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Regulating Irrigation via Block-Rate Pricing: An Econometric Analysis

by

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Abstract

In this paper we adapt Burtless and Hausman's (1978) methodology in order to estimate farmers' demand for irrigation water under increasing block-rate tariffs and empirically assess its effect on aggregate demand and inter-farm allocation efficiency. This methodology overcomes the technical challenges raised by increasing block rate pricing and accounts for both observed and unobserved technological heterogeneity among farmers. Employing a micro panel data documenting irrigation levels and prices in 185 Israeli agricultural communities in the period 1992-1997 we estimate water demand elasticity at -0.3 in the short run (the effect of a price change on demand within a year of implementation) and -0.46 in the long run. We also find that, in accordance with common belief, switching from a single to a block price regime, yields a 7% reduction in average water use while maintaining the same average price. However, based on our simulations we estimate that the switch to block prices will result in a loss of approximately 1% of agricultural output due to inter-farm allocation inefficiencies.
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1 Introduction

Recent decades of population and income growth have aggravated the problem of water shortages in many parts of the world. This is increasingly leading policy makers to be interested in the use of economic incentives to rationalize water allocation. Advocated by many economists (e.g. Yaron 1991, Michelsen, Taylor, Huffaker and McGuckin 1999, Zusman 1997) as the economic inducement of choice, increasing block-rate tariffs are gaining popularity among developed as well as developing countries (Boland and Whittington 2000). Already prevalent in the residential sector (e.g. Hewitt 2000), (e.g. Arbues, Garcia-Valinas and Martinez-Espineira 2003), in recent years block-rate pricing has been gradually introduced to regulate commercial and agricultural users. Specific examples include Californian districts, European countries and Israel (e.g. Michelsen et al. 1999, Wichelns 1991, Huffaker, Whittlesey, Michelsen, Taylor and McGuckin 1998, Tsur and Dinar 1997, Kislev and Vaksin 1997, Garrido 1999).

The main goal of block prices is to induce water use reduction without burdening farmers with the full cost that simple marginal cost pricing would entail. In particular, there is concern that marginal cost pricing in agriculture would crowd out family and small farming. In contrast, an increasing price schedule allows imposing the high, socially optimal price at the margin while maintaining a lower average price, thus keeping small farms in business. Theoretical support for this assertion is provided by Bar-Shira and Finkelshtain (2000) who showed that increasing block tariffs implements the second-best social objective of maximizing welfare subject to a desired number of firms in the industry.

While conceptually attractive, the implementation of block prices raises several practical difficulties. In particular, to achieve the second-best allocation, each and every farmer should pay, in equilibrium, the socially optimal price at the margin. However, in reality, farmers' heterogeneity virtually precludes a price schedule in which every farmer pays the same price at the margin and yet pays a lower than the marginal price on average. Hence, in practice, some farmers do not reach the high price tier at all and for others the average price is too close to the marginal, resulting in inefficient inter-farm water allocation and welfare loss. The actual severity of these inter-farm inefficiencies can only be assessed empirically.

In spite of its growing use in agriculture, few empirical studies have investigated the effect of increasing block tariffs. In a pioneering study, Wichelns (1991) examines the effect of introducing increasing block-rate tariffs in the Broadview Water District in the central valley of California. The responsiveness of irrigation depth to prices was examined by means of a linear OLS regression where variations in the ratio of applied water to crop water requirement were explained by the marginal water price, soil type, area and crop dummies. In the case of melon farms, Wichelns reports that the water price elasticity equals -0.82. However, in the cases of three other crops, the hypothesis of no effect of marginal price could not be rejected. While Wichelns asserts that the no-effect result is due to small price variations, an alternative explanation is the endogeneity of the marginal price under the block-rate tariff (for details, see Moffitt 1990). Varela-Ortega, Sumpsi, Garrido, Blanco and Iglesias (1998) study the consequences of a switching reform from a single-price to block-rate prices in European agriculture. However, they utilized simulated, rather than actual data.

The reason for the small number of empirical studies on the effect of block-pricing, is partly lack of reliable data and partly the econometric complexities involved. In particular, using the conventional OLS regressions to estimate demand under block prices would result in biased estimations due to endogeneity and selection bias (e.g. Moffitt 1986, Moffitt 1990). Modern analyses of piece-wise linear budget constraints, which block prices are a special case of, use Moffitt's exposition of Burtless and Hausman's model. This framework was first used in the context of water pricing by Hewitt and Hanemann (1995), who analyzed water demand in the residential sector.

Here we adapt this methodology in order to estimate farmers' demand for irrigation water under increasing block-rate tariffs and empirically assess its effect on aggregate demand and inter-farm allocation efficiency. This methodology overcomes the technical challenges raised by increasing block-rate pricing and accounts for both observed and unobserved technological heterogeneity among farmers. Employing a micro panel data documenting irrigation levels and prices in 185 Israeli agricultural communities in the period 1992-1997 we estimate water demand elasticity at -0.3 in the short-term (the effect of a price change on demand within a year of implementation) and -0.46 in the long-term. We also find that, in accordance with common belief, switching from a single to a block price regime yields a 7% reduction in average water use while maintaining the same average price. However, based on our simulations we estimate that the switch to block prices will result in a loss of approximately 1% of agricultural output due to inter-farm allocation inefficiencies.

Our study is the first to apply Burtless and Hausman's methodology to an estimation of agricultural water demand under tier pricing. Using similar methods, Hewitt and Hanemann (1995) estimated a water demand elasticity of -1.72 for the residential sector, while Moeltner and Stoddard (2004) found an elasticity in the range of -0.23 to -0.9 in various commercial, non-agricultural sectors.

The paper is organized as follows. The next section formalizes an economic model of irrigation by heterogeneous farmers under block-rate prices. The econometric model is introduced in the third section, followed by a detailed description of the data. Section 5 presents the estimation results, which are employed in Section 6 for a simulation of alternative pricing policies and a welfare analysis. Section 7 concludes.

2 Irrigation-Water Demand Under Block Prices

Farmers maximize their profits, $\pi(w, x, z) = f(w, x, z)) - \int_0^w p(s)ds - qx$, where w is the amount of irrigated water, x is the vector of other inputs, $p(\cdot)$ and q are their prices, z is the vector of farm and industry characteristics and the price of output is normalized to 1. The technology is represented by the production function $f(\cdot)$ which is assumed to be increasing and strictly concave in w and in x.

Let $p(w)$ be the following increasing block price schedule

$$p(w) = \begin{cases} p_1 & \text{if} \quad w \leq w_1 \\ p_2 & \text{if} \quad w_1 < w \leq w_2 \\ p_3 & \text{if} \quad w_2 < w \end{cases}$$

where $p_1 < p_2 < p_3$ are price tiers and $w_1 < w_2$ are the water quantity thresholds (see Figure 1).

Straightforward maximization (the Khun Tucker conditions) implies that the farmer's input demands, w^* and x^* solve

$$\begin{aligned} f_w(w^*, x^*, z) &\geq p(w^*) \quad \text{with equality if } w^* \notin \{w_1, w_2\} \\ f_x(w^*, x^*, z) &= q. \end{aligned} \tag{1}$$

In most economic exercises the first-order conditions hold with equality. This is not the case with block prices where the water value marginal product (VMP) curve can pass between the price tiers (curve D^1 in Figure 1). In that case, $p_1 < f_w(w^*, x^*, z) < p_2$ and the optimal irrigation level is w_1. Let $D(p, q, z)$ be the inverse of the VMP, *i.e.* the water demand function under a single-price regime. Then the water demand function under block prices is given by:

$$w(\cdot) = \begin{cases} D(p_1, q, z) & \text{if} \quad D(p_1, q, z) < w_1 \\ w_1 & \text{if} \quad D(p_2, q, z) \leq w_1 \leq D(p_1, q, z) \\ D(p_2, q, z) & \text{if} \quad w_1 < D(p_2, q, z) < w_2 \\ w_2 & \text{if} \quad D(p_3, q, z) \leq w_2 \leq D(p_2, q, z) \\ D(p_3, q, z) & \text{if} \quad D(p_3, q, z) > w_2. \end{cases} \tag{2}$$

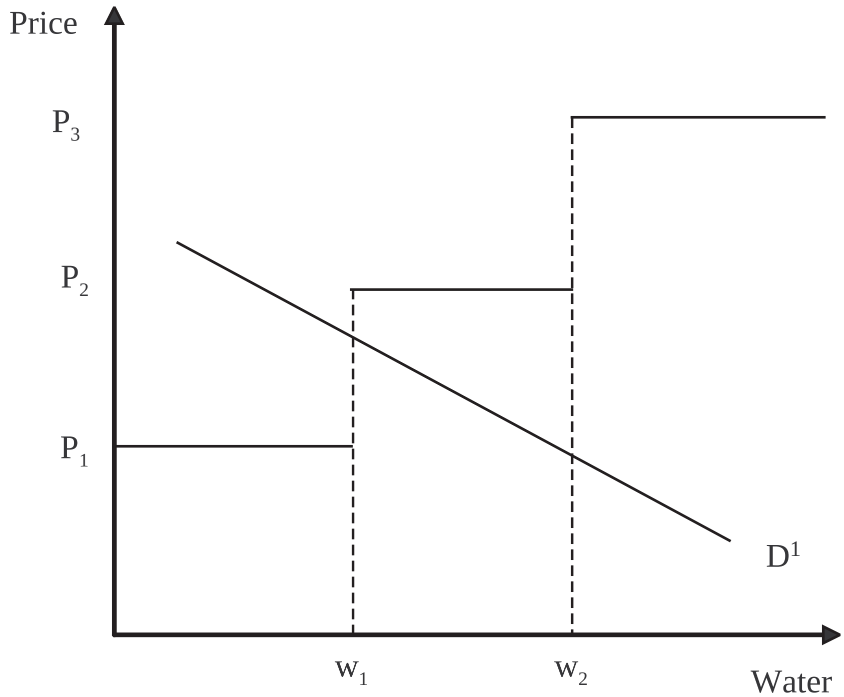

Figure 1: Increasing block price schedule

We turn now to the effect of price change on the optimal irrigation level; an infinitesimal increase in all three block prices yields:

$$w_p = \begin{cases} D_p(p_1, q, z) & \text{if} \quad D(p_1, q, z) \le w_1 \\ 0 & \text{if} \quad D(p_2, q, z) < w_1 < D(p_1, q, z) \\ D_p(p_2, q, z) & \text{if} \quad w_1 < D(p_2, q, z) \le w_2 \\ 0 & \text{if} \quad D(p_3, q, z) < w_2 < D(p_2, q, z) \\ D_p(p_3, q, z) & \text{if} \quad D(p_3, q, z) > w_2 \end{cases} \tag{3}$$

where w_p and D_p are derivatives with respect to p.[1]

It is interesting that while the VMP function is downward sloping everywhere, some price change may not induce all farmers to reduce irrigation. This can be seen in Figure 2, where we draw the water VMP curves of two farmers, D^1 and D^2, respectively. Initially, the VMP of the first intersects the price schedule at its vertical segment and the second intersects at the high price tier. A price increase of both tiers does not affect the quantity demanded by farmer 1 and therefore reflects zero elasticity. Moreover, the effect of the price increase on the second farmer is mitigated by the fact that his VMP curves intersects the price schedule at its vertical part after the change.

Hence, even if all farmers' marginal product curves reflect the same elasticities, they will react differently to a price change. The actual effect of a price change on each farmer

[1]Note that for price decrease the derivative w_p is not the same.

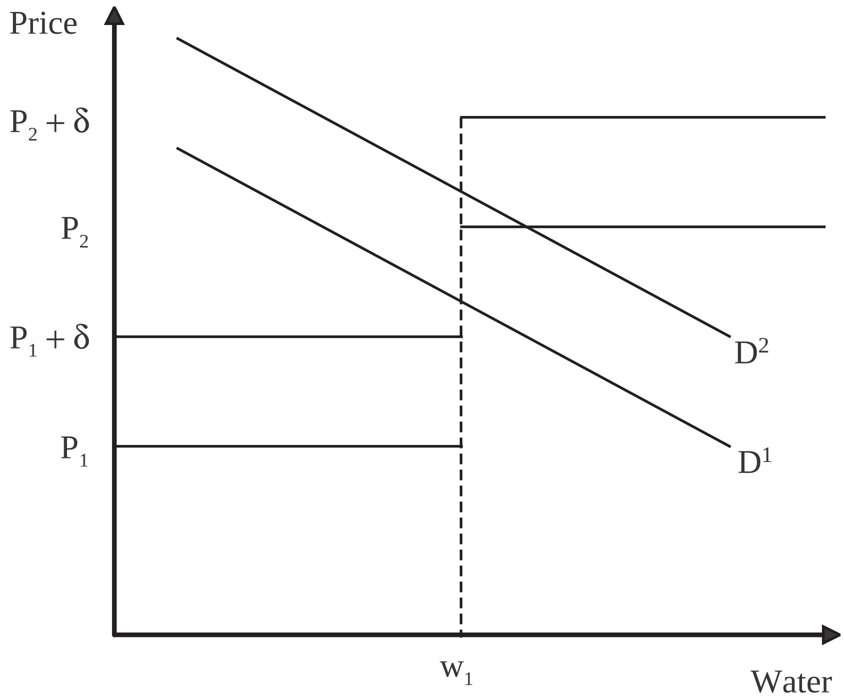

Figure 2: The effect of price change

is bounded above by the inverse elasticity of his marginal product curve. At one extreme, a farmer whose marginal product curve crosses at a vertical part of the price schedule does not react to small price changes. At the other extreme, a farmer whose marginal product curve intersects the interior of the same price block before and after a price change will show a demand elasticity as in the case of a single-price regime. Therefore, and we will elaborate on this later, under block prices water aggregate demand is less sensitive to price changes than under a single-price regime.

3 Econometric Methodology

We begin this section by introducing the statistical model, proceed with the estimation methodology and conclude with the parametric specification of demand.

3.1 The statistical model

Two sources of randomness are introduced into the economic model presented in Section 2. Farmers' heterogeneity, $\alpha \sim (0, \sigma_\alpha)$, and a measurement error, $\varepsilon \sim (0, \sigma_\varepsilon)$. The first is a *random effect* which captures variations across farms and over-time that are neither explained by the observed characteristics nor by farm dummies. These may include variations in management capabilities, local climatic conditions, etc. Although these factors are unobserved by the econometrician, they are known to the farmer and hence taken into account in his optimization process. The latter term, ε, represents a measurement error or mistakes in the farmer's optimization.

Following the literature on piece-wise linear budget constraints (e.g. Hausman 1985, Moffitt 1986, Hewitt and Hanemann 1995), we adopt a linear additive formulation. Under this assumption, we can rewrite (2) as:

$$w = \begin{cases} D(p_1,q,z)+\alpha+\varepsilon & \text{if} \quad \alpha < w_1 - D(p_1,q,z) \\ w_1+\varepsilon & \text{if} \quad w_1 - D(p_1,q,z) \le \alpha \le w_1 - D(p_2,q,z) \\ D(p_2,q,z)+\alpha+\varepsilon & \text{if} \quad w_1 - D(p_2,q,z) < \alpha < w_2 - D(p_2,q,z) \\ w_2+\varepsilon & \text{if} \quad w_2 - D(p_2,q,z) \le \alpha \le w_2 - D(p_3,q,z) \\ D(p_3,q,z)+\alpha+\varepsilon & \text{if} \quad \alpha > w_2 - D(p_3,q,z). \end{cases} \tag{4}$$

This formulation reveals that while both α and ε affect the observed demand, α alone determines the optimal choice of the price block and the optimal irrigation level. It is apparent that variations in the farmer's type, α, in some ranges, would change neither his optimal choice nor his observed one. By contrast, ε is unknown to the farmer and hence only influences the observed irrigation level. In some cases, its realization could shift the farmer's observed demand from the optimal block to another. This important distinction between α and ε is key in facilitating the econometric identification of the two errors.

3.2 The estimation approach

Following Moffitt and others, we apply a maximum likelihood approach to estimate the model. We proceed with the derivation of the observation likelihood. Denote by $Pr(w|p(\cdot),z,\theta)$ the probability of observing a certain level of irrigation, w, given the price schedule $p(\cdot)$, and the observed characteristics z and θ. The latter includes demand parameters as well the parameters of the distributions of the errors α and ε. The probability of observing w is the sum of the joint probabilities of observing that irrigation value and that the planned decision is on each block or vertical segment of the price schedule. That is,

$$\begin{aligned} & Pr(w|p(\cdot),z,\theta) = Pr[\alpha+\epsilon = w - D(p_1,q,z), \alpha \le w_1 - D(p_1,q,z)] \\ + \quad & Pr[\alpha+\epsilon = w - D(p_2,q,z), w_1 - D(p_2,q,z) < \alpha \le w_2 - D(p_2,q,z)] \\ + \quad & Pr[\alpha+\epsilon = w - D(p_3,q,z), \alpha > w_2 - D(p_3,q,z)] \\ + \quad & Pr[\epsilon = w - w_1, w_1 - D(p_1,q,z) \le \alpha \le w_1 - D(p_2,q,z)] \\ + \quad & Pr[\epsilon = w - w_2, w_2 - D(p_2,q,z) \le \alpha \le w_2 - D(p_3,q,z)]. \end{aligned} \tag{5}$$

The sample likelihood function is then given by

$$L = \prod_{all\ \ obs} Pr(w|p(\cdot), z, \theta). \tag{6}$$

Assuming that the errors are statistically independent and normally distributed, $\alpha \sim N(0, \sigma_\alpha)$ and $\varepsilon \sim N(0, \sigma_\varepsilon)$, the likelihood in (6), in terms of the standard normal density, is readily derivable. The details of the calculations are relegated to the appendix.

3.3 The empirical specification

In the previous subsections, we outlined the economic model and sketched an adequate econometric methodology to estimate it. We proceed by further parameterizing the model to be estimated. The empirical specification incorporates two additional features: first we allow Box-Cox transformation of the data, second we introduce lagged prices into the model. It follows that $D(p, q, z)$ can be expressed as[2]

$$\frac{w_{it}^\lambda - 1}{\lambda} = \beta_0 + \beta_1 \sum_{j=0}^{\infty} \gamma^j \frac{p_{t-j}^\lambda - 1}{\lambda} + \beta_2 \sum_{j=0}^{\infty} \gamma^j \frac{q_{t-j}^\lambda - 1}{\lambda} + \delta \frac{z_{it}^\lambda - 1}{\lambda}. \tag{7}$$

The Box-Cox transformation is a relatively general formulation, nesting several functional forms such as the logarithmic ($\lambda = 0$), the linear ($\lambda = 1$), the quadratic ($\lambda = 2$) and the reciprocal ($\lambda = -1$) models as special cases. The parameter λ is to be estimated simultaneously with the other parameters of the model so as to maximize the likelihood of the data.

Lagged-price effects are required here because we believe that the full

effect of price changes may take time. That is, farmers may adjust their irrigation level gradually in response to water-price changes. The reason for this is twofold. First, farmers may not change their scale of operation as long as they are not sure whether the price change is permanent or transitory. In this case, the irrigation response is limited by the horticultural requirements. Second, and not unrelated to the first reason, is the irreversible investment embodied in agriculture. An increase in input price in the presence of irreversible investment may leave the farmer covering his variable cost but not the total cost. In this case, the farmer will let his capital depreciate and irrigation will continue as long as it is unnecessary to reinvest. Consequently, water price changes have both short-term and a long-term effects.

As variables explaining heterogeneity, z_{it}, we use each farmer's water quota, which serves as an instrument for both the quantity and for his type of land, rainfall in his region,

[2]To ease the notation we treat z as a single variable in equations (7) and (8).

other input prices, a time trend to capture technological changes - in particular water saving ones - and a farmer dummy to capture a farm *fixed effect*.

Taking the lag of (7) and multiplying it by γ yields

$$\gamma \frac{w_{it-1}^{\lambda}-1}{\lambda} = \gamma\beta_0 + \gamma\beta_1 \sum_{j=0}^{\infty} \gamma^j \frac{p_{t-j-1}^{\lambda}-1}{\lambda} + \gamma\beta_2 \sum_{j=0}^{\infty} \gamma^j \frac{q_{t-j-1}^{\lambda}-1}{\lambda} + \gamma\delta \frac{z_{it-1}^{\lambda}-1}{\lambda}.$$

Subtracting the last equation from the previous one and rearranging yields:

$$\frac{w_{it}^{\lambda}-1}{\lambda} = (1-\gamma)\,\beta_0 + \gamma \frac{w_{it-1}^{\lambda}-1}{\lambda} + \beta_1 \frac{p_t^{\lambda}-1}{\lambda} + \beta_2 \frac{q_t^{\lambda}-1}{\lambda} + \delta \frac{z_{it}^{\lambda}-1}{\lambda} - \gamma\delta \frac{z_{it-1}^{\lambda}-1}{\lambda} \tag{8}$$

which is the model that we estimate.

A few points are worth mentioning. First, while β_1 measures the short-term effect of price change on water use, $\beta_1 \frac{1}{1-\gamma}$ measures its long-term effect. Second, the coefficient of the last term in (8) is restricted to be equal to $\gamma\delta$. In the empirical application we test this restriction in order to assess the validity of the lag structure in (7).

4 Institutional Description and Data

Israeli agriculture consists of two main subsectors. The "private sector" comprises several hundred individual farmers, who together cultivate 20% of irrigated crops and produce 15% of agricultural value added. The second sector is the "planned agricultural" sector, producing the remaining 85% of the agricultural product. The planned sector consists of 415 cooperative villages - *Moshavim* - and 315 collectives—*Kibutzim*. By the Israeli water law, each of the individual farmers and each of the 730 planned cooperatives is legally a single consumer receiving an annual water allocation and responsible for payment. Our micro level panel describes irrigation and geographical data for all agricultural water consumers in Israel, during the period 1992-1997. The data set was assembled in a joint project of the Israeli Central Bureau of Statistics (CBS) and the Agricultural Ministry Planning Authority and is documented in an annual publication of the Agricultural Ministry: "Agricultural Sub Industries" for various years. To ensure confidentiality, the CBS reports the data of individual farmers aggregated over large geographical areas. For that reason, we omit all individual farmers from our analysis and focus on the "planned" sector.

Farmers in our panel use three types of water for irrigation: fresh water, which accounts for 80% of the total irrigation, and semi-saline and reclaimed water comprising the rest. Since various types of water differ in their marginal product, we limit our analysis to farmers who consume only fresh water. Under the 1959 Water Law, the supply of fresh water is regulated by the state. The main water supplier - Mekorot - a state-owned company, provides 60% of the aggregate fresh-water supply to agriculture, whereas most of the rest is supplied by regional utilities and water cooperatives that are owned by the farmers themselves. Whereas Mekorot charges farmers according to the price schedule imposed by the state, the data from other suppliers, who set their prices independently, are not available for research. Thus, we are left with 185 cooperatives and semi-cooperatives, those who use only fresh water supplied by Mekorot. They comprise 20% of the water use for irrigation in Israel.

4.1 Water quotas

The Israeli water law legislated in 1959 determines that water in both underground and surface aquifers is state property. This implies that the pumping of water and its application for irrigation require permits. Each year, the state water commissioner issues pumping permits and each planned village is allocated a specific annual water quota for irrigation. Initial, historical quotas were allotted according to village's total land suitable for irrigated agriculture, its land type and to other factors such as population size, location, water usage in years prior to 1959, and political and organizational affiliations. During the 1960s and 1970s, quotas were enlarged to reflect the development of new water sources in Israel and the establishment of new agricultural settlements in the Galilee and Negev regions. Occasionally, water quotas were adjusted to reflect changes in the amount of land available for farming. In recent years, however, increased urban demand and the allocation of water to Jordan as dictated by the 1994 peace treaty, have led to a gradual reduction in the quotas of fresh water for irrigation.

Important questions for our analysis are: to what extent were the quotas enforced and were they binding? If they were, water application would reflect quotas rather than the marginal product of water. With respect to the second question, we note that in the last 20 years, with the exception of years with severe droughts, the Israeli agricultural sector consumed less fresh water than was allotted to it. In our sample, only 12% of the farmers exceeded their quotas and even then, there is not a single case recorded where any sanction was imposed. Thus, we are satisfied that the observed water application does reflect water marginal productivity.

Table 1 reports some descriptive statistics of water quotas in our panel. It shows considerable variations in the levels of the quotas across farmers and over the years. The differences across users reflect the fact that at the time when the quotas were allocated, some farmers were growing water-intensive crops, such as citrus, while others relied on crops

Table 1: Quotas (thousands of cubic meters)

Year	Mean	Max	Min	Std. Dev.
1992	773	1550	213	264
1993	927	1650	239	297
1994	844	1539	229	273
1995	848	1650	239	269
1996	833	1716	224	283
1997	978	1854	239	333
All Years	867	1854	213	295

such as wheat or olives. Aggregate variations over the years are primarily a reflection of fluctuations in annual precipitation.

4.2 Water consumption

The average cooperative in the sample cultivates 960 acres of land, of which 600 acres are allotted to field crops, 200 acres to orchards, and the rest is used for intensive farming of vegetables and flowers, mainly in greenhouses. To cultivate these crops, the average farm utilizes 650,000 m^3 (approximately 488 acre-feet) of water, an average of 677 m^3 of water per acre. However, this average consumption is misleading, since a significant share of the land is allocated to dry agriculture and consumes no water. It is interesting to note that the value of the crops per cubic meter have increased fivefold since 1950 (Kislev 2001). This reflects considerable technological improvement, in particular in the 1990s, during which our data was generated. Water productivity is estimated to have increased by 25%. Our empirical formulation of the demand function in the next section takes those changes into account by adding a productivity trend variable.

Table 2 presents the descriptive statistics of agricultural water consumption. It shows considerable variations in water application across farmers, while changes in aggregate consumption over the years are moderate. The explanations for the differences across cooperatives are similar to those provided for the dissimilarities in quotas. However, it is interesting to note that changes in consumption over the years are much smaller than the fluctuations in the quotas. This strengthens our contention that the consumed quantities reflect economic demand, rather than administrative decisions.

Table 2: Agricultural Water Consumption (thousands of cubic meters)

Year	Mean	Max	Min	Std. Dev.
1992	609	1600	83	281
1993	647	1346	83	296
1994	626	1500	83	298
1995	654	1600	100	315
1996	637	1600	38	338
1997	649	1441	38	349
All Years	637	1441	38	313

4.3 Prices

By the early 1990s Israel was going through a rapid process of depleting its water reservoirs. The 1959 Water Law, which had been enacted under entirely different circumstances, authorized the agricultural minister to levy charges to recover the cost of supplying the water. Furthermore, every price change had to be approved by the financial committee of the parliament, where the government encountered a powerful agricultural lobby. In its quest to rationalize water consumption, while adhering to the spirit of the law of preventing the crowding out of small, financially fragile farmers, the government of Israel established, in 1991, the increasing block-rate tariff.

According to the ordinances there are three tiers, which are determined by historical quotas. The lowest price block is applied for consumption of up to 50% of the (historically fixed) quota. The medium tariff is levied on consumption between 50% and 80% of the quota and consumption above 80% is charged the highest price block. Using historical quotas as benchmarks for price blocks creates exogenous variation across farmers regarding the price schedules they face, thus facilitating our estimation.

Figure 3 presents the three block-rates (deflated by the CPI) from 1992 to 1997. It can be seen that initially, real prices declined, but since 1994, all three block rates have been rising. Note, however, that the relative prices of the various tiers have changes over the period, reflecting a dynamic political compromise between the farmers' lobby and the government on the one hand, and varying amounts of precipitation on the other. Table 3 presents the distribution of irrigated water over the three tiers in the years 1992-1997. The first column in Table 3 presents the yearly proportion of water used by farmers who pay the lowest price **at the margin**, columns 2 and 3 the proportions used by farmers who pay at the margin the middle and highest prices, respectively, and column 4 shows the total water

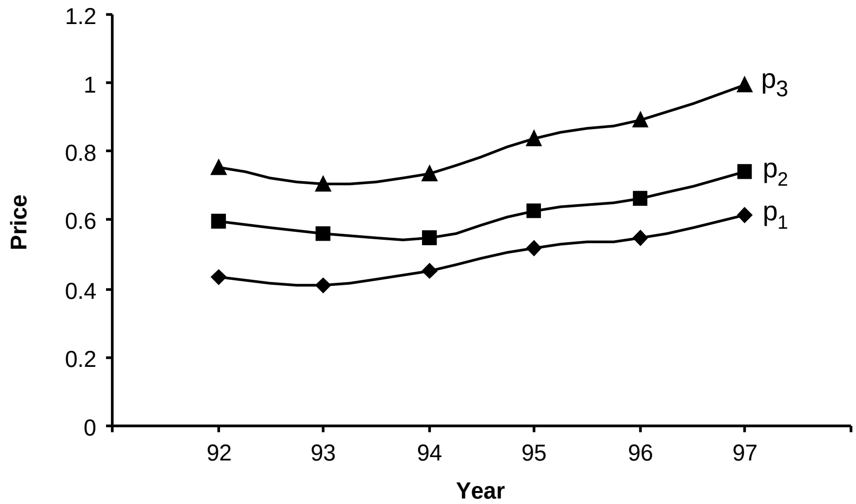

Figure 3: The block rates

Table 3: Distribution of Water (Farmers) over Tiers (percentage)

Year	1st Tier	2nd Tier	3rd Tier	Total (1000 m^3)
1992	12 (23)	58 (55)	30 (22)	112704
1993	11 (21)	44 (46)	45 (33)	119743
1994	14 (24)	44 (46)	42 (30)	115660
1995	13 (24)	39 (42)	48 (34)	120917
1996	16 (31)	41 (40)	43 (29)	117800
1997	15 (31)	40 (39)	45 (30)	120091
Average	14 (26)	44 (44)	42 (30)	117853

use in our panel. In parentheses we report the proportion of farmers paying each price at the margin.

On average, 26% of the farmers, using 14% of the water, paid only the lowest price block while 44% of the farmers who utilized 44% of the aggregate water did not reach the highest price block. Only 30% of the farmers reached the highest price tier, with a consumption share of 42%. The first column reveals that the share of farmers in the first block is nearly twice their consumption share, indicating preservation of small farms. As was mentioned earlier, under the assumption that the highest price block represents the social marginal value of water, the actual allocation is not optimal. The data also shows some dynamics over the years, but not a

monotonic trend. Our econometric findings are reported in the following sections.

5 Results and Simulations

The parameter estimates of our model are reported in the following subsection, where we discuss the effects of the various explanatory variables on irrigation, their significance and goodness of fit. In the second subsection we use our estimates to compute the elasticities under the various specifications and compare them with the aggregate elasticities obtained via simulation.

5.1 Estimation

The first column in Table 4 presents the maximum likelihood estimates of our model. We find that the maximum likelihood estimate of λ is 0.653, an intermediate case between the linear and logarithmic models. We tested both hypotheses, that $\lambda = 0$ and that $\lambda = 1$, and rejected them at a 1% significance level. Thus, the Box-Cox specification is significantly better than both restricted forms – the logarithmic and the linear. Nevertheless, we also report the estimates for the logarithmic and linear models in order to assess the sensitivity of our results to the functional form.

In all three specifications, water prices have a significant, negative effect on water use. In contrast, the price of other inputs has a small insignificant effect, implying small substitutability between water and other inputs. As expected, 'Rain' has a significant (with the exception of the linear specification) negative effect on water demand. This variable indicates the amount of rainfall during the months of April and October. In Israel, there is virtually no rain between May and September, which is when crops are irrigated extensively and there is abundant precipitation between November and February, during which time there is no irrigation at all. During the months of April and October rainfall is spurious with high variability. Therefore, relatively high precipitation levels during April and October reduce irrigation during those months and thereby annual irrigation. As expected, more rain during those months significantly reduces farmers' annual irrigation levels.

Lagged water use turns out to be highly significant in our data, suggesting that the effect of price changes extends beyond its first-year effect. Specifically, the long-term effect, $\sum_{j=0}^{\infty} \gamma^j = 1/(1-\gamma)$, varies between 1.62 in the linear to 1.81 in the logarithmic specifications. That is, over-time, water use changes by an additional 62 to 81% of its first-year effect.

Water quota is also highly significant, increasing irrigation between 0.21 m^3 for each additional cubic meter of quota in the Box-Cox, to 0.31 in the logarithmic specification. That is the case although the quotas were not binding for almost 90% of the farmers and penalties

Table 4: Maximum Likelihood Estimates

Specification Coefficient	Box-Cox	Logarithmic ($\lambda = 0$)	Linear ($\lambda = 1$)
λ	0.65 (17.13)	– –	– –
Constant	-116.70 (-1.66)	-5.85 (-3.20)	16.77 (0.05)
Water Price	-31.90 (-2.25)	-1.32 (-6.56)	-141.68 (-2.00)
Other Input Prices	1.66 (0.21)	0.30 (0.93)	16.88 (0.32)
Rain	-0.24 (-2.23)	-0.04 (-2.08)	-0.37 (-1.40)
Lagged Water Use	0.40 (13.62)	0.45 (16.64)	0.38 (13.53)
Water Quota	0.26 (5.23)	0.35 (3.60)	0.22 (5.78)
Year	1.72 (2.34)	0.08 (4.69)	5.12 (1.27)
σ_α	8.82 (3.88)	0.27 (25.80)	98.40 (17.17)
σ_ε	7.57 (3.52)	0.11 (14.69)	37.87 (2.45)
Log Likelihood	-6638	-6744	-6678
Number of Obs.	1110	1110	1110
Number of Coefs.	194	193	193

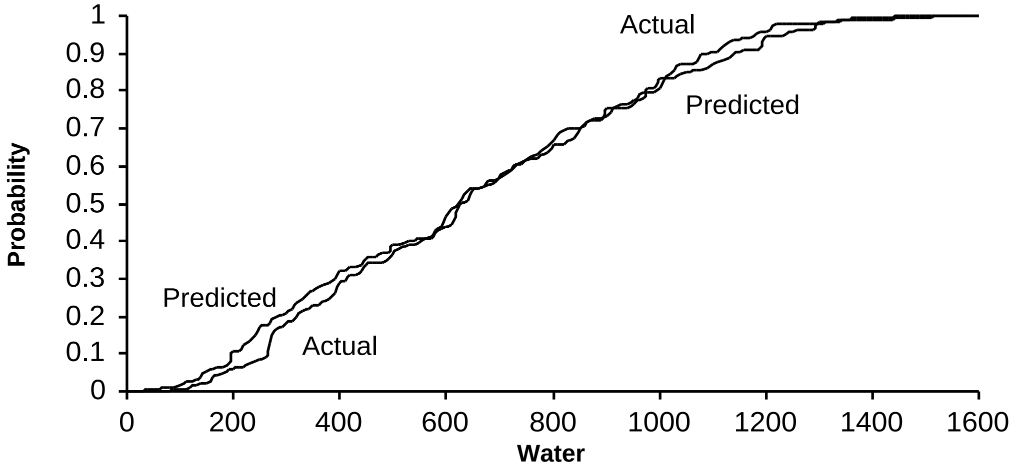

Figure 4: Predicted versus Actual Water Distributions

were never imposed when they were surpassed. The explanation for this correlation is that quotas serve as a proxy for the farmers' amount and type of land, which are known to the water commissioner and to the farmer, but are unobservable for the researcher. For example, it may be known to the commissioner that some farmers' lots are sandy and therefore require more irrigation than the region's average. In this example, farmers with higher quotas irrigate more than the average farmer, generating a positive correlation.

The total 'Year' effect is positive indicating, ceteris paribus, augmentation of water use over-time. The 'Year' effect is twofold. On the one hand, it captures productivity improvements in agriculture, increasing the VMP and therefore raising water use. On the other hand, over-time improvements in water saving technology may under certain circumstances reduce water use. Our findings indicate that the productivity effect dominates the water saving effect.

It is conventional to assess the estimation's goodness of fit by comparing the predicted to the actual properties of the water use distributions.[3] Figure 4 draws those two distributions and reveals that the two are quite similar. Table 5 lists the moments of the two distributions. In particular, note that the predicted average use in our sample is 654 (1000 m^3) compared with the actual average use of 649.

Finally, we tested the lag structure by estimating the model twice; once under the constraint that the coefficient of $\frac{Z_{it-1}^{\lambda}-1}{\lambda}$ in equation (8) equals $\gamma\delta$ and once without that constraint. Comparing the maximum likelihood of the two models, we could not reject, at

[3]The predicted values were obtained by simulating our sample at the actual prices, as outlined in the Appendix.

Table 5: Moments of Actual and Predicted Water Use

	Observed Use	Predicted Use
Mean	649	654
Median	624	625
Maximum	1441	1518
Minimum	38	87
Std. Dev.	348	322
Skewness	0.228	0.261
Kurtosis	2.007	2.178

a 5% significance level, the restricted model in favor of the unrestricted one, supporting the validity of the lag structure.

5.2 Elasticities

To make the results reported in Table 4 comparable across the different specifications, we compute the demand elasticities with respect to the variables and report them in Table 6. With respect to water prices we report in the first row the estimated elasticities for the short-run, computed as $\beta_1(p/w)^\lambda$. These elasticities reflect the slope of the VMP curve and would have been the actual elasticities under a single-price regime. However, as we argued earlier, under block-rate pricing, the estimated elasticity overestimates the aggregate elasticity because some farmers are located at the horizontal section of the price schedule. To evaluate the aggregate elasticity we simulated the irrigated agricultural sector to predict water use under the actual price schedule and under the actual price schedule increased by 1%. It is evident that the greatest gap between the aggregate and estimated elasticities is with the logarithmic specification and the smallest is with the linear. The reason for this is that the steeper the demand curve, the less farmers will optimally choose to use the threshold water levels between blocks. We estimate the proportion of farmers that would optimally choose thresholds w_1 or w_2 to be 35% under the logarithmic, 25% under the Box-Cox and 10% under the linear specification.

Rain in April and October turns out to have a smaller than expected effect, although recall from Table 4 that it has the expected sign and is significant under the Box-Cox and logarithmic specifications. As the precipitation variability during these months is large, it is not uncommon to have years with either half or twice the long-term average. Comparing such two years, the irrigation levels between them will differ by 6 to 7 percents, according to the first two specifications. Finally, it is interesting to note that the effects of water quotas

Table 6: Demand Elasticities

		Box-Cox	Logarithmic	Linear
Estimated price elasticity	Short-Run	−0.358	−1.323	−0.146
	Long-Run	−0.594	−2.401	−0.237
Aggregate price elasticity	Short-Run	−0.299	−0.812	−0.133
	Long-Run	−0.496	−1.474	−0.216
Other-input price elasticity	Short-Run	0.033	0.303	0.041
	Long-Run	0.055	0.550	0.067
Rain elasticity		−0.047	−0.043	−0.029
Water Quota elasticity		0.347	0.348	0.333

and past use are quite similar in all three specifications.

To summarize the main finding of this section, although there are considerable differences across model specifications farmers' demand appears to be sensitive to water price changes and almost half of the effect occurs after the first year of the price change. In the following section we study the effects of switching from a single-price regime to a block-rate one on water demand and social welfare.

6 Comparing Increasing Block-Rate Tariff and the Single-Price Regimes

Block-rate pricing advocates argue that facing a marginal price that reflects the marginal social cost of water, farmers will use the socially optimal quantity while paying less, on average, than the marginal social cost. This way, they argue, society can reach the optimal water allocation without burdening farmers with the costs associated with marginal pricing. A necessary condition to achieving the optimal allocation is that *every* farmer's demand function intersect the price schedule at the highest tier. Unfortunately, that is practically impossible due to farmers' heterogeneity. In practice, the more productive farmers, who, on average, use larger quantities of water, reach the highest price block, while less productive farmers, who irrigate less, pay a lower marginal price. Figure 5 demonstrates the differential effect of switching from a single-price p_0 to the block prices p_1, p_2 on farmers with different demand functions. Following the switch, lower-productivity farmers enjoy a *marginal price reduction* from p_0 to p_1 and consequently increase their water use from w_1 to w_2, while high-productivity farmers experience a marginal price increase from p_0 to p_2 and reduce their use from w_4 to w_3. Thus, if low-productivity farmers increase their use and high-productivity

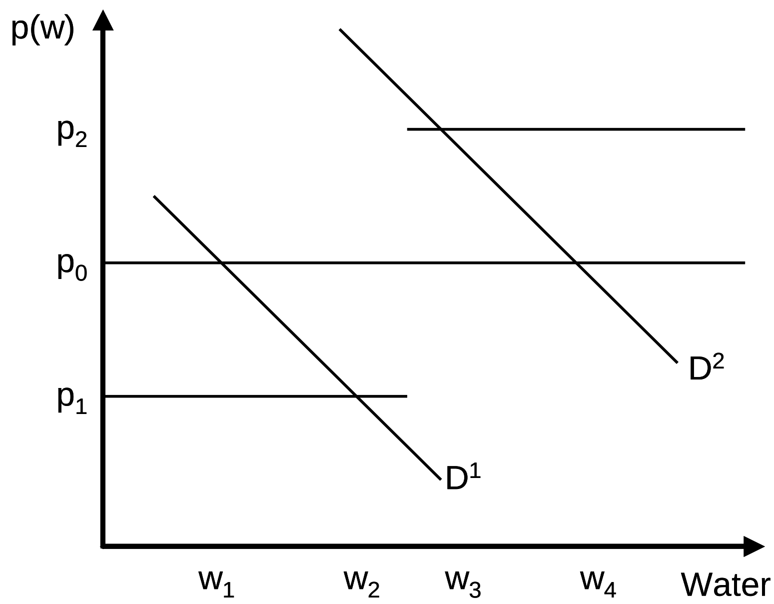

Figure 5: Welfare effect of block-rate pricing

farmers reduce it, what is the net effect of a change in price regime on aggregate water use? Analytically, if the average water price remains the same under the two regimes, the answer is equivocal. Empirically, we can answer this question by simulating a price-regime switch.

We first simulate the agricultural sector to find the predicted water demand under a single-price that equals the 1997 weighted average price. That is, we examine a hypothetical switch from the actual block-price regime that existed in 1997 to a single-price regime that would have maintained the same average price. It turns out that the aggregate demand under the single-price regime is 7% higher than under block pricing. This outcome confirms the basic intuition that block prices can substantially reduce the level of irrigation without raising the cost to farmers. Thus, in Israel, where the switch from a single-price regime to block prices was designed specifically for that purpose, it achieved its purpose.

Alternatively, suppose that the policy-maker chooses a single-price that would maintain the same actual aggregate irrigation level. It turns out that such a single-price would be 20% higher than the average price under block prices. Again, this outcome confirms our prior belief that overall, farmers pay less under block prices. However, as we saw in Figure 5 a switch to block prices has different effects on farmers with different productivity levels and therefore has potentially important welfare consequences, an issue that we explore next.

By welfare, we mean the sum of all farmers' profits, plus the total amount they pay for water, minus the cost of water extraction, C. Since we compare welfare under the two price

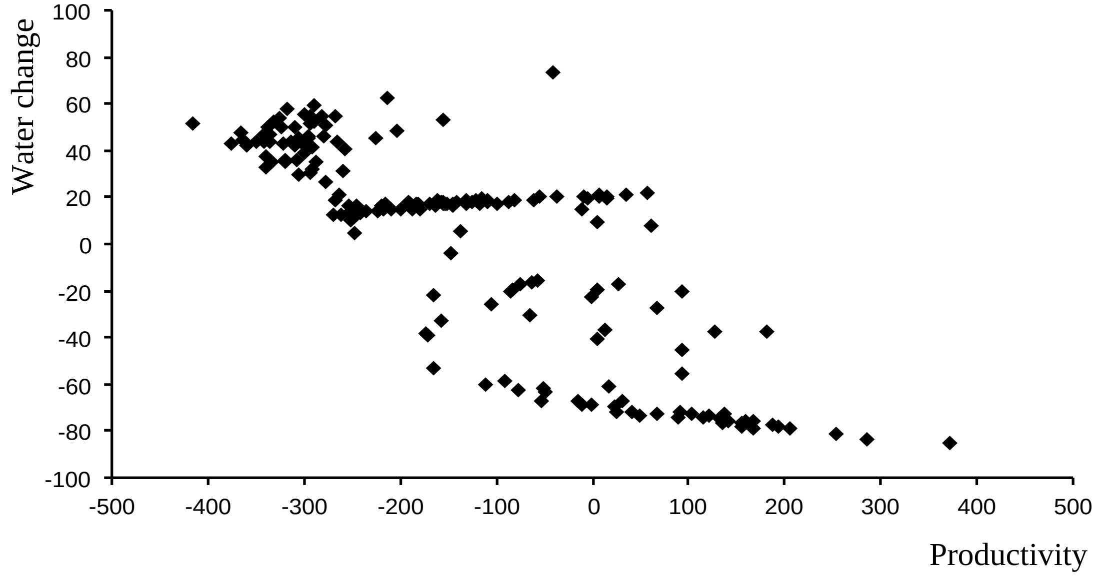

Figure 6: Inter farm inefficiency

regimes with the same aggregate water use, C is the same in both cases and can be ignored.

In practice, we calculate for each farmer the area under the estimated demand function (8) in 1997 and sum it over all farmers. We find that, empirically, switching to a block-price regime in a way that would maintain the same aggregate use raises all farmers' profits and reduces the revenues to water suppliers.[4] Since the loss to the water suppliers was larger than the farmers' aggregate gain, the switch to block prices lowered social welfare. This does not mean that **all** farmers always gain from block prices; rather it is the outcome of the specific price schedule that was enacted in Israel at the time of our study. What is inevitable is the aggregate welfare loss that follows a switch to block-rate pricing and the reason for this is quite simple. Let $\tilde{p}$ denote the single-price that maintains the same aggregate use. It must be that the lowest price tier, p_1, is lower than $\tilde{p}$ and the highest price p_3 is higher than $\tilde{p}$; that is, $p_1 < \tilde{p} < p_3$. Therefore, the smaller the quantity a farmer uses, the closer the average price he pays to p_1 and the greater his gain from switching to block prices. In contrast, the greater the quantity he uses, the closer the average price he pays to p_3 and the less he gains from the switch. As we saw in Figure 5, less productive farmers have lower demand curves than the more productive ones. Hence, the former increase their water use at the expense of the latter which is socially inefficient. This inter-farm inefficiency, is demonstrated in Figure 6, where for each farmer we chart the change in predicted water use when switching from a single-price

[4]Actually, since in 1997 there was a block-rate regime, we estimated the effect of switching from it to the single price that would have maintained the same aggregate use.

to block pricing, against his productivity level.[5] Overall, farmers with higher productivity irrigate more under a single-price regime, while less productive farmers irrigate more under block prices. This outcome should not come as a surprise as one of the manifested goals of block pricing is to encourage small and medium farming.

In the above simulation, we compared price regimes that yield the same aggregate irrigation level. We also examined a switch from a single-price regime to block pricing that would maintain the same average cost (rather than maintaining the same aggregate use in the previous exercise). In that case, we find that aggregate demand declines by approximately 14%. Since under this alternative the average cost of water is the same in the two regimes, some farmers lose from the switch. Again, since block pricing favors small farmers, less productive farmers benefit while the more productive farmers lose from the switch. The difference between this and the previous scenario is that the cost of switching to block pricing shifts from the water suppliers to the more productive farmers, who are inadvertently subsidizing their less productive colleagues.

7 Conclusions

We found that farmers' water demand is responsive to prices. We estimate that within the first year after a price increase water demand responds with an elasticity of 0.3 and that demand continues declining thereafter, reaching a long-term elasticity of 0.46. Our findings corroborate the common belief that using block pricing enables policy-makers to lower aggregate demand without raising the average cost of water to farmers. Comparing the actual block prices that existed in 1997 to the theoretical single-price that would maintain the same average cost, we estimated that under block prices, the aggregate water use would be 7% lower, that small farmers would pay a lower average price and use more, and that large farmers would pay a higher average price and use less than under the single-price regime. Hence, switching to block pricing achieves the dual goal of reducing aggregate water use without increasing the water cost for small family farms.

And yet, these achievements should be balanced against the inter-farm inefficiencies that are typically the outcome of block pricing. These inefficiencies stem from the fact that in reality there is large heterogeneity among farmers and therefore it is practically impossible to design a block-pricing schedule in which all farmers will pay the same marginal price. In reality, small farmers will pay the lowest price tier whereas large farmers will pay the

[5] As proxy to farmers' productivity we use their estimated dummy's parameter value.

highest (socially optimal) price tier at the margin. Hence, in switching from a single-price regime to a block price regime that results in the same average cost, small farmers would increase irrigation while large farmers will reduce it. If smaller farms are typically less productive (if they were more productive they would have earned more and would have expanded production), then block pricing benefits the less productive at the expanse of the more productive. The alternative of course is to benefit all farmers at the expense of the water suppliers.

While potential inter-farm inefficiencies are a serious concern, our estimates indicate that with our dataset, the switch to block pricing creates an efficiency loss of less than 1% of total agricultural output. Given the alternative of switching to block pricing at the same average water cost and bearing an efficiency loss of 0.4% or raising water prices by 20% (in both cases the aggregate water use declines by 7%), most policy-makers would find the block-pricing option more appealing.

8 Appendix

Once the normality assumption is imposed on the errors, we can formulate the observation likelihood in terms of the standard normal distribution. We then outline the simulation procedure. To this end, let $v = \alpha + \varepsilon$ and let $g_{v\alpha}(v, \alpha)$ denote the joint density of v and α. The density $g_{v\alpha}$ is bivariate normal with parameters $\sigma_v^2 = \sigma_\alpha^2 + \sigma_\varepsilon^2$, σ_α^2 and $\rho = \sigma_\alpha / \sigma_v$. Likewise, $g_{\varepsilon\alpha}$ denotes the density of the variable(s) in the subscript. We will also use f and F to denote the density and the cumulative distribution functions of a standard normal random variable, respectively.

The probability of observing a certain irrigation level w (given in equation (5)) is essentially the observation likelihood function, which can be expressed in terms of g as follows

$$
\begin{aligned}
L(w, \theta) &= \int_{-\infty}^{\alpha_1} g_{v\alpha}(w - D(p_1, q, z), \alpha) d\alpha + \int_{\alpha_1}^{\alpha_2} g_{\varepsilon\alpha}(w - w_1, \alpha) d\alpha \\
&+ \int_{\alpha_2}^{\alpha_3} g_{v\alpha}(w - D(p_2, q, z), \alpha) d\alpha + \int_{\alpha 3}^{\alpha_4} g_{\varepsilon\alpha}(w - w_2, \alpha) d\alpha \\
&+ \int_{\alpha_4}^{\infty} g_{v\alpha}(w - D(p_3, q, z), \alpha) d\alpha \qquad (9)
\end{aligned}
$$

where $\alpha_1 \equiv w_1 - D(p_1, q, z)$, $\alpha_2 \equiv w_1 - D(p_2, q, z)$, $\alpha_3 \equiv w_2 - D(p_2, q, z)$, and $\alpha_4 \equiv w_2 - D(p_3, q, z)$.

The definition of conditional distribution implies that $g_{v,\alpha}(v, \alpha) = g_{\alpha|v}(\alpha \mid v) g_v(v)$, and the independence of ε and α implies that $g_{\varepsilon\alpha} = g_\varepsilon g_\alpha$. Hence, equation (9) becomes

$$
\begin{aligned}
L(w, \theta) &= g_v(w - D(p_1, q, z)) \int_{-\infty}^{\alpha_1} g_{\alpha|v}(\alpha) d\alpha + g_\varepsilon(w - w_1) \int_{\alpha_1}^{\alpha_2} g_\alpha(\alpha) d\alpha \\
&+ g_v(w - D(p_2, q, z)) \int_{\alpha_2}^{\alpha_3} g_{\alpha|v}(\alpha) d\alpha + g_\varepsilon(w - w_2) \int_{\alpha_3}^{\alpha_4} g_\alpha(\alpha) d\alpha \\
&+ g_v(w - D(p_3, q, z)) \int_{\alpha_4}^{\infty} g_{\alpha|v}(\alpha) d\alpha \qquad (10)
\end{aligned}
$$

Since $g_{v,\alpha}$ is bivariate normal it follows that $g_{\alpha|v}(\alpha \mid v)$ is $N(\rho^2 v, \sigma_\alpha^2(1 - \rho^2))$. We are now ready to write the observation likelihood in term of f and F:

$$
\begin{aligned}
L(w, \theta) &= \frac{1}{\sigma_v} f(h_1) F(r_1) + \frac{1}{\sigma_\varepsilon} f(s_1) [F(r_2) - F(r_1)] \\
&+ \frac{1}{\sigma_v} f(h_2) [F(r_3) - F(r_2)] + \frac{1}{\sigma_\varepsilon} f(s_2) [F(r_4) - F(r_3)] \\
&+ \frac{1}{\sigma_v} f(h_3) [1 - F(r_4)] \qquad (11)
\end{aligned}
$$

where

$$
\begin{aligned}
h_j &= [w - D(p_j, q, z)]/\sigma_v \qquad j = 1, 2, 3 \\
s_j &= (w - w_{ij})/\sigma_\varepsilon \qquad j = 1, 2 \\
t_1 &= \alpha_1/\sigma_\alpha \\
t_2 &= \alpha_2/\sigma_\alpha \\
t_3 &= \alpha_3/\sigma_\alpha \\
t_4 &= \alpha_4/\sigma_\alpha \\
r_1 &= \frac{(\alpha_1 - \rho^2(w - D(p_1, q, z))}{\sigma_\alpha\sqrt{1-\rho^2}} \\
r_2 &= \frac{(\alpha_2 - \rho^2(w - D(p_2, q, z))}{\sigma_\alpha\sqrt{1-\rho^2}} \\
r_3 &= \frac{(\alpha_3 - \rho^2(w - D(p_2, q, z))}{\sigma_\alpha\sqrt{1-\rho^2}} \\
r_4 &= \frac{(\alpha_4 - \rho^2(w - D(p_3, q, z))}{\sigma_\alpha\sqrt{1-\rho^2}.}
\end{aligned}
$$

8.1 The simulation procedure

The simulation procedure is aimed at assessing the industry response to a price change under block-prices and the effect of a price regime change on welfare. To evaluate the effect of a change in the price schedule on aggregate water demand one cannot trust the estimates of 4, as we already saw that the latter is only an upper bound on the aggregate effect. In principle, the aggregate effect has to be calculated as follows: evaluate the expected water use of each farmer twice – before and after the change in the price schedule, where the expectation is with respect to the density of individual water use. The aggregate effect is the sum of the individual effects.

The key element in our simulation is the density of individual water use which is given in equation (11). Note that this density depends on the variables in z as well as the individual farm thresholds w_1 and w_2 and consequently depends on their observed values. We calculate the expected individual water use, conditional on the price schedule and the other farm characteristics, as follows

$$\int_0^\infty wL(w, \hat{\theta}|z_i, w_{1i}, w_{2i}) \tag{12}$$

where i is a farm index and $\hat{\theta}$ is the vector of maximum likelihood estimates of the parameters in θ. The integral in (12) is calculated numerically over the domain of observed individual water use where the integral increments were chosen so that the estimated densities of all observation sum to one (up to 10^{-5}). The welfare level can be expressed as a function of water use and thus, evaluations of welfare level and welfare changes are not conceptually different. The calculation will be based on the expected individual welfare, as a function of individual water use, with respect to the same density as in (12).

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2.05 Zvi Lerman and Natalya Shagaida – Land Reform and Development of Agricultural Land Markets in Russia.

3.05 Ziv Bar-Shira, Israel Finkelshtain and Avi Simhon – Regulating Irrigation via Block-Rate Pricing: An Econometric Analysis.