

BLOCK-RATE VERSUS UNIFORM WATER PRICING IN AGRICULTURE: AN EMPIRICAL ANALYSIS

ZIV BAR-SHIRA, ISRAEL FINKELSHTAIN, AND AVI SIMHON

In this article 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 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.

Key words: block-rate pricing, irrigation.

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 et al. 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; Arbues, Garcia-Valinas, and Martinex-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 et al. 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.

Block-rate pricing advocates argue that faced with 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. 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

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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 ordinary least square (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 et al. (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, 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 interfarm 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 com-

munities 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 interfarm 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 -1.6 for the residential sector. Other studies of residential demand under tier pricing (e.g., Arbues et al. 2003) which did not apply Burtless and Hausman's methodology found significantly lower elasticities. Moreover, Moeltner and Stoddard (2004) applied a 2SLS procedure to estimated water demand under tier pricing in various commercial, nonagricultural sectors and found elasticities in the range of -0.23 to -0.9 .

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 subsequent section, followed by a detailed description of the data. The next section presents the estimation results, which are employed in following section for a simulation of alternative pricing policies and a welfare analysis. Final section concludes.

Irrigation-Water Demand under Block Prices

Consider a farmer who irrigates n crops and faces an increasing block-rate schedule $p(w)$, where $w = \sum_{i=1}^n w_i$ is the farm's aggregate water use and w_i is the quantity of water applied to irrigate crop i . The block-rate schedule $p(w)$ is given by the following (continuous from the left) step function:

$$p(w) = \begin{cases} p^1 & \text{if } w \leq w^1 \\ p^2 & \text{if } w^1 < w \leq w^2 \\ p^3 & \text{if } w^2 < w \end{cases}$$

where $p^1 < p^2 < p^3$ are price tiers and $w^1 < w^2$ are the water-quantity thresholds (see figure 1).

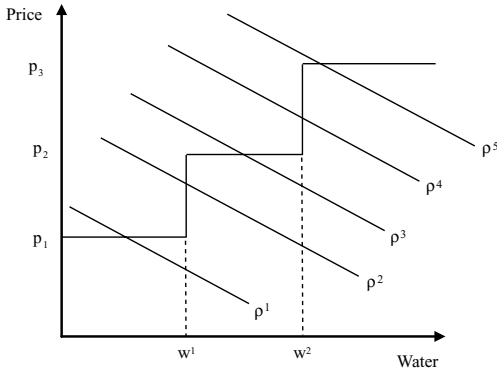


Figure 1. Increasing block price schedule

The farmer maximizes profits, given by: $\sum_{i=1}^n \pi^i f^i(w_i, x^i) - \int_0^w p(s) ds - \sum_{i=1}^n q^i x^i$, where π^i is the price of crop i , x^i is a vector of other inputs used in the production process of crop i , and q is the corresponding inputs price vector. The production technology of crop i is $f^i(\cdot)$, and is assumed monotonically increasing and strictly concave in w_i and in x^i . The value of the marginal product of water in the production of crop i is denoted by $\pi^i f_w^i(\cdot)$. The horizontal sum of these functions, denoted $\rho(w, x^1, \dots, x^n, \pi^1, \dots, \pi^n)$, is the value of the marginal product (VMP) of water on the farm. By concavity of f^i for all i , $\rho(\cdot)$ is decreasing in w .

The demand for x^i is determined by the usual equalities of the value of marginal product and its price. The derivation of the water demand is more complex; to simplify, we rely on graphical arguments.¹ The downward sloping lines in figure 1 represent alternative VMP curves. The increasing stepwise graph that we call the block-price graph is the union of $p(w)$ and two vertical segments that connect the price tiers (which, formally, are not part of $p(w)$). Since the VMP curves are decreasing everywhere, and the block-price graph is nondecreasing, they intersect only once. The intersection may take place either at a horizontal segment of the block prices or at a vertical segment. In both cases, the intersection determines the farmer's optimal aggregate water use. To see this, note that the farmer's profit equals the area below the VMP function minus the area below the block-price graph. Obviously, this area is maximized when the two intersect.

To see the quantities demanded, let $D(s, q) = \rho^{-1}(s)$ denote the inverse VMP (at the optimal levels of x^i for all i). Consider first

the curve ρ^1 in figure 1. The fact that it intersects the block-price graph at $w < w^1$ amounts to saying that $D(\rho^1, q) < w^1$. The optimality of the water use at the intersection implies that the farmer's water demand equals $D(\rho^1, q)$. Suppose now that the actual VMP curve is ρ^2 , intersecting the vertical segment connecting p^1 and p^2 . Then it follows that $D(\rho^1, q) \geq w^1$ and $D(\rho^2, q) \leq w^1$. As before, the optimality of the water use at the intersection implies that demand equals w^1 , and so on. This is summarized by the demand equation (1).

(1)

$$w(\cdot) = \begin{cases} D(\rho^1, q) & \text{if } D(\rho^1, q) < w^1 \\ w^1 & \text{if } D(\rho^2, q) \leq w^1 \\ & \leq D(\rho^1, q) \\ D(\rho^2, q) & \text{if } w^1 < D(\rho^2, q) < w^2 \\ w^2 & \text{if } D(\rho^3, q) \leq w^2 \leq D(\rho^2, q) \\ D(\rho^3, q) & \text{if } D(\rho^3, q) > w^2 \end{cases}$$

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

(2)

$$\frac{\partial w}{\partial p} = \begin{cases} D_p(\rho^1, q) & \text{if } D(\rho^1, q) \leq w^1 \\ 0 & \text{if } D(\rho^2, q) < w^1 \\ & < D(\rho^1, q) \\ D_p(\rho^2, q) & \text{if } w^1 < D(\rho^2, q) \leq w^2 \\ 0 & \text{if } D(\rho^3, q) < w^2 \\ & < D(\rho^2, q) \\ D_p(\rho^3, q) & \text{if } D(\rho^3, q) > w^2 \end{cases}$$

where w_p and D_p are derivatives with respect to p .²

It is interesting that while the VMP function is downward sloping everywhere, some price changes may not induce all farmers to reduce irrigation. A small price increase of all tiers does not affect the quantity demanded by the farmers who are represented by ρ^2 and ρ^4 . Moreover, the effect of a large price change on the farmers represented by ρ^1 , ρ^3 , and ρ^5 may be mitigated if their VMP curves intersect 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 is bounded above by the inverse elasticity of

¹ A calculus-based derivation is available from the authors upon request.

² Note that for a price decrease the derivative w_p is not the same.

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.

Econometric Methodology

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

The Statistical Model

Two sources of randomness are introduced into the economic model presented in Section Irrigation-water demand under block price. Farmers’ heterogeneity, α_{it} , and a measurement error, ε_{it} where $i = 1, 2, \dots, N$ denotes farm i and $t = 1, 2, \dots, T$ indicating a year t . Farmers’ heterogeneity, α_{it} , is a random term that captures variations across farms and over time that are neither explained by the observed characteristics, z_{it} , nor by farm fixed effects (dummies), d_i . Random heterogeneity may include variations in management capabilities and other variables that are unobserved by the econometrician, but are known to the farmer and taken into account in his optimization process. The error term, ε_{it} 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. Denoting by w_{it} the total quantity of water applied in i th farm in year t , and using the linearity assumption, we can rewrite (1) as

(3)

$$w_{it} = \begin{cases} D(p_t^1, q_t, d_i, z_{it}) + \alpha_{it} + \varepsilon_{it} & \text{if } \alpha_{it} < w_{it}^1 - D(p_t^1, q_t, d_i, z_{it}) \\ w_{it}^1 + \varepsilon_{it} & \text{if } w_{it}^1 - D(p_t^1, q_t, d_i, z_{it}) \leq \alpha_{it} \leq w_{it}^1 - D(p_t^2, q_t, d_i, z_{it}) \\ D(p_t^2, q_t, d_i, z_{it}) + \alpha_{it} + \varepsilon_{it} & \text{if } w_{it}^1 - D(p_t^2, q_t, d_i, z_{it}) < \alpha_{it} < w_{it}^2 - D(p_t^2, q_t, d_i, z_{it}) \\ w_{it}^2 + \varepsilon_{it} & \text{if } w_{it}^2 - D(p_t^2, q_t, d_i, z_{it}) \leq \alpha_{it} \leq w_{it}^2 - D(p_t^3, q_t, d_i, z_{it}) \\ D(p_t^3, q_t, d_i, z_{it}) + \alpha_{it} + \varepsilon_{it} & \text{if } \alpha_{it} > w_{it}^2 - D(p_t^3, q_t, d_i, z_{it}) \end{cases}$$

Note that all the firms in our economy face the same price tiers, $p_t^j, j = \{1, 2, 3\}$ but may have different threshold levels w_{it}^j . 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.

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 $\Pr(w_{it} | p_{it}(\cdot), z_{it}, d_i, \theta)$ the probability of observing a certain level of irrigation, w_{it} by firm i in year t , given the price schedule $p_{it}(\cdot)$, the observed characteristics z_{it} , and parameter vector θ . The latter includes demand parameters as well the parameters of the distributions of the errors α_{it} and ε_{it} . The probability of observing w_{it} 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,

$$\begin{aligned} \Pr_{it}(w_{it} | p_{it}(\cdot), z_{it}, d_i, \theta) &= \Pr[\alpha_{it} + \varepsilon_{it} = w_{it} - D(p_t^1, q_t, d_i, z_{it}), \\ &\quad \alpha_{it} < w_{it}^1 - D(p_t^1, q_t, d_i, z_{it})] \\ &\quad + \Pr[\alpha_{it} + \varepsilon_{it} = w_{it} - D(p_t^2, q_t, d_i, z_{it}), \\ &\quad w_{it}^1 - D(p_t^2, q_t, d_i, z_{it}) \leq \alpha_{it} \leq w_{it}^2 \\ &\quad - D(p_t^2, q_t, d_i, z_{it})] \end{aligned}$$

$$\begin{aligned}
 &+ \Pr[\alpha_{it} + \varepsilon_{it} = w_{it} - D(p_t^3, q_t, d_i, z_{it}), \\
 &\alpha_{it} > w_{it}^2 - D(p_t^3, q_t, d_i, z_{it})] \\
 &+ \Pr[\varepsilon_{it} = w_{it} - w_{it}^1, w_{it}^1 - D(p_t^1, q_t, d_i, z_{it}) \\
 &\leq \alpha_{it} \leq w_{it}^1 - D(p_t^2, q_t, d_i, z_{it})] \\
 &+ \Pr[\varepsilon_{it} = w_{it} - w_{it}^2, w_{it}^2 - D(p_t^2, q_t, d_i, z_{it}) \\
 &\leq \alpha_{it} \leq w_{it}^2 - D(p_t^3, q_t, d_i, z_{it})].
 \end{aligned}$$

The likelihood function of firm *i* is then given by

$$L_i = \prod_{t=1}^T \Pr(w_{it} | p_{it}(\cdot), z_{it}, d_i, \theta).$$

The sample likelihood function is

$$(4) \quad L = \prod_{i=1}^N L_i.$$

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 (4), in terms of the standard normal density, is readily derivable. For a detailed likelihood function see, for example, Hewitt and Hanemann (1995).

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 geometric lag price distribution into the model.³ It follows that $D(p, q, d, z)$ can be expressed as

$$\begin{aligned}
 (5) \quad \frac{w_{it}^\lambda - 1}{\lambda} &= \beta_0 + \beta_1 \sum_{j=0}^\infty \gamma^j \frac{p_{t-j}^\lambda - 1}{\lambda} \\
 &+ \beta_2 \sum_{j=0}^\infty \gamma^j \frac{q_{t-j}^\lambda - 1}{\lambda} \\
 &+ \delta' \frac{z_{it}^\lambda - 1}{\lambda} + \mu' d_i.
 \end{aligned}$$

The Box–Cox transformation is a relatively general formulation, nesting several functional forms such as the logarithmic ($\lambda \rightarrow 0$), 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 captured by the parameter γ , the rate at which the effect of prices decay over time.⁴ Those 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 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, and a time trend (a variable indicating the calendar year) to capture technological changes—in particular water saving ones. Finally, we use farm dummies to capture a farm *fixed effect*.

Taking the lag of (5) and multiplying it by yields

$$\begin{aligned}
 \gamma \frac{w_{it-1}^\lambda - 1}{\lambda} &= \gamma \beta_0 + \gamma \beta_1 \sum_{j=0}^\infty \gamma^j \frac{p_{t-j-1}^\lambda - 1}{\lambda} \\
 &+ \gamma \beta_2 \sum_{j=0}^\infty \gamma^j \frac{q_{t-j-1}^\lambda - 1}{\lambda} \\
 &+ \gamma \delta' \frac{z_{it-1}^\lambda - 1}{\lambda} + \gamma \mu' d_i.
 \end{aligned}$$

Subtracting the last equation from the previous one and rearranging yields

³ For the estimation procedure, we replace the other inputs and their prices by a quantity index of all other inputs and its associated price index. These indexes are compiled by The Israeli Central Bureau of Statistics comprising all other inputs excluding water.

⁴ For the role of geometric and other lag distributions in econometric models see, for example, chapter 19.3 in Greene (2003).

$$\begin{aligned}
 (6) \quad \frac{w_{it}^\lambda - 1}{\lambda} &= (1 - \gamma)\beta_0 + \gamma \frac{w_{it-1}^\lambda - 1}{\lambda} \\
 &+ \beta_1 \frac{p_t^\lambda - 1}{\lambda} + \beta_2 \frac{q_t^\lambda - 1}{\lambda} \\
 &+ \delta' \frac{z_{it}^\lambda - 1}{\lambda} - \gamma \delta' \frac{z_{it-1}^\lambda - 1}{\lambda} \\
 &+ (1 - \gamma)\mu' d_i
 \end{aligned}$$

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, $\beta_1/(1 - \gamma)$ measures its long-term effect. Second, the coefficient of the term before the last one in (6) 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 (5).

Institutional Description and Data

Our estimation is based on a micro level panel, which describes irrigation and geographical data for 185 cooperatives (*Kibutzim*) and semi-cooperatives (*Moshavim*), during the period 1992–97. These cooperatives comprise 20% of the water use for irrigation in Israel. The data for each year are documented in the Agricultural Ministry’s annual publication: “Agricultural Sub Industries” for the corresponding year.

Three criteria were employed in the selection of the 185 cooperatives. First, since our interest lies in estimating the marginal product of water, we limit our sample to farmers who consume only fresh water. Second, to minimize aggregation problems, we confine the analysis to observations that describe the consumption of a legally single consumer, receiving an annual water allocation and responsible for payment. Finally, to gain access to billing and pricing data, we confine our analysis to farmers who purchase the water from Mekorot—a state-owned company that supplies 60% of total irrigated fresh water.

Water Quotas

The Israeli water law legislated in 1959 determines that water in both underground aquifers and surface reservoirs is state property. Each year, each planned village is allocated an annual water quota for irrigation. Initial, historical quotas were allotted according to the village’s total arable land, crop type, location,

water usage in years prior to 1959, and political and organizational affiliations. Occasionally, water quotas were adjusted to reflect changes in the amount of land and water available for farming. 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 fresh-water quotas.

Important questions for our analysis were: to what extent are the quotas enforced and are they binding? If they are, water application would reflect quotas rather than the marginal product of water. 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 in which any sanctions were imposed. Thus, we are satisfied that the observed water application does reflect the marginal productivity of water.

Table 1 shows considerable variation in the levels of the quotas across farmers and over the years. The differences across users reflect the afore-listed allocation criteria. Aggregate variations over the years are primarily a reflection of fluctuations in annual precipitation.

Water Consumption

The average farm utilizes 650,000 m³ of water, an average of 677 m³ (approximately 0.5 acre-feet) of water per acre. Interestingly, the value of the crops per cubic meter has increased fivefold since 1950, reflecting considerable technological improvement (Kislev 2001). Our empirical formulation of the demand function in the next section takes those changes into account by adding a productivity trend variable. Table 2 shows considerable variations in water application across farmers, while changes in aggregate consumption over the years are moderate. The explana-

Table 1. Quotas (in thousands of cubic meters)

| Year | Mean | Max | Min | Std. Dev. |
|-----------|------|-------|-----|-----------|
| 1992 | 773 | 1,550 | 213 | 264 |
| 1993 | 927 | 1,650 | 239 | 297 |
| 1994 | 844 | 1,539 | 229 | 273 |
| 1995 | 848 | 1,650 | 239 | 269 |
| 1996 | 833 | 1,716 | 224 | 283 |
| 1997 | 978 | 1,854 | 239 | 333 |
| All years | 867 | 1,854 | 213 | 295 |

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

| Year | Mean | Max | Min | Std. Dev. |
|-----------|------|-------|-----|-----------|
| 1992 | 609 | 1,600 | 83 | 281 |
| 1993 | 647 | 1,346 | 83 | 296 |
| 1994 | 626 | 1,500 | 83 | 298 |
| 1995 | 654 | 1,600 | 100 | 315 |
| 1996 | 637 | 1,600 | 38 | 338 |
| 1997 | 649 | 1,441 | 38 | 349 |
| All years | 637 | 1,441 | 38 | 313 |

tions 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.

Prices

By the early 1990s, Israel was going through a rapid process of depleting its water reservoirs. In a 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 2 presents the three block rates (New Israeli Shekels per m^3 at 1997 prices) 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 changed over that period.

Table 3 presents the distribution of irrigated water over the three tiers in the years 1992 to 1997. 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%. Note that the share of farmers in the first block is nearly twice their consumption share. This latter observation and additional calculations, which are based on the table, suggest that the water quotas are systematically more generous for the smaller farmers, than for the large ones. This is in concert with the notion that the block-rate regime in Israel supports small farms. Our econometric findings are reported in the following sections.

Results and Simulations

The parameter estimates of our model are reported in the following subsection, where we

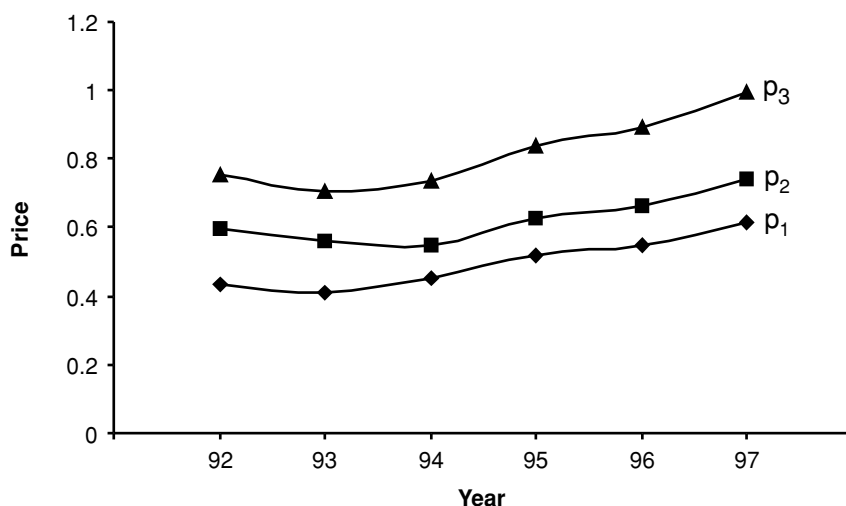
**Figure 2. The block rates**

Table 3. Proportions of All Farms with Operating Margins in Different Price Tiers and Proportion of Total Water Consumed by These Farms

| Year | 1st Tier | | 2nd Tier | | 3rd Tier | | Total (1,000 m ³) |
|---------|----------|---------|----------|---------|----------|---------|-------------------------------|
| | Water % | Farms % | Water % | Farms % | Water % | Farms % | |
| 1992 | 12 | 23 | 58 | 55 | 30 | 22 | 112,704 |
| 1993 | 11 | 21 | 44 | 46 | 45 | 33 | 119,743 |
| 1994 | 14 | 24 | 44 | 46 | 42 | 30 | 115,660 |
| 1995 | 13 | 24 | 39 | 42 | 48 | 34 | 120,917 |
| 1996 | 16 | 31 | 41 | 40 | 43 | 29 | 117,800 |
| 1997 | 15 | 31 | 40 | 39 | 45 | 30 | 120,091 |
| Average | 14 | 26 | 44 | 44 | 42 | 30 | 117,853 |

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.

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 0.65, an intermediate case between the linear and logarithmic models. We tested both hypotheses, that $\lambda \rightarrow 0$ 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 by 0.22 m³ for each additional cubic meter of quota in the linear case and with coefficients of 0.26 and 0.35 in the Box-Cox and the logarithmic specifications respectively. That is the case although the quotas were not surpassed by almost 90% of the farmers and penalties were never imposed when they were exceeded. 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 "Trend" effect is positive indicating, *ceteris paribus*, augmentation of water use over time. The "Trend" 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, improvements made over time in water-saving technology may under certain circumstances

Table 4. Maximum Likelihood Estimates

| Specification | Box-Cox | Logarithmic ($\lambda \rightarrow 0$) | 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) |
| Trend | 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 | –6,638 | –6,744 | –6,678 |
| Number of obs. | 1,110 | 1,110 | 1,110 |
| Number of coefs. | 194 | 193 | 193 |

Note: Numbers in parenthesis are asymptotic *t* statistics.

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.⁵ Figure 3 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 (1,000 m³) 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 $\frac{Z_{i-1}^\lambda - 1}{\lambda}$ in equation (6) equals $\gamma\delta$ and once without that constraint. Comparing the maximum likelihood of the two models, we could not reject, at a 5% significance level, the restricted model in favor of the unrestricted model, supporting the validity of the lag structure.

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 vertical section of the price schedule. To evaluate the aggregate elasticity we simulated the irrigated agricultural sector to predict water use; once under the actual price schedule and once when it is increased by 1%. It is evident that the greatest gap between the simulated and estimated elasticities is with the logarithmic specification and the smallest is with the linear. Intuitively, the less elastic the demand, 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%

⁵ The predicted individual water use, conditional on the price schedule and the other farm characteristics, were calculated by numerical calculation of $\int_0^\infty wL(w, \hat{\theta} | z_i, w_i^1, w_i^2) dw$, where i is a farm index and $\hat{\theta}$ is the vector of maximum likelihood estimates of the parameters in θ .

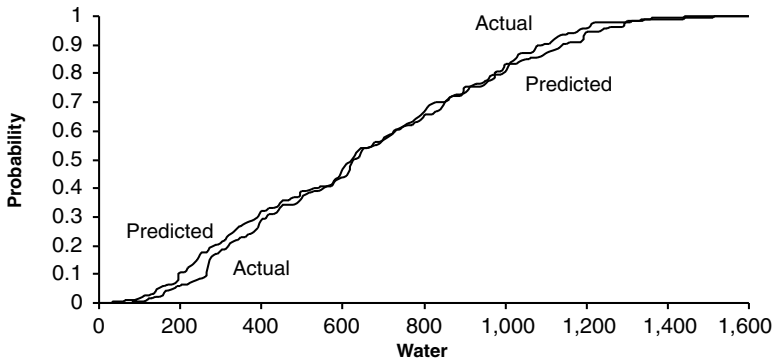


Figure 3. Predicted versus actual water distributions

Table 5. Moments of Actual and Predicted Water Use

| | Observed Use | Predicted Use |
|-----------|--------------|---------------|
| Mean | 649 | 654 |
| Median | 624 | 625 |
| Maximum | 1,441 | 1,518 |
| Minimum | 38 | 87 |
| Std. Dev. | 348 | 322 |
| Skewness | 0.228 | 0.261 |
| Kurtosis | 2.007 | 2.178 |

specifications. Finally, it is interesting to note that the effects of water quotas 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 regime on water demand and social welfare.

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 effect than expected, 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 two such years, the irrigation levels between them will differ by 6%–7%, according to the first two

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

Table 7 presents the differential effect of water price changes on farmers' profitability for different water demand groups. We divide the farmers in our sample into four equal size groups by their water use: the first quartile contains those with the lowest water demand, and the fourth quartile contains those who use the most. The average is over all the farmers. Then, we normalize to one the predicted profits of each group under the actual 1997 block-rate

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 |
| Simulated 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 |

Table 7. Proportion of Profits Lost by Farmers from Switching from Block Rate Pricing to a Single Price, by Quartile of Farm Profits

| Quartile | Actual Water Use | | Water Use + 20% | | Water Use - 20% | |
|----------|------------------|------|-----------------|------|-----------------|------|
| | Block | Flat | Block | Flat | Block | Flat |
| I | 1 | 0.80 | 1.44 | 1.27 | 0.51 | 0.40 |
| II | 1 | 0.89 | 1.25 | 1.17 | 0.69 | 0.60 |
| III | 1 | 0.93 | 1.18 | 1.13 | 0.76 | 0.71 |
| IV | 1 | 0.96 | 1.14 | 1.11 | 0.82 | 0.79 |
| Average | 1 | 0.94 | 1.18 | 1.13 | 0.78 | 0.73 |

prices and calculate the effect of several price and regime changes on each of them.

The second column shows the effect of switching from block prices to a single-price regime, choosing the price such that the same aggregate use is maintained. Clearly, the farmers most affected are the smallest who suffer a 20% loss in profits, while the least affected are the largest water users, who lose only 4%. Interestingly, switching to a single price lowered the profitability of all groups (while raising the revenues of the water supplier).

The third and fourth columns present the effect of reducing water prices, using block prices and a single price, respectively, such that there is a 20% increase in aggregate water use. The last two columns present the effect of price

increases that result in a 20% decrease in aggregate use. Again, the choice of regime has a very large effect on the small farmers, but not so large an effect on the big farmers.

Despite the apparent advantages of regulating water demand through block prices, it should be emphasized that the second best is nonimplementable as it requires that every farmer's demand function intersect the price schedule at the highest tier. Such a scheme is practically impossible due to farmers' heterogeneity and the obvious difficulty for policy makers to observe each farmer's demand function.

Figure 4 demonstrates the differential effect of switching from a single-price p^0 to the block prices p^1, p^2 on farmers with different

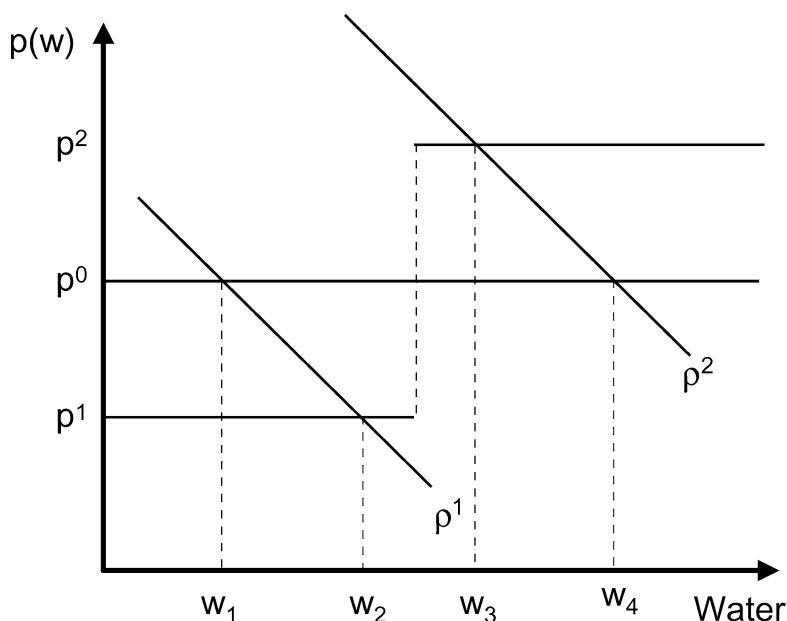


Figure 4. Welfare effect of block-rate pricing

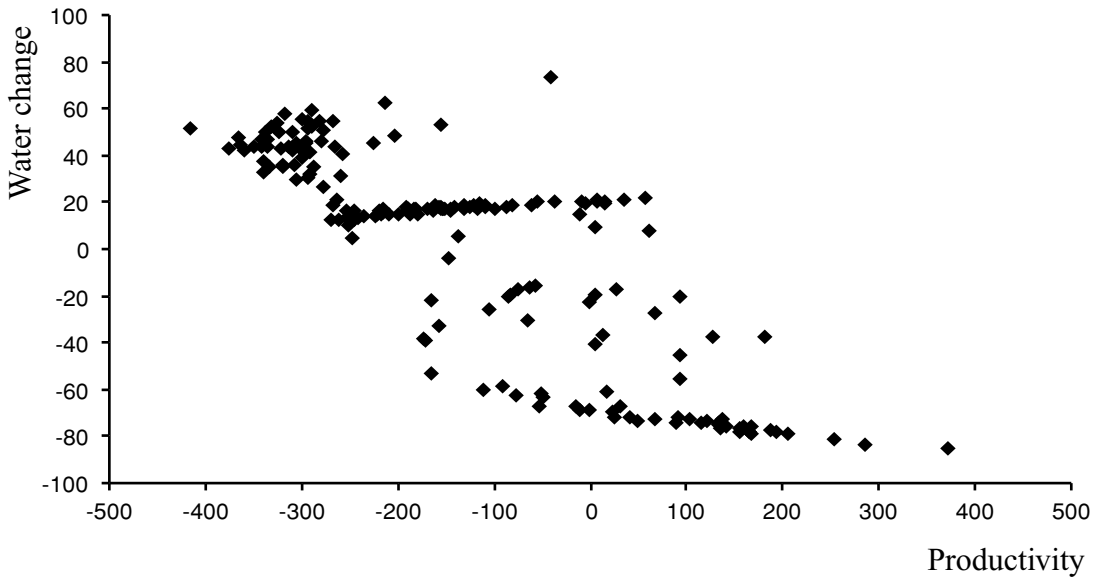


Figure 5. Relationship between the change in water use when switching to block pricing and productivity

VMP 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 farmers reduce it, what is the net effect on aggregate water use? Analytically, if the weighted 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 using our data.

We first simulate the agricultural sector to find the predicted water demand under a single price that equals the 1997 weighted average price. It turns out that the aggregate demand under the single-price regime is 7% higher than under block pricing. Thus, in Israel, where the switch from a single-price regime to block prices was designed to reduce aggregate water use without raising its cost to farmers, it achieved its goal.

Alternatively, suppose that a 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. However, as we saw in figure 4, a switch to block prices has different effects on farmers with different productivity levels and therefore

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

In practice, we calculate the areas under the estimated demand function (6) in 1997 for each farmer and sum them over all farmers. We find that 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.⁶ Since the loss to the water suppliers was larger than the farmers' aggregate gain, the switch to block prices lowered social welfare. The welfare loss is a general outcome that stems from the fact that switching to block prices while maintaining the average price raises the water use of the less-productive farmers and depresses that of the most productive farmers. This is demonstrated in figure 5, where for each farmer we chart the change in predicted water use when switching from a single price to block pricing, against his productivity level.⁷

⁶ It is not always the case that all the farmers gain from a switch to block prices; rather it is the outcome of the specific price schedule that was enacted in Israel at the time of our study.

⁷ As farmers' productivity we use their estimated value of d_i .

Table 8. Reducing Water Use While Preserving Small Farms

| | Block Pricing | Flat Pricing |
|-------------------------------------|---------------|--------------|
| Water change | -13.5% | -6.3% |
| Profit change | -14.7% | -12.7% |
| Welfare enhancement (million \$) | 8.8 | 5.4 |

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 here 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, they would benefit from the switch to block pricing while the larger farmers lose from it. 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.

As we mentioned before, an important reason for switching to block prices is to divert water from agricultural to urban use at a lower cost to small farmers than a rise in a single price would entail. To study this issue, we examine a reduction in water use in agriculture, under the constraint that the profits of the lowest farmers quartile are unchanged. Table 8 presents a comparison between two possible policies: (i) raising the actual 1997 block prices by 50%, and (ii) setting a single price that would result in the same average profit for the bottom quartile of farmers as in alternative (i).

Since the two regimes lead to different water use in the agricultural sector, the comparison requires an evaluation of the water value in other sectors of the economy. To this end, we use the 2000 financial statement of the national water supplier, Mekorot. According to that statement, the cost of pumping and delivering 1 m³ of water including an aquifer levy is \$ 0.36. The aquifer toll is levied to rationalize water pumping and to reflect the economic shadow price of water. Therefore, this number is a close proxy for the marginal value of fresh water in Israel and is employed in the simulation below.

The first row of table 8 shows that water use declines by 13.5% per farmer (from the actual use in 1997) under block prices, and by 6.3% under the single price. That is, block prices achieve more than double the reduction of water use in agriculture at the same cost for the bottom quartile. The second row presents the average decline of profits for all the farmers—14.7% under block prices and by 12.7% under the single price. The last row shows that the social welfare increases by \$8.8 million under block prices and by only \$5.4 million under the single price.⁸

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 interfarm 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 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 will increase irrigation while large farmers will reduce it. If

⁸ The change in welfare was calculated under the assumption that our sampled farms represent 25% of Israeli irrigated agriculture. Thus, the quantity of saved water in the sample was amplified fourfold.

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 expense of the more productive. The alternative of course is to benefit all farmers at the expense of the water suppliers.

While potential interfarm inefficiencies are a serious concern, our estimates indicate that with our data set, 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.

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