint.

As a rule, the decision process is explained as if there were two distinct steps: (a) description of the probability consumption will be in one of the budget segments (or at a kink)[6] and (b) conditional on a budget segment there is a conditional demand function describing how the quantity of water demanded relates to the marginal price[7]. At a kink, there is no such relationship because the model implies demand is higher than the highest value of the lower block but lower than the lowest value of the next highest block. The parameters of the conditional demand are assumed constant across segments to identify the model.[8] This assumed constant parameter structure for all budget segments does not simplify the relationship between conditional and unconditional price elasticities.

As Olmstead et al. [2007] demonstrate, the unconditional price elasticity is a function of both the conditional price elasticity and the income elasticity. As a result, the relationship of the unconditional price elasticity to the conditional elasticity is not clear. Both the ranking of the two measures and the magnitude of the absolute difference between them cannot be signed *a priori*. For their application Olmstead et al. found the unconditional price elasticity of demand was smaller in absolute magnitude than the conditional elasticity. Because the rate subsidies associated with increasing block price

[6] There will be multiple kinks if there are several steps in the increasing block structure.
[7] In practice we simply observe households consumption that implies their consumption affects the marginal price they face. The steps are used to separate the way the statistical model uses available information to treat the inherent simultaneity consistently.
[8] See Strong and Smith [forthcoming] for a discussion of the limitations of this assumption.

structures are small, the price differences between blocks are small, and the expenditure share for water in a household's overall budget is small; the authors suggest their results may provide a reasonable guide for the relationship with most communities' residential water demands.

Nonetheless, despite this relatively optimistic conclusion of the Olmstead et al study, the experience with the DCC model for water applications has been mixed. Hewitt and Hanemann [1995] report the first such application. They found large (in absolute value) conditional price elasticities and report an unconditional elasticity for a price structure change that is also large compared to the literature. Both the Espey et al. [1997] and the Dalhausen et al. [2003] meta analyses of water demand studies found the majority of the estimates for the price elasticity were less than unity in absolute magnitude. Olmstead et al.'s results are consistent with these findings in that their conditional elasticity estimate was approximately -0.34 and simulated unconditional estimates was -0.59.

The Olmstead et al. sample is quite unique in that it pools household level data across 11 urban areas in the U.S.[9] One concern the authors raise is that the very advantage of their data in displaying how households responded to a variety of price structures may create a cause for concern. That is, the community's tastes for water conservation may be reflected in the provider's rate structure. In this case, the selection of a demand function that is defined conditional to a rate structure may be reflecting this unobserved taste for conservation rather than offering an exogenous source of variation in price structures necessary to recover how consumers water consumption changes with

[9] The data are relatively old as are the data underlying Hewitt and Hanemann [1995].

them. They conclude, after an interesting set of comparative assessments of alternative models, to address this question, by observing that they cannot dismiss the hypothesis that the underlying city level taste for conservation, and thus the price structure, may contribute to the observed higher (in absolute value) elasticities. It is not possible to reproduce their study using a single water provider because most water systems' price schedules contain insufficient variability in price schedules for identification.[10] Moreover, the meter level records of water usage contain no information on the household characteristics. These data are important to being able to distinguish the selection of consumption blocks from the amount of use given a block has been chosen. Overall, the composite of the challenges arising from rate structures intended to encourage water conservation together with exceptionally limited detail about the factors that might distinguish different households' water use, aside from price, provide the motivations for our new approach.

B. A Quasi-Experimental Approach for Measuring Price Elasticities for Residential Water Demand

Our analysis considers a single municipal water provider with a two block pricing structure. Rates change in different months throughout the year corresponding to low, medium, and high usage periods. Water and waste water are priced on the same metered

[10] One exception to this is Pint (1999) who estimates the DCC model for the Alameda County Water District but argues that it would be difficult to take the results from the estimation to other parts of California since the results are sensitive to climatic conditions, a key point of our analysis. Additionally, she has good resolution of household characteristics including house size and lot size.

records. Thus, changes in either price affect incentives for water usage. The block cutoff for the increase in marginal price changes from 600 cubic feet in low water use periods (winter) to 1000 cubic feet in high water use periods (summer). Our estimation strategy exploits increases in rates over time that vary by different amounts for each usage period along with the ability to map meters to census block groups. To consider the first component of our analysis, the changes in marginal prices, we use a linear specification for water demand as in equation (1).

$$w_{ijt} = \alpha_0 + \alpha_1 p_{jtb} + f(z_i) + g(T_{jt}) + u_{ijt} \tag{1}$$

$w_{ijt}$ is the water consumed by household *i* in the month *j* of year *t*. Based on the metered records for $w_{ijt}$, we know $p_{jtb}$, which is the *b*th block's marginal price (water and waste water) in month *j* of year *t*. This is the block that corresponds to the water consumption $w_{ijt}$. $f(z_i)$ is a function for the effects of observable household attributes ($z_i$) and $g(T_{jt})$ is the effect of weather related variables – temperatures and precipitation in month *j* and year *t*. $u_{ijt}$ is an error that can be a composite of unobserved individual heterogeneity and measurement error.

By selecting years that experienced marginal price changes for at least one of the three periods and considering the households in a given price block we can difference equation (1) between two years in which a marginal price change occurs to derive equation (2).

$$w_{ijt+a} - w_{ijt} = \alpha_1 \bullet k + (g(T_{jt+a)} - g(T_{jt}) + (u_{ijt+a} - u_{ijt}) \tag{2}$$

$k$ is the constant increment to the marginal price experienced between years. Using this specification the intercept in this difference equation estimates the effect of price on water demand assuming the price change between the two months is identical for each consumer. These differences can be pooled over months if the price increment is the same. In addition, if the price block remains constant we do not need to consider the issues of a household adjusting to price changes by altering its consumption block. This type of change would be a key requirement to use the DCC model. Of course, if we imposed a restriction that we would only consider households who did not change blocks we do not avoid the problem of endogeneity as that restriction would create a selection effect. This effect occurs because we would be leaving out the individuals who changed blocks and failed to account for how the features of those who did not are likely to change their demand responses to prices.

Fortunately, these issues can be avoided by focusing on the distribution of responses to price changes. Following Borenstein's [2009] proposal for the case of modeling electricity demand, we estimate price responses using percentile thresholds of water consumption for the 10, 25, 50, 75, and 90 percentiles by month for households grouped into census block groups. By constructing differences in each block group for the same percentile across years with different prices we difference out cross-sectional neighborhood differences, such as xeric or mesic landscape, that might interact with the consumption responses to marginal price changes. This strategy together with the Phoenix pricing structure results in only one of the threshold quantities displaying price responses that is sufficient to lead consumption to be in a different price block -- the tenth percentile. For this percentile, the change in blocks is limited to one month, May, which precedes the high use and highest price season. All of the remaining points on the distribution fall in the highest block. We avoid these issues by estimating demand for the winter (low) and summer (high) periods which do not include the month of May.

Thus, for this application, with only two price segments and a relatively low quantity threshold for the highest block, there is no scope for differences between conditional and unconditional elasticities. As a result, we cannot comment on the importance of the issue for the design of price structures. Our strategy does allow us to estimate different conditional price elasticities for different sized residential users. To our knowledge, this issue has not been considered in the past literature. It appears these distinctions can be important for pricing policy when there are concerns with assuring affordable access to low income users who tend to be in the lowest use percentile.

## III. Data and Price Elasticity Estimates

Our analysis is based on records for residential meters served by the municipal Phoenix water system. Aside from the year, month, amount of water used and the price schedule each user faced we have no other information about these households. It is only possible to match the meter records to one of approximately 1000 census block groups served by the Phoenix water system When we match the water meter data and information on weather conditions experienced in each block group for each month and year we are left with between 993 and 996 records for each month in the three years comprising our sample.

As noted earlier our analysis relies on isolating changes in the marginal prices for water over time. These changes are not associated with any household moving between pricing blocks. Table 1 displays the marginal prices for the highest water usage blocks by month for the three years in our sample. There are some months that display differences from the low (December, January, February, March), medium (April, May, October, November), and high (June, July, August, September) periods due to the timing of allowed adjustments in the rate structure.

Our analysis uses these monthly differences in prices for the low and the high usage period as exogenous price differences to form a quasi-experiment. By considering the difference in the quantity thresholds for each consumption percentile at the block group level, we are differencing away the effects of the housing distribution and neighborhood attributes to the extent they are relatively constant at the block group level. Given that the bulk of the new housing construction in the Phoenix area during this time period occurred outside the Phoenix water providers district, this assumption appears reasonable.

As Table 1 suggests, we selected three years for analysis and treat the year 2000 as our base year. This strategy provides two different sets of price changes to evaluate using differences between the years 2000 versus 2002 and 2000 versus 2003. The years 2000 and 2003 were considered a pair of normal years in terms of precipitation while year 2002 was unusually dry. We explore the effects of these climate differences on our elasticity estimates later on. Table 2 compares the weather variables and percentiles of consumption use for each of the three years.

Our models also include controls for monthly differences in the minimum temperature, the precipitation, and the days of precipitation interpolated for each block group. This data came from approximately 15 NOAA monitoring stations located around the Phoenix area which were collecting data during the months used for our analysis. Each monitoring station reports daily data on temperatures and precipitation. To form the climate variables which we hypothesize to be important in determining water consumption, we averaged temperatures in each month at each monitor to form a series of monthly specific temperature variables. For precipitation, we formed monthly counts of the total number of days in which rainfall was detected as well as the total amount of rainfall (measured in inches). Each block group in our sample is assigned the temperature data from the closest monitor.

As Table 1 demonstrates, different months experience different price changes. It is essential to focus on months where a constant price change was experienced by consumers. Comparing the price changes over all months in the 2000 and 2003 difference we observe price changes varying from \$0.14 to \$0.31 (per 100 cubic feet). Eight months fall between the ranges of \$0.14 to \$0.17. Comparing 2000 and 2002 the range is \$0.08 to \$0.19.

We estimated the model for each of two sub-samples – restricting the data to include only winter (low-December, January, and February) and then only summer (high-June, July, August, and September).

Table 3 reports our estimated first difference model for these two seasons and for each paired price change (2000-2003 and 2000-2002). The intercept in each set of models is our key focus. This parameter is precisely estimated in comparing summer consumption for normal weather years. For winter use, large users are not responsive to price but otherwise respond during summer months. What is especially striking is the contrast when we consider a price change comparing a normal and a dry year. Here all users are price responsive but the pattern of adjustment is informative. In the summer months all users, but especially large users, are less responsive to price changes during dry conditions. This suggests that water users exercise savings in the winter months – price changes when less external water is needed are largest for the large users. In comparing a dry and normal year changes in precipitation are the same sign in both winter and summer months. The amount of precipitation, as Table 2 shows, has declined so the change is negative between 2002 and 2000. This precipitation effect increases demand in both winter and summer. By contrast, in comparing the two normal years the effect depends upon season. To assess the overall impact we need to consider changes in the amount of precipitation and its distribution over days. Changes in each of these variables have opposite signs for the summer and winter models. For the comparison of the normal and dry year the two effects are consistent between the winter and summer models. Reductions in precipitation lead to increases in water use and reductions in the number of days of precipitation reduce water usage for both seasons. Precipitation declines in the winter increase water use, while in the summer the effect depends on both the change in amount and days of precipitation. All the T-tests for testing the effects of individual variables use Huber robust standard errors.

Table 4 presents the primary results of our analysis – estimates of the demand elasticity using comparisons between two normal years versus a normal and a dry year to identify the price effect. Our estimates using percentiles can distinguish the size of residential customers and appear to play an important role in parsing out the heterogeneity in household level estimates for the price elasticities.

Comparing the first panel of table 4 for normal / normal versus the second panel for normal / dry we see directly that price responsiveness is reduced quite substantially for the summer when the overall situation is abnormally dry. All classes of users (as reflected by the consumption percentiles) respond to prices in the winter when it is dry. For normal years, there is greater responsiveness to price among all users in the summer than for the case of the dry year. Also, in the winter, we see large users show no response to prices in the summer. To our knowledge this is the first time a water demand study has been able to extract these distinctions and control for the effects of the block structure. Overall, households appear to economize in the winter months and are more responsive to price in winter months compared with summer months because outdoor uses (vegetation and pools) are less. Moreover, dry conditions do reduce price responsiveness.

## IV. Implications

When the DCC approach was first proposed to model labor supply and evaluate the effect of policies to change income taxes Heckman [1983] raised a number of questions about the ability of analysts to assume the actual points specifying the kinks in the budget constraint, due to tax structure, were exogenous. In addition, he noted that with exogenous kink points there should be "bunching at the kinks."[11] These kink points are exogenous in the case of increasing block price structures for water. If the behavior

[11] See Heckman [1983] p. 71.

hypothesized in the DCC model is important we should observe this bunching. In preliminary analysis of the Phoenix residential water use records we considered this issue and the results were not convincing. This finding should not be surprising given the rate structure facing households. Most residential customers are in the highest block throughout the year. As a result, there would be no meaningful variation in the marginal price for these customers. Even the seasonal changes in rates would be experienced at the same time and responses to the price changes would be difficult to separate from changes in water needs due to the temperature changes throughout the year. The Phoenix system's price structure is not unique. Thus it may be reasonable to assume the opportunities for adjustments among blocks for many water systems are limited and the distinction between conditional and unconditional price responses small.

Moreover, as Olmstead et al. suggest, using situations where there are large differences in price structures to indentify price responses within cross community studies must consider whether the price structure is truly exogenous. Community conservation motives may lead to conservation oriented pricing structures and therefore it may be difficult to assume the variation in price structures in a cross community analysis is exogenous to preferences. This concern would lead to preference for models that do not start with the conditional demand function defined for a given price structure and instead a framework that begins from a preference specification. Of course, this approach adds considerable structure in the assumed specification for preferences.

We have proposed a different strategy for estimating demand responses more consistent with the compromise advocated by Chetty [2009]. It combines two elements. The first is Borenstein's proposal to consider how order statistics summarizing the distribution of use, constructed at the census block group level, change with exogenous price changes. The second element involves exploiting changes in marginal prices over time (to reflect cost increases) that can be matched by month so the position of water consumption in the price structure is unchanged. This matching process assures the increment to the marginal price can be used as an exogenous change from the household's perspective. As a result, the price change can be treated as a type of treatment for evaluating the change in water usage.

Our application to the Phoenix residential market indicates this approach is remarkably effective. It does not require the structural assumptions of the DCC approach, allows estimates of price responsiveness by size (in terms of water use) of customer, and was very effective at detecting how lower than normal precipitation in an arid desert environment can reduce price responsiveness, especially for the large users. We interpret this ability to detect a role for differences in seasonal rainfall as confirming evidence that supports our direct approach. It confirms that matching a simple theoretical model of demand response with an exogenous source of price variation can allow measurement of demand elasticities. Our use of the order statistic to characterize demand provides a strategy that recognizes the importance of a price change to any individual household will depend on the amount of water used. More complex increasing block price schedules can be accommodated in this logic with enough ability to observe within block changes in marginal prices over time.

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**Table 1: Monthly Variation in Marginal Prices for the Highest Water Blocks: 2000, 2002, 2003**

| | Year | | | | | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | 2000 | | | | 2002 | | | | 2003 | | | |
| | Marginal Prices | | | | | | | | | | | |
| Month | 1.09 | 1.12 | 1.32 | 1.68 | 1.17 | 1.24 | 1.47 | 1.87 | 1.24 | 1.26 | 1.49 | 1.89 |
| January | X | | | | X | | | | X | | | |
| February | X | | | | X | | | | X | | | |
| March | X | | | | | X | | | | X | | |
| April | | | X | | | | X | | | | X | |
| May | | | X | | | | X | | | | X | |
| June | | | | X | | | | X | | | | X |
| July | | | | X | | | | X | | | | X |
| August | | | | X | | | | X | | | | X |
| September | | | | X | | | | X | | | | X |
| October | | | X | | | | X | | | | X | |
| November | | | X | | | | X | | | | X | |
| December | | X | | | | X | | | | X | | |

**Table 2: Comparison of Weather Conditions in Normal Versus Dry Years**[a]

| | Year / Classification | | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|
| | 2000 / Normal | | | 2002 / Dry | | | 2003 / Normal | | |
| Water Usage Quantile | *m* | σ | | *m* | σ | | *m* | σ | |
| 10% | 6.36 | 3.53 | | 6.22 | 3.33 | | 5.76 | 3.4 | |
| 25% | 10.24 | 4.92 | | 10.07 | 4.62 | | 9.42 | 4.67 | |
| 50% | 16.11 | 7.33 | | 10.07 | 4.62 | | 9.42 | 4.67 | |
| 75% | 23.99 | 2-Nov | | 23.95 | 11.04 | | 22.68 | 11.33 | |
| 90% | 33.66 | 16.21 | | 33.71 | 16.29 | | 32.09 | 16.68 | |
| Weather[b] | *m* | Min | Max | *m* | Min | Max | *m* | Min | Max |
| Precipitation | 0.65 | 0 | 4.86 | 0.3 | 0 | 1.78 | 0.75 | 0.1 | 5.07 |
| Precipitation Days | 2.5 | 0 | 11 | 1.78 | 0 | 6 | 2.78 | 0 | 9 |

[a] m considers to the sample mean and σ the standard deviation; Water is measured in units (100 cubic feet)

[b] Precipitation is in inches and precipitation days are a count of days with measurable precipitation at the linked NOAA weather station

**Table 3: Water Consumption Difference Models: Normal Years and Normal / Dry Year**

| Percentile | Season | Intercept | Δ Min Temperature | Δ Precip | Δ Precip Days | $R^2$ | N |
|---|---|---|---|---|---|---|---|
| | | 2003-2000 (Normal / Normal ) | | | | | |
| 10% | Summer | -0.826 | 0.135 | 0.342 | -0.144 | 0.014 | 398 |
| | | (-15.22) | -5.95 | -4.61 | (-5.54) | 0.07 | 2 |
| | Winter | -0.269 | -0.114 | -0.498 | 0.194 | | 298 |
| | | (-3.90) | (-8.49) | (-13.94) | -6.76 | | 7 |
| 25% | Summer | -1.18 | 0.207 | 0.481 | -0.195 | 0.025 | 398 |
| | | (-20.34) | -8.42 | -5.74 | (-7.03) | 0.121 | 2 |
| | Winter | -0.175 | -0.181 | -0.689 | ..231 | | 298 |
| | | (-2.17) | (-11.02) | (-15.99) | -6.9 | | 7 |
| 50% | Summer | -1.518 | 0.309 | 0.689 | -0.265 | 0.038 | 398 |
| | | (-22.25) | -9.98 | -6.78 | (-8.24) | 0.147 | 2 |
| | Winter | -0.077 | -0.295 | -0.989 | 0.319 | | 298 |
| | | (-0.71) | (-12.55) | (-14.77) | -6.45 | | 7 |
| 75% | Summer | -1.885 | 0.454 | 1.252 | -0.382 | 0.041 | 398 |
| | | (-19.42) | -10.3 | -8.51 | (-8.43) | 0.153 | 2 |
| | Winter | -0.144 | -0.436 | -1.484 | 0.512 | | 298 |
| | | (-0.90) | (-12.45) | (-16.59) | -7.74 | | 7 |
| 90% | Summer | -2.160 | 0.523 | 1.641 | -0.415 | .027 | 398 |
| | | (-14.94) | -8.02 | -7.93 | (-6.25) | 0.094 | 2 |
| | Winter | 0.052 | -0.659 | -2.01 | 0.661 | | 298 |
| | | -0.21 | (-11.89) | (-12.99) | -6.26 | | 7 |
| | | 2002-2000 (Normal / Dry) | | | | | |
| Percentile | Season | Intercept | Δ Min Temperature | Δ Precip | Δ Precip Days | $R^2$ | N |
| 10% | Summer | -0.28 | 0.118 | -0.088 | 0.035 | 0.006 | 398 |
| | | (-4.54) | -5 | (-0.78) | -1.2 | 0.03 | 2 |
| | Winter | -0.352 | 0.025 | -0.727 | .260 | | 298 |
| | | (-7.92) | -1.57 | (-2.18) | -5.99 | | 8 |
| 25% | Summer | -0.427 | 0.181 | -0.137 | 0.075 | 0.013 | 398 |
| | | (-6.28) | -6.86 | (-1.15) | -2.34 | 0.052 | 2 |
| | Winter | -0.461 | 0.023 | -1.494 | 0.404 | | 298 |
| | | (-10.03) | -1.35 | (-4.09) | -8.83 | | 8 |
| 50% | Summer | -0.625 | 0.226 | -0.228 | 0.134 | 0.018 | 398 |
| | | (-7.87) | -6.83 | (-1.64) | -3.63 | 0.07 | 2 |
| | Winter | -0.598 | 0.032 | -1.356 | 0.531 | | 298 |
| | | (-11.05) | -1.37 | (-2.77) | -8.6 | | 8 |
| 75% | Summer | -0.597 | 0.332 | -0.243 | 0.148 | 0.015 | 398 |
| | | (-4.71) | -6.91 | (-1.12) | -2.52 | 0.068 | 2 |
| | Winter | -0.751 | 0.031 | -1.983 | 0.75 | | 298 |
| | | (-9.67) | -1.04 | (-2.83) | -8.6 | | 8 |
| 90% | Summer | -0.595 | 0.502 | -0.541 | 0.14 | 0.011 | 398 |
| | | (-2.99) | -6.23 | (-1.60) | -1.56 | 0.025 | 2 |
| | Winter | -1.047 | -0.025 | -5.561 | 1.238 | | 298 |
| | | (-6.27) | (-0.33) | (-1.89) | -5.1 | | 8 |

**Table 4: Price Elasticity for Residential Water Demand[a]**

| | 2003-2000 (Normal / Normal) | | | 2002-2000 (Normal / Dry) | | |
|---|---|---|---|---|---|---|
| Percentile | Overall | Winter | Summer | Overall | Winter | Summer |
| 10 | -1.068 | -0.528 | -0.959 | -0.296 | -0.758 | -0.362 |
| | (-27.78) | (-3.9) | (-15.22) | (-7.37) | (-7.92) | (-4.54) |
| 25 | -0.899 | -0.215 | -0.823 | -0.143 | -0.627 | -0.335 |
| | (-37.19) | (-2.17) | (-20.34) | (-5.54) | (-10.03) | (-6.28) |
| 50 | -0.743 | -0.061 | -0.652 | -0.99 | -0.524 | -0.307 |
| | (-40.13) | (-0.71) | (-22.25) | (-5.16) | (-11.05) | -7.87 |
| 75 | -0.625 | -0.075 | -0.537 | -0.003 | -0.438 | -0.195 |
| | (-35.21) | (-0.91) | (-19.42) | (-0.15) | (-9.67) | (-4.71) |
| 90 | -0.528 | * | -0.437 | * | -0.428 | -0.138 |
| | (-27.38) | | (-14.94) | | (-6.27) | (-2.99) |

[a]The numbers in parentheses are asymptotic Z statistics, treating the price difference, price and quantity at their sample means as constants for estimating the variance of the estimated price elasticity.

[*]Positive and statistically insignificant